TRENDS IN THE MEASUREMENT OF U.S. EMPLOYMENT STATUSES, 1972-1987

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I. INTRODUCTION. This paper has two goals: first, to describe trends in random measurement error (measurement variability or reliability) in classifications of employment statuses of the Current Population Survey (CPS) between 1972 and 1987. During this period, the percent of adult women who were members of the labor force increased from about 45 in 1973 to about 55 in 1985 (Bureau of Labor Statistics, 1986; U.S. Bureau of the Census, 1986).

During the same period, CPS procedures for measuring employment statuses were unchanged. It seems plausible that some of the factors that caused increasing female labor force participation, such as changes in the composition and economic conditions of families, also affected the reliability of measurement.

The second goal is to compare two statistical models of measurement variability: the misclassification model (Bailar and Biemer, 1984; U.S. Bureau of the Census, 1985) and the loglinear model (Goodman, 1969; Hauser and Massagli, 1983). The sensitivity of measurement error statistics to variations in the prevalences of categories complicates the analysis of trends and other comparisons since many extraneous factors affect category prevalences. We show that the two models can be viewed as different strategies for analytically separating trends in measurement error from trends in category prevalences.

II. DATA. Table 1 presents the data for the first semi-annual period, Jan.-June, 1972, and for the last semi-annual period, Jan.-June, 1987. The data are in the form of interview-reinterview cross classifications computed separately for adult males and adult females. The complete data consist of male and female tables representing all semi-annual periods from Jan.-June, 1972 through Jan.-June, 1987, a total of 31 periods.

The data are restricted to the CPS national sample (called the "combined A and C samples" in U.S. Bureau of the Census, 1978), to nonagricultural workers (excluding the rare category "employed in agriculture"), and to what is called the CPS "20 percent" or "unreconciled" reinterview sample. Every month, about 1 out of 18 of the 60,000 or so CPS sample households is reinterviewed. Of the households that are reinterviewed, 80 percent are subject to further inquiries for the purpose of reconciling discrepancies between the interview and reinterview. The 80 percent sample is used to estimate measurement bias, while the 20 percent sample is used to estimate measurement variability.

Other design features are important for evaluating the assumptions of measurement models: The reinterview takes place, on average, about one week after the original interview. A higher percentage of reinterviews than of original interviews are conducted by telephone rather than by personal visit. A higher percentage of reinterviews than of original interviews are conducted by highly experienced interviewers.

The same questions are used to measure employment statuses in the two interviews. Household respondents are asked about their own and other adult members' work-related activities during the "reference week", the calendar week preceding the week in which the original interview takes place. A recoding procedure (Bureau of Labor Statistics and Bureau of the Census, 1976) is used to assign the five employment statuses presented in the margins of Table 1. "Employed persons" are those who did any work for pay during the reference week, who were temporarily absent from a job (that is, "not at work" or "NAW", as abbreviated in Table 1), or who did 15 or more hours of work in a family business. Those who worked less than 35 hours are classified as "part-time".

The "labor force" consists of the "employed" together with the "unemployed", i.e. those who did not have a job during the reference week but who had actively "looked for work" during the previous 3 weeks or who were waiting to report to a new job. All others are categorized as "not in the labor force" or "NILF", as abbreviated in Table 1.

III. MODELS AND MEASURES. A. <u>Misclassification model</u>: This approach analyzes each employment status separately. For example, to make inferences about the category "full-time" from the first contingency table in Table 1 (males in Jan.-June, 1972), the data are collapsed to the 2x2 cross classification of "Full" against all other categories combined:

		Interview	
		Full	Other
Reint.	Full	a=788	b= 47
	Other	c= 29	d= 661

Collapsing the data *a priori* might result in some loss or distortion of information in the 5x5 tables. In the sequel, we refer to the set of five 2 x 2 arrays obtained by cross classifying each category against all others as the *collapsed data*. The symbols "a", "b", "c", and "d" are used to denote the cell counts of the collaped data.

The misclassification model assumes that each measurement is vulnerable to two kinds of misclassification errors: "false positives" (classifications as "full" when the true status is "other") and "false negatives" (classifications as "other" when the true status is "full"). The probabilities of false positives and false negatives are assumed to have a joint probability distribution in the population.

