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AUTHOR Mislevy, Robert J.
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A closed form approximation is given for the variance of examinee proficiency estimates in the Rasch model for dichotomous items, under the condition that only estimates, rather than true values, of item difficulty parameters are available. The term that must be added to the usual response-sampling variance is inversely proportional to both the number of examinees in the item calibration sample and the length of the test. Illustrative numerical values suggest that the impact of uncertainty about Rasch item parameters on subsequent estimates of examinee proficiencies is less than has been observed in the two- and three-parameter logistic item response theory models. Two tables present approximate error variances, and an eight-item list of references is included. (Author/SLD)

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THE VARIANCE OF RASCH ABILITY ESTIMATES FROM PARTIALLY-KNOWN ITEM PARAMETERS

Robert J. Mislevy
Educational Testing Service

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Key words: Cohen's approximation, Rasch model, variance components.

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The Variance of Rasch Ability Estimates from Partially-Known Item Parameters

Robert J. Mislevy

Educational Testing Service

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The Variance of Rasch Ability Estimates from Partially-Known Item Parameters

Abstract

A closed-form approximation is given for the variance of examinee proficiency estimates in the Rasch model for dichotomous items, under the condition that only estimates, rather than true values, of item difficulty parameters are available. The term that must be added to the usual response-sampling variance is inversely proportional to both the number of examinees in the item calibration sample and the length of the test. Illustrative numerical values suggest that the impact of uncertainty about Rasch item parameters on subsequent estimates of examinee proficiencies is less than has been observed in the 2- and 3-parameter logistic IRT models.

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Introduction

It is common practice in applications of item response theory (IRT) to estimate item parameters from a calibration sample of examinees, then use these estimates as if they were known true parameters when estimating the proficiency parameters of these and subsequent examinees. Tsutakawa and his colleagues have pointed out that ignoring the uncertainty about item parameters leads to inferential errors, and offered a Taylor series approximation to incorporate this uncertainty into proficiency estimates and statements about their precision. Results for the two-parameter logistic IRT model appear in Tsutakawa and Soltys (1988) and for the three-parameter logistic model appear in Tsutakawa and Johnson (1990). (Also see Lewis, 1985, for an alternative approach to the problem, and Mislevy, Sheehan, & Wingersky, in press, for the application of these ideas to test equating.) Wright and Panchapekesan (1969) present without derivation a closed-form approximation for the variance of examinee proficiency estimates under the Rasch model for dichotomous items (Rasch, 1960/1980). Utilizing a series of computing approximations and simplifying assumptions, the present paper gives a somewhat simpler approximation. *Neither numerical accuracy nor practical application is the point of this presentation—the aforementioned alternatives are preferable in this regard. Rather, by giving an approximation with fewer "moving parts," the objective is to aid intuition into the sources, the structures, and the consequences of uncertainty.*

The Rasch Model

The Rasch model for dichotomous items gives the probability of a correct response to Item j from an examinee with proficiency θ as

$$P_j(\theta) = \frac{\exp(\theta - b_j)}{1+\exp(\theta - b_j)}, \qquad (1)$$

where b_j is the difficulty parameter of Item j. Letting x_j denote the response to Item j, 1 for correct and 0 for incorrect, the assumption of local independence gives the conditional probability of the response vector $\mathbf{x}=(x_1,\ldots,x_L)$ on an L-item test as

$$P(\mathbf{x}|\theta,\mathbf{b}) = \prod_{j=1}^{L} P_j(\theta)^{x_j}\left[1-P_j(\theta)\right]^{1-x_j}, \tag{2}$$

where $\mathbf{b}=(b_1,\ldots,b_L)$.

Given item difficulty parameters and a realization of $\mathbf{x}$, (2) can be interpreted as a likelihood function for θ. Unless all the responses are correct or all are incorrect, it attains its maximum at a unique point, the maximum likelihood estimate (MLE) $\hat{\theta}$. The negative reciprocal of its second derivative at this point approximates the sampling variance of $\hat{\theta}$, or $\mathrm{Var}(\hat{\theta}|\theta,\mathbf{b})$. Given response vectors from a calibration sample of N examinees, the product of expressions like (2) serves similarly as a basis for the estimation of $\mathbf{b}$ and the sampling variance of the resulting estimate $\hat{\mathbf{b}}$.

