

**THE SURVEY OF INCOME AND  
PROGRAM PARTICIPATION**

**Testing a New Attrition Nonresponse  
Adjustment Method for SIPP**

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## Testing A New Attrition Nonresponse Adjustment Method For SIPP

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### 1. Introduction

While unit or questionnaire nonresponse can seriously degrade the quality of any survey, nonparticipation is particularly threatening to a longitudinal survey like the U.S. Census Bureau's Survey of Income and Program Participation (SIPP). To minimize the potential biasing effects of second and subsequent wave attrition from SIPP panels, staff at the Census Bureau perform weight adjustments. These adjustments are designed to make the self-selected subsample of longitudinal respondents more representative of the initial wave one sample. The Bureau's SIPP weight adjustments take the form of post-stratum or weighting class specific multipliers applied to the wave one base sample weight. The associated weighting classes are defined by collapsing cells in a multi-way cross classification of categorical variables until each resulting cell satisfies two conditions: cells are collapsed until the sample size and the estimated response propensity are greater than predefined thresholds.

The Bureau funded research project reported here (Folsom and Witt, 1994) tests a new nonresponse adjustment methodology recently developed at the Research Triangle Institute (RTI) (Folsom, 1991). This new method uses weight adjustment multipliers defined at the person level. These multipliers are created by modeling a sample person's response propensity using constrained forms of either a logistic or exponential model. This nonresponse adjustment methodology was tested on data from the 1987 SIPP panel. The goal of this project was to develop a new SIPP nonresponse adjustment that would reduce the attrition bias in cross sectional and longitudinal estimates derived from the survey.

### 2. Overview of the 1987 SIPP

The 1987 SIPP was a longitudinal household, panel survey designed to collect demographic and economic data on all household occupants over a 28 month data collection period. The initial sample of households was divided into four *rotation groups*, and data was collected from respondents in each rotation group every four months. Each four month data collection period is referred to as a sample *wave*; consequently,

the 1987 SIPP had seven waves of data collection. During each interview, demographic and economic data was collected by month, for the four previous months. By carefully matching waves of data collection, rotation groups and reference months one can see that monthly data is available for all four rotation groups for 25 continuous reference months and monthly data for subsets of rotation groups are available for an additional 6 months.

The SIPP was designed to be an ongoing survey with new, seven or eight wave samples (or panels) introduced each year. In our evaluation of the new weight adjustment, we take advantage of this panel overlap noting that the reference period for the first wave of the 1989 SIPP is equivalent to the reference period in the last wave of the 1987 SIPP. To increase the length and sample size of individual panels without increasing survey costs, the Bureau plans to change the design of future SIPP surveys to nonoverlapping panels with 48 months of data collection. The impetus for this research project was based on the likelihood that the increase in panel length and associated respondent burden would increase the attrition rate and the associated nonresponse bias.

At the first wave of data collection for the 1987 SIPP, a sample of roughly 11,700 responding households was established. For the months in which they remained survey eligible, person-level data was sought on residents of these households at each subsequent wave of the SIPP panel. Generally, people can become ineligible for the survey if they died or move out of the country. Within this sample of first wave responding households, a person sample of roughly 33,100 was identified.

While SIPP person nonresponse occurs at all waves of data collection, the wave specific attrition rates decline monotonically through the sample waves. At wave one, 30,767 persons were interviewed and of these, 24,429 responded for each month in which they were eligible during the seven waves of data collection.

For this project, we were interested in developing a nonresponse adjustment for the SIPP longitudinal respondent weights. Longitudinal respondents are wave one participants that provide data for all subsequent waves in which they were eligible. Consequently, this project began with an initial

sample of 30,767 people; 24,429 longitudinal respondents and 6,338 nonrespondents for a 20.6% attrition rate. The initial sample weight (or base weight) for the 30,767 people used for this analysis retained the Bureau's adjustment for wave one household and person nonresponse.

One appealing feature of this two step (wave one plus subsequent attrition) weight adjustment is that the wave one data can be used as explanatory information in our models of longitudinal response propensity. Since our response propensity models can effectively incorporate more of these wave one variables than the weighting class method, the potential clearly exists for a significant reduction in attrition bias.

Before launching into a description of how our new weights were created, it will be useful to consider how the Census Bureau currently creates their longitudinal nonresponse adjustment. The Bureau refers to their weighting class adjustment for longitudinal nonresponse as the *first stage* adjustment. Weighting classes are formed using such wave one auxiliary information as race, education, welfare and unemployment benefits indicators, a labor force status indicator, a bonds indicator and categorized average monthly household income (126 total classes). The nonresponse adjustment is defined as the inverse of the base weighted longitudinal response rate observed within each weighting class. In order to minimize the effect of unequal weights on the variance of estimates, weighting classes are collapsed to avoid cells with less than 30 longitudinal respondents or cells with inverse response rates (nonresponse adjustments) exceeding 2.00.

After performing the weighting class adjustments, Census Bureau statisticians performed a *second stage* adjustment (also commonly referred to as a post-stratification adjustment) to January 1987, person level control totals derived from the Current Population Survey (CPS). The second stage adjustment was applied using standard raking methods in order to preserve CPS counts for the cross classifications defined by:

- Age Group x Race x Gender x Hispanic Indicator. Several cells in this cross-classification were collapsed; 98 population totals initially controlled for.
- Complex Interaction Of Householder Status, Living With Relative Indicator, Spouse and/or Child Present Indicator; 19 population totals initially controlled for.

As with the first stage adjustment procedure, marginal cells in the second stage adjustment are collapsed when the total number of respondents is less than 30 or the marginal adjustment is greater than 2.00.