Let y(j,t) = 1 if the observed datum for the j-th sample unit on the t-th measurement is "full" and y(j,t) = 0 if the observed datum is "other". Let u(j) = 1 if the true status for the j-th sample unit is "full" and u(j) = 0 otherwise. The misclassification model assumes that the initial interview (t = 1) and reinterview (t = 2) measurements are generated according to

$$y(j,t) = u(j) + e(j,t), \quad j=1,\ldots,n;\ t = 1, 2; \qquad (1)$$

where u(j) is the "true value" and e(j,t) is the "measurement error".

If u(j) = 0, the error e(j,t) can equal either 0 or 1 depending upon the probability of a false positive. If u(j) = 1, the error e(j,t) can equal either -1 or 0 depending upon the probability of a false negative. Note that u(j) is assumed unchanged between measurements.

If the data arise from a simple random sample, if the probabilities of misclassification are the same on both measurements, and if e(j, 1) and e(j, 2) are independent for all j, then the <u>index of inconsistency</u>, defined as

$$I = \frac{(b+c)/n}{[(a+b)(b+d)+(a+c)(c+d)]/n^2} \qquad (2)$$

is a consistent estimator of the ratio of simple response variance (i.e., the expectation over the sample of the variance in repeated measurements of the same unit) to the total variance (i.e., the sum of simple response variance and sampling variance). That is, I estimates "the <u>impact</u> of misclassification errors on the <u>total variance</u> of an observation" (U.S. Bureau of the Census, 1985, p. 231).

The assumption of simple random sampling is an approximation since the CPS is based on a clustered areal probability sample of households (U.S. Bureau of the Census, 1978b). The assumptions of iid repeated measurements for each unit j may be invalid in the CPS reinterview (O'Muircheartaigh, 1986). Since the interval between measurements is only one week, respondents might remember their previous responses, in which case e(j,1) and e(j,2) might not be independent. Since more reinterviews than original interviews are conducted by telephone and by experienced interviewers, e(j,1) and e(j,2) might not be identically distributed. Indeed, the last two assumptions are challenged by the rejection of the loglinear model of symmetry (below).

Since the assumptions may not be satisfied, note that I has an alternative interpretation: The numerator of (2) is the proportion inconsistent, while the denominator is the estimated expected proportion inconsistent under the model of independence of the interview and reinterview measurements. Since the denominator is computed from the row and column margins, it can be regarded as an adjustment for category prevalence. The adjustment is <u>ad hoc</u> because the independence model seldom fits this kind of data (Hauser and Massagli, 1983; see Tab. 1).

B. <u>Loglinear model</u>: Rather than collapsing the data <u>a priori</u>, this approach tests a variety of models to find out how to reduce or summarize the data in each 5x5 contingency table, if any reduction is empirically defensible. The particular loglinear model which satisfactorily fit the 31 male and 31 female 5x5 employment status tables is called "quasi-symmetry" or "QS" (e.g., Fienberg, 1980; Hauser and Massagli, 1983; Hout, <u>et al.</u>, 1987).

Let F(i,j) denote the expected cell frequency in the cell located at the intersection of the ith row and jth column, i = 1, 2, ..., 5; j = 1, 2, can be represented as follows:

$$F(i,j) = a(i)b(j)c(i,j), \quad i=1,\ldots,5;\ j=1,\ldots,5; \qquad (3)$$

where c(i,j) = c(j,i) for all i not equal to j. To uniquely define the parameters, the product of the row effects (a's), the product of the column effects (b's), and the products of interaction effects (c's) across any fixed row or column are conventionally set equal to unity. Under QS, the interaction effects, c(i,j)'s, are symmetric across the main diagonal, but row main effects do not generally equal the corresponding column main effects. That is, a(i) need not equal b(i).

Applied to each 5 x 5 table, QS has 6 degrees of freedom. Under the assumption that the cell counts of each table arise from a multinomial distribution (an approximation given the complex sampling design of CPS), iterative methods were used to compute the maximum likelihood estimates of expected cell frequencies under QS for each of the 62 tables. Only 4 of the 62 likelihood-ratio statistics, comparing fitted and observed cell frequencies, were significant at the .01 level, when compared to a chi-square distribution with 6 df. In short, QS provided an acceptable fit to almost all the tables.