Cohen's Approximations

Cohen (1979) gives closed form approximations for $\hat{\theta}$, $\mathrm{Var}(\hat{\theta}|\theta,\mathbf{b})$, $\hat{\mathbf{b}}$, and $\mathrm{Var}(\hat{\mathbf{b}}|\mathbf{b})$ for long tests when both θ and b are normally distributed, in the samples of examinees and items respectively. Under these circumstances, item difficulty parameters can be approximated by rescaling the logits, or log-odds, of item percents-correct in the calibration sample, and examinees' proficiencies can be approximated by rescaling logits of their total scores by functions of item parameters. Adapting Wright's (1977) presentation, we begin with a calibration data set from which zero and perfect scoring examinees and items with all or no correct responses have been excised, then calculate the following quantities:

$\overline{P}_j$ — the proportion of the calibration sample that answered Item j correctly, for j=1,...,L,

n_r — the number of examinees with a total score of r, for r=1,...,L-1,

$t_j = \ln\left(\frac{1-\overline{P}_j}{\overline{P}_j}\right)$ — the log ratio of wrong to right answers on Item j,

$t_. = \sum_{j=1}^{L} t_j/L$ — the mean of t_j over L items,

$T = \sum_{j=1}^{L} (t_j-t_.)^2/(L-1)$ — the variance of t_j over L items,

$y_r = \ln\left(\frac{r}{L-r}\right)$ — the log ratio of right to wrong answers for score r,

$y_. = \sum_{r=1}^{L-1} n_r y_r/N$ — the mean of y_r over N examinees,

$Y = \sum_{r=1}^{L-1} n_r(y_r-y_.)^2/(N-1)$ — the variance of y_r over N examinees, and

$W = \left(\frac{1+Y/2.89}{1-TY/8.35}\right)^{1/2}$ — an expansion factor, in which 1.7 is used to match the logistic and normal distributions, and $2.89=1.7^2$ and $8.35=1.7^4$.

Then, from Cohen (1979), item parameters can be approximated as follows:

$$\hat{b}_j \approx W(t_j-t_.) \qquad (3)$$

and

$$\mathrm{Var}(\hat{b}_j|b_j) \approx \frac{1}{N}\left(\frac{W^2}{\bar{P}_j(1-\bar{P}_j)}\right) \equiv \frac{C_j}{N},$$

where $C_j=W^2/[\bar{P}_j(1-\bar{P}_j)]$. The sampling covariances among the parameters of different items tend toward zero as L increases, as long as the items are not too dissimilar (Mislevy, 1981, shows that the error correlations among equivalent items are $-1/(L-1)$). We shall therefore approximate the L-by-L covariance matrix of sampling errors among item parameters as simply

$$\mathrm{Var}(\hat{\mathbf{b}}|\mathbf{b}) \approx \mathrm{diag}(\frac{C_1}{N},\ldots,\frac{C_L}{N}). \tag{4}$$

If item parameters are known, examinee proficiencies can be similarly approximated by rescaling logits of their total scores. Define three additional quantities:

$b_. = \sum_{j=1}^{L} b_j/L$ — the mean of b_j over L items,

$U = \sum_{j=1}^{L} (b_j-b_.)^2/(L-1)$ — the variance of b_j over L items, and

$X = (1+U/2.89)^{1/2}$ — an expansion factor.

Then, for a total score of r,

$$\hat{\theta} \approx b_. + X \ln\left(\frac{r}{L-r}\right) = b_. + X \ln\left(\frac{p_r}{1-p_r}\right), \tag{5}$$

where $p_r \equiv r/L$, and

$$\mathrm{Var}(\hat{\theta}|\theta,\mathbf{b}) \approx X^2 \left(\frac{L}{r(L-r)}\right) \equiv \frac{1}{L}\left(\frac{X^2}{p_r(1-p_r)}\right). \tag{6}$$

Estimating Ability with Item Parameter Estimates

When **b** is known, (5) gives a proficiency estimate from **x**, and (6), evaluated with the true value of θ, gives its sampling variance. What is the sampling variance of $\hat{\theta}$ when (5) is evaluated with $\hat{\mathbf{b}}$ rather than **b**? In this presentation, this amounts to using item parameter estimates rather than true parameter values to calculate $b_.$ and X. The setup corresponds to a two-stage experiment:

Stage 1: A sample of N examinees is drawn from a normal population. They respond to the L items in the test. Their responses are used to calculate $\hat{\mathbf{b}}$ via (3).