Simpler models, those with fewer parameters and more degrees of freedom than QS, were much more commonly rejected. In particular, the misclassification model implies that the model of symmetry, obtained by setting a(i) = b(i) for all i in (3), should fit. Symmetry was rejected for 19 of the 62 tables at the .01 level. The major departure from symmetry is that it is more common for part-time to be reported in the original interview and full-time in the reinterview than vice-versa (see Tab. 1). Part-time work during the reference week might be easily forgotten by those who usually work full-time.

The model of quasi-independence (QI), obtained by setting c(i,j) in (4) equal to unity for all cells off the main diagonal, would have to fit the data to justify restricting attention to the collapsed data that are utilized by the survey error misclassification model (Hauser and Massagli, 1983). QI was rejected at the .01 level for each of the 62 tables.

QS implies that the information in each 5x5 table is described by a set of ten 2x2 tables, representing the 5-choose-2 or ten unordered pairs of employment statuses. Each 2x2 table is defined by the cell counts located at the intersection of a pair of row categories with the same pair of column categories. For example, information about inconsistencies involving switches of the categories "part-time" and "full-time" is contained in the following 2 x 2 table (Table 1, first panel):

		Interview Full	Interview Part
Reint.	Full	f(1,1)=788	f(1,2)= 37
	Part	f(2,1)= 14	f(2,2)=126

where f(i,j) denotes the cell count in cell (i,j). Since the reduction is based upon goodness-of-fit tests, we refer to the set of 10 "pairwise" tables as the uncollapsed data.

Sufficient statistics for the c(i,j)'s in (3) are the natural logs of the odds-ratios ("log odds-ratios") of the 10 pairwise tables, i.e.

$$LOR(i,j) = \ln\{ \frac{f(i,j)f(j,i)}{f(i,i)f(j,j)} \} \quad (4)$$

Like I, LOR in (4) is a measure of measurement inconsistency. That is, the larger (in this paper, closer to zero) LOR, the higher the inconsistency between original interview and reinterview measurements. Unlike I, which ranges between 0 and 2 (between 0 and 200, if, as in Table 2, it is multiplied by 100), LOR ranges between minus and plus infinity. LOR is always negative in this report because cases are clustered on the main diagonals. For most purposes, one would change LOR from negative to positive and report it as a measure of consistency (see (4)).

With respect to dependence upon prevalence, LOR has a unique property, namely invariance against row and column multiplications (Fienberg, 1980): If the counts in any row or column are multiplied by a constant, LOR is unchanged.

C. Comparison of measures: To compare the sensitivities of I and LOR to category prevalences, we examined the correlations between each measure and the percentage in class, as measured by the reinterview. The correlations were computed for all time series (n=31) of I and LOR, using both collapsed and uncollapsed data for both males and females. For the collapsed data for the two rarest response categories, NAW and Unemp., I was invariably significantly correlated with the percent in class, whereas LOR was in no case signicantly correlated. This agrees with the simulation results in U.S. Bureau of the Census (1985, p. 231): For highly skewed data, I depends, in a complex fashion, upon category prevalence, even when the misclassification probabilities are constant. For the more prevalent employment statuses, Full, Part, and NILF, on the other hand, we found that the correlations of both measures with prevalences were generally small and that the correlation between I and LOR was always greater than .93.

The sensitivity of I to category prevalence, when applied to rare categories, was also suggested by the autocorrelation functions and male-female cross-correlations of the time series of measurement error statistics.

Figure 1 illustrates this point. It shows the autocorrelation functions of I and LOR for the female NAW series (collapsed data). In the autocorrelation function of I, seasonal variation is indicated by negative autocorrelations at odd lags and positive autocorrelations at even lags (see Box and Jenkins, 1976, Chap. 2).

The seasonal pattern probably results from the contamination of I by changes in category prevalence rather than from any cyclical variation in measurement error. In particular, NAW tends to be about 50% more prevalent in the second half of the year when more workers go on vacation. The corresponding autocorrelation function of LOR evidences a similar pattern for the first three lags, but the magnitudes of the autocorrelations are much smaller.

If invariance to changes in category prevalences is a reasonable criterion for choosing a statistic and form of data, one prefers LOR applied to the uncollapsed data, as stipulated by the quasi-symmetry model. For the ten male and ten female series of this type, we found that the numbers of autocorrelations and male-female correlations that were significant at alpha = .05 approximately equaled the corresponding numbers expected under the hypothesis of "white noise" in all of the twenty time series. (Tests of significance of the autocorrelations were performed using Bartlett's test; Box and Jenkins, 1976, Chap. 2).

Details of the comparisons of the alternative measures are given in our longer paper (available upon request).

IV. TRENDS. Table 2 shows the mean I's and mean LOR's for collapsed and uncollapsed data computed separately for each sex in each of four time periods: 1972-75, 1976-79, 1980-84, and 1985-87. Also shown, for each sex and period, are the percentages of inconsistent responses which involved each category ("collapsed data") and each category pair ("uncollapsed data"). The base of each percentage is obtained by multiplying the sample size by the proportion inconsistent. (The percentages for the collapsed data add to 200 since each observation is counted twice.)

On the basis of either I or LOR, Table 2 suggests substantial time- and sex-invariance in the pattern of random measurement error. The rankings of employment statuses and of pairs of statuses according to the magnitudes of I and LOR are similar but not the same: I ranks the least prevalent category, NAW, more highly than does LOR with respect to the relative seriousness of measurement error.

For each of the n=31 semi-annual observations of each time series of LOR, the sampling variance was estimated using the sum of the reciprocals of the cell counts, a consistent estimator under multinomial sampling (Fienberg, 1980). For each observation of each series of I, a Taylor series approximation of the sampling variance, a consistent estimator under simple random sampling (U.S. Bureau of the Census, 1985), was computed. The variance estimates for semi-annual periods of each time series of each statistic were pooled to carry out two-sample t-tests comparing the average levels of the statistic in 1972-75 and 1985-87.

We inferred only one consistent difference at the .05 level: For both males and females, using either I or LOR, measurement variability involving switches of the categories Part and NILF is smaller in 1985-87 than in 1972-75. For both sexes, the magnitude of the estimated change, relative to the initial level of the series, is greater when I is used than when LOR is used.

Perhaps with the increase in its prevalence, part-time work has become less episodic, more highly institutionalized, and more readily distinguished from NILF by respondents. Inspection of the trends across all four periods in Table 2 suggests that the Part:NILF type of inconsistency declined most rapidly between 1972-75 and 1976-79.

We also investigated male-female differences and found one difference that was persistent over time: According to both I and LOR, applied to collapsed data, part-time work is reported less reliably for males than for females (see Tab. 2). Perhaps part-time work is more frequently of short duration, hence more readily forgotten, among males than among females.

The detailed description provided by the uncollapsed data suggests hypotheses for future research: For each sex in each period, more than 50 percent of the measurement inconsistencies involved either Full:Part or Unemp.:NILF. Perhaps Full and Part are frequently confused because the distinction between less than 35 hours and 35 or more is arbitrary and easily forgotten. Perhaps Unemp. and NILF are confused because the concept "looking for work" is ambiguous for many persons.

V. CONCLUSIONS. There are two methodological conclusions: 1) When applied to highly skewed data, the misclassification approach (i.e., the index of inconsistency applied to collapsed data) is more sensitive to variations in category prevalence than the loglinear approach (i.e., the log odds-ratio applied to reduced data based upon a model which fits the data); 2) Applied to the CPS employment status tables, the loglinear approach preserved useful information about pairs of categories that was not preserved by the misclassification approach.

There are three substantive conclusions: 1) During 1972-1987, males reported part-time work less reliably than females; 2) Measurement variability involving switches of the employment statuses "part-time" and "not in the labor force" declined between 1972 and 1987; 3) Otherwise, the pattern of measurement variability in CPS employment statuses appears sex- and time-invariant during 1972-1987. The distinctions between "full-time" and "part-time" and between "unemployed" and "not in the labor force" are most likely to give rise to inconsistent reports.

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FOOTNOTES: 1. This paper reports the general results of research undertaken by Census Bureau staff. The views expressed are attributable to the authors and do not necessarily reflect those of the U.S. Census Bureau. Thanks to Larry Cahoon, Donna Kostanich, Elizabeth Martin, Gary Shapiro, and Henry Woltman for helpful comments.