Stage 2: A new examinee is administered the test and the test score r is observed. A proficiency estimate is calculated using (5) with $\hat{\mathbf{b}}$. We can write this estimate as $\hat{\theta}(r,\hat{\mathbf{b}})$ to emphasize that it is a function of two statistics—which, it will be noted, are independent, given **b** and θ.

The variance of $\hat{\theta}$ in this setting can be decomposed into two terms:

$$\mathrm{Var}(\hat{\theta}|\theta) = E_{\hat{b}}\left[\mathrm{Var}(\hat{\theta}|\theta,\hat{\mathbf{b}})\right] + \mathrm{Var}_{\hat{b}}\left[E(\hat{\theta}|\theta,\hat{\mathbf{b}})\right]. \tag{7}$$

The first term on the right of (7) is the expectation of (6) as evaluated with $\hat{\mathbf{b}}$, over the distribution of $\hat{\mathbf{b}}$ given **b**. It depends on **b** only through X^2, or 1+U/2.89, where U is the variance of the true item parameters. The expected variance of the item parameter *estimates* is the variance of the item parameters themselves plus the average estimation error variance, $C_./N=(\Sigma C_j/L)/N$. Thus,

$$E_{\hat{b}}\left[\mathrm{Var}(\hat{\theta}|\theta,\hat{\mathbf{b}})\right] = E_{\hat{b}}\left[\frac{1}{L}\left(\frac{\widehat{X^2}}{p_r(1-p_r)}\right)\right]$$

$$= \frac{1}{L}\left(\frac{1+(U+C_./N)/2.89}{p_r(1-p_r)}\right)$$

$$= \frac{1}{L}\left(\frac{X^2}{p_r(1-p_r)}\right) + \frac{C_.}{LN}\left(\frac{1}{2.89\,p_r(1-p_r)}\right). \tag{8}$$

This is the sum of the error variance for $\hat{\theta}$ with known item parameters and a term depending on their error variance that is inversely proportional to both L and N.

The second term in (7) concerns the variation in $\hat{\theta}$ associated with the expected test score r for the true θ, as induced by variation in $\hat{\mathbf{b}}$. This term can be approximated by the delta method as follows:

$$\mathrm{Var}_{\hat{b}}\left[E(\hat{\theta}|\theta,\hat{\mathbf{b}})\right] \approx \frac{\partial\hat{\theta}}{\partial\hat{\mathbf{b}}'}\,\mathrm{Var}(\hat{\mathbf{b}}|\mathbf{b})\,\frac{\partial\hat{\theta}}{\partial\hat{\mathbf{b}}}$$

$$\approx \sum_{j=1}^{L}\left[\left(\frac{\partial\hat{\theta}}{\partial\hat{b}_j}\right)^2 \mathrm{Var}(\hat{b}_j|b_j)\right] \qquad \text{[by (4)]}$$

$$= \sum_{j=1}^{L}\left[\left\{\frac{\partial}{\partial\hat{b}_j}\left[\hat{b}_. + \hat{X}\ln\left(\frac{p_r}{1-p_r}\right)\right]\right\}^2 \mathrm{Var}(\hat{b}_j|b_j)\right]$$

$$\approx \sum_{j=1}^{L}\left[\frac{1}{L} + \frac{(b_j - b_.)\ln\left(\frac{p_r}{1-p_r}\right)}{2.89\,LX}\right]^2 \frac{C_j}{N}.$$

Approximating all C_js by their average, say $C_.$, leads to the following simplification:

$$\mathrm{Var}_{\hat{b}}\left[E(\hat{\theta}|\theta,\hat{\mathbf{b}})\right] \approx \frac{C_.}{LN}\left[1 + \frac{U\left\{\ln\left(\frac{p_r}{1-p_r}\right)\right\}^2}{8.35\,X^2}\right]. \tag{9}$$

Substituting (8) and (9) back into (7) yields the final approximation:

$$\mathrm{Var}(\hat{\theta}|\theta) \approx \frac{1}{L}\left(\frac{X^2}{p_r(1-p_r)}\right) + \frac{C_.}{LN}\left(1 + \frac{1}{2.89\,p_r(1-p_r)} + \frac{U\left\{\ln\left(\frac{p_r}{1-p_r}\right)\right\}^2}{8.35\,X^2}\right)$$

$$\approx \mathrm{Var}(\hat{\theta}|\theta,\mathbf{b}) + \frac{C_.}{LN}\left(1 + \frac{1}{2.89\,p_r\,(1-p_r)} + \frac{U\left\{\ln\left(\frac{p_r}{1-p_r}\right)\right\}^2}{8.35\;X^2}\right). \tag{10}$$

Thus the sampling variance of $\hat{\theta}$ with item parameter estimates can be approximated as the sum of its variance with true item parameters and a correction term. The following observations can be made about the correction term:

1. It is always positive.
2. It is inversely proportional to test length (L).
3. It is inversely proportional to the size of the examinee calibration sample (N).
4. It is directly proportional to $C_.$. Since

   $$C_j = \frac{\left(\frac{1+Y/2.89}{1-TY/8.35}\right)}{\bar{P}_j(1-\bar{P}_j)},$$

   where Y is the variance of the logits of item percents-correct and $\bar{P}_j$ is the percent-correct for Item j in the calibration sample, the impact of the uncertainty about item parameters is greater when...

   a. the items are more dispersed, so that Y departs from zero, or
   b. items are farther from 50-percent correct, so that $\bar{P}_j(1-\bar{P}_j)$ terms fall farther from the maximum value of .25.
5. Values of θ for which expected proportions of correct response are near 50-percent are less affected by uncertainty about item parameters than more extreme values of θ. This follows from the second and third terms inside the brackets in (10): As p_r approaches .5, the second term approaches its minimum value of 4/2.89, and, because the logit of .5 is zero, the third term approaches zero.

Illustrative Numerical Values

To provide a feel for the relative contributions of sources of uncertainty this section evaluates the preceding formulae with some illustrative values. Standard normal distributions are assumed for θ and b in the item calibration sample, implying the following values for true parameter values:

$$t. = y. = b. = 0$$

$$T = Y = .743$$

$$X = W = 1.160$$

$$U = 1 .$$

Equation (10) is used to approximate $\hat{\theta}$ sampling variances, for $\theta = 0$ and 2, for test lengths of 10, 20, 40, 80, and 160 items, after item calibration with examinee samples of 50, 100, 250, 500, 1000, 2500, and 10000. For C., we employ a representative value, namely that for an item one standard deviation from the mean. For such an item, $\bar{P}$ is about .7, in which case

$$C. = W^2/(.7 \times .3) = 6.40 .$$

The values of p_r that correspond to $\theta = 0$ and 2 are .50 and .85.

Table 1 gives results for $\theta = 0$, in terms of approximate sampling variances and proportional increases in sampling variance over those based on known item parameters. The proportional increases in standard errors are the square roots of the values in the rightmost column. Table 2 gives similar results for $\theta = 2$.

[Tables 1 and 2 about here]

The most striking feature of these results is how small the corrections are. Even with a calibration sample of only 50 examinees estimation error variances for subsequent θ estimates increase by only about 5-percent. This contrasts with the increases of up to 30-percent Tsutakawa and Soltys (1988) observed in posterior variances for θ with a calibration sample of 100 under the 2-parameter logistic model. Tsutakawa and Johnson's (1990) results for the 3-parameter logistic were even more extreme, with increases in posterior variances more than doubling at higher levels of proficiency with a calibration sample of 400.

Conclusion

Cohen's (1979) closed-form approximations for the parameters in the Rasch model support a closed-form approximation for the variance of these examinee proficiency estimates when they are calculated with item parameter *estimates* rather than true item parameter values. The approximation is the sum of the sampling variance with known item parameters and a correction term. The correction term is inversely proportional to both test length and the size of the examinee item-calibration sample. Illustrative numerical values suggests the additional variance is quite small, yielding increases in standard errors of less than 5-percent even with calibration samples of only 50 examinees.

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Table 1

Approximate Error Variances at θ=0

N	L: 10	20	40	80	160	$\frac{\widehat{Var}(\theta\mid\theta,b)}{\widehat{Var}(\theta\mid\theta,\hat{b})}$
50	0.5689	0.2845	0.1422	0.0711	0.0356	1.0567
100	0.5537	0.2768	0.1384	0.0692	0.0346	1.0283
250	0.5445	0.2723	0.1361	0.0681	0.0340	1.0113
500	0.5415	0.2707	0.1354	0.0677	0.0338	1.0057
1000	0.5399	0.2700	0.1350	0.0675	0.0337	1.0028
2500	0.5390	0.2695	0.1348	0.0674	0.0337	1.0011
10000	0.5386	0.2693	0.1346	0.0673	0.0337	1.0003
∞	0.5384	0.2692	0.1346	0.0673	0.0337	1.0000