FIGURE 1

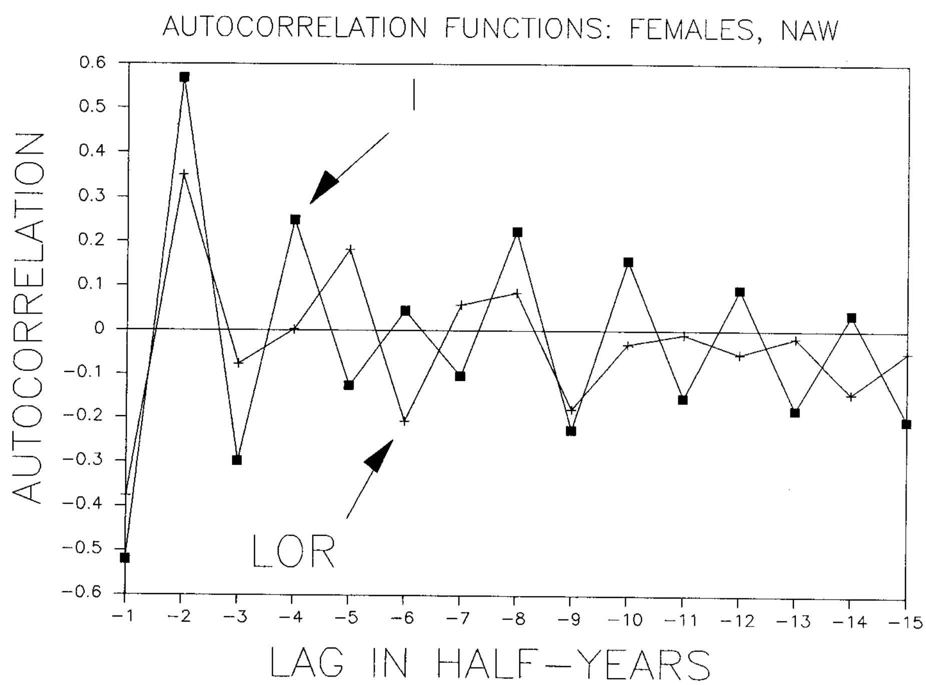

Table 1. Cross classifications of original interview and reinterview measurements of employment statuses*. CPS Reinterview twenty percent (unreconciled) sample. National sample design. Males and females, aged 16 and over. Jan.-June, 1972 and Jan.-June, 1987.

MALES, JAN.-JUNE, 1972

	Original interview						
Reinterview	Full	Part	NAW	Unemp.	NILF	Total	Percent
Full-time	788	37	3	3	4	835	54.8
Part-time	14	126	2	4	8	154	10.1
Not at work	5	0	26	0	4	35	2.3
Unemployed	4	2	1	49	12	68	4.5
NILF	6	8	4	12	403	433	28.4
Total	817	173	36	68	431	1525	100.0
Percent	53.5	11.3	2.4	4.5	28.3	100.0	

MALES, JAN.-JUNE, 1987

	Original interview						
Reinterview	Full	Part	NAW	Unemp.	NILF	Total	Percent
Full-time	1006	47	6	7	8	1074	53.7
Part-time	39	146	8	4	14	211	10.6
Not at work	2	4	28	4	2	40	2.0
Unemployed	10	8	1	78	19	116	5.8
NILF	5	7	3	16	528	559	28.0
Total	1062	212	46	109	571	2000	100.0
Percent	53.1	10.6	2.3	5.5	28.6	100.0	

FEMALES, JAN.-JUNE, 1972

	Original interview						
Reinterview	Full	Part	NAW	Unemp.	NILF	Total	Percent
Full-time	407	34	3	2	10	456	24.4
Part-time	9	191	0	1	14	215	11.5
Not at work	1	5	15	3	1	25	1.3
Unemployed	0	0	0	37	19	56	3.0
NILF	4	16	5	19	1076	1120	59.8
Total	421	246	23	62	1120	1872	100.0
Percent	22.5	13.1	1.2	3.3	59.8	100.0	

FEMALES, JAN.-JUNE, 1987

	Original interview						
Reinterview	Full	Part	NAW	Unemp.	NILF	Total	Percent
Full-time	638	58	7	5	10	718	30.5
Part-time	35	325	7	1	13	381	16.2
Not at work	2	3	33	1	5	44	1.9
Unemployed	5	1	1	52	24	83	3.5
NILF	5	21	6	13	1083	1128	47.9
Total	685	408	54	72	1135	2354	100.0
Percent	29.1	17.3	2.3	3.1	48.2	100.0	

*See Section II for definitions.