Table 2

Approximate Error Variances at θ=2

	L					
N	10	20	40	80	160	$\frac{Var(\hat{\theta}\|\theta,\mathbf{b})}{Var(\hat{\theta}\|\theta,\hat{\mathbf{b}})}$
50	1.1066	0.5533	0.2767	0.1383	0.0692	1.0483
100	1.0812	0.5406	0.2703	0.1351	0.0676	1.0241
250	1.0659	0.5329	0.2665	0.1332	0.0666	1.0097
500	1.0608	0.5304	0.2652	0.1326	0.0663	1.0048
1000	1.0582	0.5291	0.2646	0.1323	0.0661	1.0024
2500	1.0567	0.5284	0.2642	0.1321	0.0660	1.0010
10000	1.0559	0.5280	0.2640	0.1320	0.0660	1.0002
∞	1.0557	0.5278	0.2639	0.1320	0.0660	1.0000

Dr. Terry Ackerman
Educational Psychology
260C Education Bldg.
University of Illinois
Champaign, IL 61801

Dr. Terry Allard
Code 1142CS
Office of Naval Research
800 N. Quincy St.
Arlington, VA 22217-5000

Dr. Nancy Allen
Educational Testing Service
Princeton, NJ 08541

Dr. Gregory Anrig
Educational Testing Service
Princeton, NJ 08541

Dr. Phipps Arabie
Graduate School of Management
Rutgers University
92 New Street
Newark, NJ 07102-1895

Dr. Isaac I. Bejar
Law School Admissions
Services
Box 40
Newtown, PA 18940-0040

Dr. William O. Berry
Director of Life and
Environmental Sciences
AFOSR/NL, N1, Bldg. 410
Bolling AFB, DC 20332-6448

Dr. Thomas G. Bever
Department of Psychology
University of Rochester
River Station
Rochester, NY 14627

Dr. Menucha Birenbaum
Educational Testing
Service
Princeton, NJ 08541

Dr. Bruce Bloxom
Defense Manpower Data Center
99 Pacific St.
Suite 155A
Monterey, CA 93943-3231

Dr. Gwyneth Boodoo
Educational Testing Service
Princeton, NJ 08541

Dr. Richard L. Branch
HQ, USMEPCOM/MEPCT
2500 Green Bay Road
North Chicago, IL 60064

Dr. Robert Brennan
American College Testing
Programs
P. O. Box 168
Iowa City, IA 52243

Dr. David V. Budescu
Department of Psychology
University of Haifa
Mount Carmel, Haifa 31999
ISRAEL

Dr. Gregory Candell
CTB/MacMillan/McGraw-Hill
2500 Garden Road
Monterey, CA 93940

Dr. Paul R. Chatelier
Perceptronics
1911 North Ft. Myer Dr.
Suite 1100
Arlington, VA 22209

Dr. Susan Chipman
Cognitive Science Program
Office of Naval Research
800 North Quincy St.
Arlington, VA 22217-5000

Dr. Raymond E. Christal
UES LAMP Science Advisor
AL/HRMIL
Brooks AFB, TX 78235

Dr. Norman Cliff
Department of Psychology
Univ. of So. California
Los Angeles, CA 90089-1061

Director
Life Sciences, Code 1142
Office of Naval Research
Arlington, VA 22217-5000

Commanding Officer
Naval Research Laboratory
Code 4827
Washington, DC 20375-5000

Dr. John M. Cornwell
Department of Psychology
I/O Psychology Program
Tulane University
New Orleans, LA 70118

Dr. William Crano
Department of Psychology
Texas A&M University
College Station, TX 77843

Dr. Linda Curran
Defense Manpower Data Center
Suite 400
1600 Wilson Blvd
Rosslyn, VA 22209

Dr. Timothy Davey
American College Testing Program
P.O. Box 168
Iowa City, IA 52243

Dr. Charles E. Davis
Educational Testing Service
Mail Stop 22-T
Princeton, NJ 08541

Dr. Ralph J. DeAyala
Measurement, Statistics,
and Evaluation
Benjamin Bldg., Rm. 1230F
University of Maryland
College Park, MD 20742

Dr. Sharon Derry
Florida State University
Department of Psychology
Tallahassee, FL 32306

Hei-Ki Dong
Bellcore
6 Corporate Pl.
RM: PYA-1K207
P.O. Box 1320
Piscataway, NJ 08855-1320

Dr. Neil Dorans
Educational Testing Service
Princeton, NJ 08541

Dr. Fritz Drasgow
University of Illinois
Department of Psychology
603 E. Daniel St.
Champaign, IL 61820

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Dr. Richard Duran
Graduate School of Education
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210 Fishburne Bldg.
Atlanta, GA 30322

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2440 Research Blvd., Suite 550
Rockville, MD 20850-3238

Dr. Marshall J. Farr
Farr-Sight Co.
2520 North Vernon Street
Arlington, VA 22207

Dr. Leonard Feldt
Lindquist Center
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Iowa City, IA 52242

Dr. Richard L. Ferguson
American College Testing
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912 South Wood Street
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Learning Research
& Development Center
University of Pittsburgh
3939 O'Hara Street
Pittsburgh, PA 15260

Dr. Susan R. Goldman
Peabody College, Box 45
Vanderbilt University
Nashville, TN 37203

Dr. Timothy Goldsmith
Department of Psychology
University of New Mexico
Albuquerque, NM 87131

Dr. Joseph McLachlan
Navy Personnel Research
and Development Center
Code 14
San Diego, CA 92152-6800

Alan Mead
c/o Dr. Michael Levine
Educational Psychology
210 Education Bldg.
University of Illinois
Champaign, IL 61801

Dr. Timothy Miller
ACT
P. O. Box 168
Iowa City, IA 52243

Dr. Robert Mislevy
Educational Testing Service
Princeton, NJ 08541

Dr. Ivo Molenaar
Faculteit Sociale Wetenschappen
Rijksuniversiteit Groningen
Grote Kruisstraat 2/1
9712 TS Groningen
The NETHERLANDS

Dr. E. Muraki
Educational Testing Service
Rosedale Road
Princeton, NJ 08541

Dr. Ratna Nandakumar
Educational Studies
Willard Hall, Room 213E
University of Delaware
Newark, DE 19716

Academic Progs. & Research Branch
Naval Technical Training Command
Code N-62
NAS Memphis (75)
Millington, TN 30854

Dr. W. Alan Nicewander
University of Oklahoma
Department of Psychology
Norman, OK 73071

Head, Personnel Systems Department
NPRDC (Code 12)
San Diego, CA 92152-6800

Director
Training Systems Department
NPRDC (Code 14)
San Diego, CA 92152-6800

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Special Assistant for Research
Management
Chief of Naval Personnel (PERS-01JT)
Department of the Navy
Washington, DC 20350-2000

Dr. Judith Orasanu
Mail Stop 239-1
NASA Ames Research Center
Moffett Field, CA 94035

Dr. Peter J. Pashley
Educational Testing Service
Rosedale Road
Princeton, NJ 08541

Wayne M. Patience
American Council on Education
GED Testing Service, Suite 20
One Dupont Circle, NW
Washington, DC 20036

Dept. of Administrative Sciences
Code 54
Naval Postgraduate School
Monterey, CA 93943-5026

Dr. Peter Pirolli
School of Education
University of California
Berkeley, CA 94720

Dr. Mark D. Reckase
ACT
P. O. Box 168
Iowa City, IA 52243

Mr. Steve Reise
Department of Psychology
University of California
Riverside, CA 92521

Mr. Louis Roussos
University of Illinois
Department of Statistics
101 Illini Hall
725 South Wright St.
Champaign, IL 61820

Dr. Donald Rubin
Statistics Department
Science Center, Room 608
1 Oxford Street
Harvard University
Cambridge, MA 02138

Dr. Fumiko Samejima
Department of Psychology
University of Tennessee
310B Austin Peay Bldg.
Knoxville, TN 37966-0900

Dr. Mary Schratz
4100 Parkside
Carlsbad, CA 92008

Mr. Robert Semmes
N218 Elliott Hall
Department of Psychology
University of Minnesota
Minneapolis, MN 55455-0344

Dr. Valerie L. Shalin
Department of Industrial
Engineering
State University of New York
342 Lawrence D. Bell Hall
Buffalo, NY 14260

Mr. Richard J. Shavelson
Graduate School of Education
University of California
Santa Barbara, CA 93106