Table 2. For each of four periods, mean index of inconsistency (I), mean log-odds ratio (LOR), and percent of total inconsistent responses involving category or category pair (%). Collapsed and uncollapsed data for males and females.

MALES

Category	Period and Statistic											
	1972-75			1976-79			1980-84			1985-87		
or Pair	I	LOR	%	I	LOR	%	I	LOR	%	I	LOR	%
Full-time	13	-5.4	56	13	-5.5	59	13	-5.3	56	12	-5.5	57
Part-time	32	-4.2	57	32	-4.3	60	33	-4.1	59	32	-4.3	55
Not at work	35	-5.2	20	32	-5.5	20	36	-5.2	19	33	-5.5	18
Unemployed	33	-5.0	26	31	-5.1	26	28	-5.1	28	29	-5.1	28
NILF	12	-5.8	41	10	-6.2	36	11	-5.9	38	11	-6.0	41
Full:Part	24	-4.7	37	26	-4.5	42	26	-4.4	39	24	-4.7	36
Full:NAW	20	-6.2	10	17	-6.6	9	19	-6.2	8	17	-7.1*	8
Full:Unemp	7	-8.6	4	5	-9.0	3	4	-9.1	3	7	-8.2	5
Full:NILF	2	-9.5	6	2	-10.0	5	2	-9.3	6	3	-8.8	9
Part:NAW	12	-6.6	4	11	-6.5	4	15	-5.7	5	13	-6.0	4
Part:Unemp	9	-6.4	4	9	-6.7	4	9	-6.3	5	10	-6.2	5
Part:NILF	12	-5.9	13	8	-6.8	10	9	-6.5	10	9*	-6.5*	11
NAW:Unemp	8	-6.9	2	9	-6.9	2	10	-6.6	3	7	-7.2	2
NAW:NILF	13	-6.7	5	11	-7.1	5	11	-7.3	4	12	-7.0	5
Unemp:NILF	26	-4.6	16	25	-4.8	16	22	-4.8	18	22	-4.9	17
Base N	16994			17690			23652			10476		
%Inconsistent	11.3			10.7			11.8			10.7		

FEMALES

Category	Period and Statistic											
	1972-75			1976-79			1980-84			1985-87		
or Pair	I	LOR	%	I	LOR	%	I	LOR	%	I	LOR	%
Full-time	13	-5.7	46	13	-5.7	48	13	-5.6	48	13	-5.6	52
Part-time	26	-4.6	58	25	-4.6	61	25	-4.6	62	24	-4.7	61
Not at work	39	-5.3	17	34	-5.6	15	34	-5.5	16	32*	-5.6	16
Unemployed	37	-5.0	24	35	-5.0	25	34	-5.0	26	34	-5.1	25
NILF	11	-5.6	54	11	-5.8	51	10	-5.8	49	10	-5.9*	47
Full:Part	20	-4.6	33	21	-4.5	40	20	-4.5	39	20	-4.6	44
Full:NAW	16	-6.6	6	12	-7.1	5	14	-6.8	5	15	-6.7	6
Full:Unemp	6	-8.9	3	3	-9.9	1	3	-9.7	2	5	-9.1	3
Full:NILF	3	-8.8	9	3	-9.1	9	2	-9.7	7	2	-9.2	8
Part:NAW	17	-5.9	5	15	-6.3	5	16	-6.0	5	12	-6.5	4
Part:Unemp	7	-7.9	3	7	-7.4	4	7	-7.1	4	8	-7.3	4
Part:NILF	12	-6.2	24	9	-6.5	21	9	-6.5	20	8*	-6.7*	19
NAW:Unemp	8	-7.0	1	7	-7.2	2	6	-7.4	1	5	-7.7	1
NAW:NILF	20	-6.8	7	17	-7.0	7	15	-7.4	6	15	-7.0	6
Unemp:NILF	31	-5.1	20	30	-4.9	22	29	-4.9	21	28	-5.1	21
Base N	20158			20524			28111			12315		
%Inconsistent	10.4			10.4			10.7			10.5		

*I or LOR for 1985-87 is significantly different from corresponding statistic for 1972-75, 2-tailed test, alpha = .05.