Ms. Kathleen Sheehan
Educational Testing Service
Princeton, NJ 08541

Dr. Kazuo Shigemasu
7-9-24 Kugenuma-Kaigan
Fujisawa 251
JAPAN

Dr. Randall Shumaker
Naval Research Laboratory
Code 5500
4555 Overlook Avenue, S.W.
Washington, DC 20375-5000

Dr. Judy Spray
ACT
P.O. Box 168
Iowa City, IA 52243

Dr. Martha Stocking
Educational Testing Service
Princeton, NJ 08541

Dr. William Stout
University of Illinois
Department of Statistics
101 Illini Hall
725 South Wright St.
Champaign, IL 61820

Dr. Kikumi Tatsuoka
Educational Testing Service
Mail Stop 03-T
Princeton, NJ 08541

Dr. David Thissen
Psychometric Laboratory
CB# 3270, Davie Hall
University of North Carolina
Chapel Hill, NC 27599-3270

Mr. Thomas J. Thomas
Federal Express Corporation
Human Resource Development
3035 Director Row, Suite 501
Memphis, TN 38131

Mr. Gary Thomasson
University of Illinois
Educational Psychology
Champaign, IL 61820

Dr. Howard Wainer
Educational Testing Service
Princeton, NJ 08541

Elizabeth Wald
Office of Naval Technology
Code 227
800 North Quincy Street
Arlington, VA 22217-5000

Dr. Michael T. Waller
University of
Wisconsin-Milwaukee
Educational Psychology Dept.
Box 413
Milwaukee, WI 53201

Dr. Ming-Mei Wang
Educational Testing Service
Mail Stop 03-T
Princeton, NJ 08541

Dr. Thomas A. Warm
FAA Academy
P.O. Box 25082
Oklahoma City, OK 73125

Dr. David J. Weiss
N660 Elliott Hall
University of Minnesota
75 E. River Road
Minneapolis, MN 55455-0344

Dr. Douglas Wetzel
Code 15
Navy Personnel R&D Center
San Diego, CA 92152-6800

German Military
Representative
Personalstammamt
Koelner Str. 262
D-5000 Koeln 90
WEST GERMANY

Dr. Sherrie Gott
AFHRL/MOMJ
Brooks AFB, TX 78235-5601

Dr. Bert Green
Johns Hopkins University
Department of Psychology
Charles & 34th Street
Baltimore, MD 21218

Prof. Edward Haertel
School of Education
Stanford University
Stanford, CA 94305-3096

Dr. Ronald K. Hambleton
University of Massachusetts
Laboratory of Psychometric
and Evaluative Research
Hills South, Room 152
Amherst, MA 01003

Dr. Delwyn Harnisch
University of Illinois
51 Gerty Drive
Champaign, IL 61820

Dr. Patrick R. Harrison
Computer Science Department
U.S. Naval Academy
Annapolis, MD 21402-5002

Ms. Rebecca Hetter
Navy Personnel R&D Center
Code 13
San Diego, CA 92152-6800

Dr. Thomas M. Hirsch
ACT
P. O. Box 168
Iowa City, IA 52243

Dr. Paul W. Holland
Educational Testing Service, 21-T
Rosedale Road
Princeton, NJ 08541

Prof. Lutz F. Hornke
Institut fur Psychologie
RWTH Aachen
Jaegerstrasse 17/19
D-5100 Aachen
WEST GERMANY

Ms. Julia S. Hough
Cambridge University Press
40 West 20th Street
New York, NY 10011

Dr. William Howell
Chief Scientist
AFHRL/CA
Brooks AFB, TX 78235-5601

Dr. Huynh Huynh
College of Education
Univ. of South Carolina
Columbia, SC 29208

Dr. Martin J. Ippel
Center for the Study of
Education and Instruction
Leiden University
P. O. Box 9555
2300 RB Leiden
THE NETHERLANDS

Dr. Robert Jannarone
Elec. and Computer Eng. Dept.
University of South Carolina
Columbia, SC 29208

Dr. Kumar Joag-dev
University of Illinois
Department of Statistics
101 Illini Hall
725 South Wright Street
Champaign, IL 61820

Professor Douglas H. Jones
Graduate School of Management
Rutgers, The State University
of New Jersey
Newark, NJ 07102

Dr. Brian Junker
Carnegie-Mellon University
Department of Statistics
Pittsburgh, PA 15213

Dr. Marcel Just
Carnegie-Mellon University
Department of Psychology
Schenley Park
Pittsburgh, PA 15213

Dr. J. L. Kaiwi
Code 442/JK
Naval Ocean Systems Center
San Diego, CA 92152-5000

Dr. Michael Kaplan
Office of Basic Research
U.S. Army Research Institute
5001 Eisenhower Avenue
Alexandria, VA 22333-5600

Dr. Jeremy Kilpatrick
Department of
Mathematics Education
105 Aderhold Hall
University of Georgia
Athens, GA 30602

Ms. Hae-Rim Kim
University of Illinois
Department of Statistics
101 Illini Hall
725 South Wright St.
Champaign, IL 61820

Dr. Jwa-keun Kim
Department of Psychology
Middle Tennessee State
University
Murfreesboro, TN 37132

Dr. Sung-Hoon Kim
KEDI
92-6 Umyeon-Dong
Seocho-Gu
Seoul
SOUTH KOREA

Dr. G. Gage Kingsbury
Portland Public Schools
Research and Evaluation Department
501 North Dixon Street
P. O. Box 3107
Portland, OR 97209-3107

Dr. William Koch
Box 7246, Meas. and Eval. Ctr.
University of Texas-Austin
Austin, TX 78703

Dr. James Kraatz
Computer-based Education
Research Laboratory
University of Illinois
Urbana, IL 61801

Dr. Patrick Kyllonen
AFHRL/MOEL
Brooks AFB, TX 78235

Ms. Carolyn Laney
1515 Spencerville Rod
Spencerville, MD 20868

Richard Lanterman
Commandant (G-PWP)
US Coast Guard
2100 Second St., SW
Washington, DC 20593-0001

Dr. Michael Levine
Educational Psychology
210 Education Bldg.
1310 South Sixth Street
University of IL at
Urbana-Champaign
Champaign, IL 61820-6990

Dr. Charles Lewis
Educational Testing Service
Princeton, NJ 08541-0001

Mr. Hsin-hung Li
University of Illinois
Department of Statistics
101 Illini Hall
725 South Wright St.
Champaign, IL 61820

Library
Naval Training Systems Center
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Orlando, FL 32826-3224

Dr. Marcia C. Linn
Graduate School
of Education, EMST
Tolman Hall
University of California
Berkeley, CA 94720

Dr. Robert L. Linn
Campus Box 249
University of Colorado
Boulder, CO 80309-0249

Logicon Inc. (Attn: Library)
Tactical and Training Systems
Division
P.O. Box 85158
San Diego, CA 92138-5158

Dr. Richard Luecht
ACT
P. O. Box 168
Iowa City, IA 52243

Dr. George B. Macready
Department of Measurement
Statistics & Evaluation
College of Education
University of Maryland
College Park, MD 20742

Dr. Evans Mandes
George Mason University
4400 University Drive
Fairfax, VA 22030

Dr. Paul Mayberry
Center for Naval Analysis
4401 Ford Avenue
P.O. Box 16268
Alexandria, VA 22302-0268

Dr. James R. McBride
HumRRO
6430 Elmhurst Drive
San Diego, CA 92120

Mr. Christopher McCusker
University of Illinois
Department of Psychology
603 E. Daniel St.
Champaign, IL 61820

Dr. Robert McKinley
Educational Testing Service
Princeton, NJ 08541

Dr. David Wiley
School of Education
and Social Policy
Northwestern University
Evanston, IL 60208

Dr. Bruce Williams
Department of Educational
Psychology
University of Illinois
Urbana, IL 61801

Dr. Mark Wilson
School of Education
University of California
Berkeley, CA 94720

Dr. Eugene Winograd
Department of Psychology
Emory University
Atlanta, GA 30322

Dr. Martin F. Wiskoff
PERSEREC
99 Pacific St., Suite 4556
Monterey, CA 93940

Mr. John H. Wolfe
Navy Personnel R&D Center
San Diego, CA 92152-6800

Dr. Kentaro Yamamoto
03-0T
Educational Testing Service
Rosedale Road
Princeton, NJ 08541

Ms. Duanli Yan
Educational Testing Service
Princeton, NJ 08541

Dr. Wendy Yen
CTB/McGraw Hill
Del Monte Research Park
Monterey, CA 93940

Dr. Joseph L. Young
National Science Foundation
Room 320
1800 G Street, N.W.
Washington, DC 20550