### 3. Exponential and Logistic Model

Extending a constrained exponential model suggested by Deville and Särndal (Deville and Särndal, 1992), our constrained exponential and logistic weight adjustment multipliers may be written as:

$$\lambda_i = \begin{cases} \alpha_o^{-1} \alpha_i & \text{\{Exponential Model\}} \\ (1 + \alpha_o^{-1} \alpha_i) & \text{\{Logistic Model\}} \end{cases} \quad (1)$$

Where:

$$\alpha_i = \left\{ \frac{\hat{L}(\hat{U}-1) + \hat{U}(1-\hat{L})\exp(-AX_i\beta)}{(\hat{U}-1) + (1-\hat{L})\exp(-AX_i\beta)} \right\} \quad (2)$$

$i$  = Indexes the sample units (wave one respondents),

$X_i$  = Vector of explanatory variables known for all  $i$ ,

$\lambda_i$  = the inverse response propensity ( $\rho_i^{-1}$ ) weight adjustment,

$A$  =  $(\hat{U}-\hat{L})/(1-\hat{L})(\hat{U}-1)$ . This is simply a scale factor that helps minimize the effects of the constraints on the shape of the exponential function.

The scale factors  $\alpha_o$  are derived from the base sample weighted overall response rate,  $\rho_o$ ; specifically,

$$\alpha_o = \begin{cases} \rho_o & \text{\{Exponential Model\}} \\ \rho_o + (1-\rho_o) & \text{\{Logistic Model\}} \end{cases} \quad (3)$$

$$\rho_o = (\sum w_i r_i) \div (\sum w_i),$$

$r_i$  = 0/1 Sample response indicator, and

$w_i$  = Sample base weight.

And the bounds  $\hat{L}$  and  $\hat{U}$  in  $\alpha_i$  [see equation (2)] are set as follows:

$$\begin{aligned} \hat{L} &= \alpha_o L, & \hat{U} &= \alpha_o U & \text{\{Exponential Model\}} \\ \hat{L} &= \alpha_o (L-1), & \hat{U} &= \alpha_o (U-1) & \text{\{Logistic Model\}} \end{aligned} \quad (4)$$

Our estimates for  $\beta$  are found by solving the following generalized raking equation using a Newton-Raphson algorithm:

$$\sum w_i r_i \lambda_i X_i = \sum w_i X_i \quad (5)$$

Notice, the constants  $U$  and  $L$  in (4) are chosen to bound (or constrain) the resulting adjustment factor  $\lambda_i$ :

As  $X_i \beta \Rightarrow +\infty$  then  $\alpha_i \Rightarrow \hat{L}$  in (2) and  $\lambda_i \Rightarrow L$  in (1),

As  $X_i \beta \Rightarrow -\infty$  then  $\alpha_i \Rightarrow \hat{U}$  in (2) and  $\lambda_i \Rightarrow U$  in (1).

Bounding the adjustments constrains the associated variance inflation which results from the increase in adjusted weight variability compared to the original base weights.

In order to obtain a solution to (2),  $\hat{L}$  and  $\hat{U}$  must satisfy:  $0 < \hat{L} < 1 < \hat{U}$ . Notice in (4) that the scale factors  $\alpha_o$  are introduced to simply shift the limits on  $\hat{L}$  and  $\hat{U}$ . This shifting allows us to obtain feasible inverse response propensities from the exponential and logistic models as follows:

#### Exponential Model:

$\hat{L} \in (0, 1)$  shifts to  $L \in (0, \rho_o^{-1})$

$\hat{U} \in (1, \infty)$  shifts to  $U \in (\rho_o^{-1}, \infty)$ .

Without this shift, we note from (1) that a lower bound could not be set that would force the  $\lambda_i$ 's to be greater than one since the lower bound on  $\alpha_i$  could not be set to one without forcing all  $\alpha_i$  to equal one uniformly.

#### Logistic Model:

$\hat{L} \in (0, 1)$  shifts to  $L \in (1, \rho_o^{-1})$  and

$\hat{U} \in (1, \infty)$  shifts to  $U \in (\rho_o^{-1}, \infty)$ .

Without this shift, equation (1) implies that an upper bound could not be chosen that would force  $\lambda_i$ 's to be less than or equal to 2 since the upper bound on  $\alpha_i$  could not be set to one without again forcing uniformity on all  $\alpha_i$ .

Thus the scale factor is introduced in order to allow one to use either the logistic or exponential response propensity model and achieve the desirable property:  $1 \leq L \leq \lambda_i \leq U$  where  $U$  and  $L$  are predetermined constants.

For the unconstrained cases, (i.e. when  $\alpha_o = 1$ ,  $\hat{U} = +\infty$ , and  $\hat{L} = 0$ ), the terms  $A \rightarrow 1$ , and  $(\hat{U} - 1)^{-1} \rightarrow 0$  in equation (2). In these cases, the model reduces to the familiar form:

$$\rho_i = \begin{cases} [\exp(-X_i \beta)]^{-1} & \{\text{Exponential Model}\} \\ [1 + \exp(-X_i \beta)]^{-1} & \{\text{Logistic Model}\} \end{cases}$$

Folsom (1991) proposed the unconstrained logistic model for nonresponse adjustment and the unconstrained exponential model for sampling error adjustments akin to post-stratification or double sampling ratio estimation. In particular, Folsom (1991) discusses the variance and bias reduction properties which result from an equivalence between logistic re-weighted respondent means and regression imputation-based estimates, and between exponential reweighted sample means and a survey regression estimator.

In practice, we most often use the exponential model for second stage, post-stratification type adjustments since the required adjustment factors are not logically bound below by one as they are with adjustments resulting from the logistic response propensity model. Further, to facilitate obtaining a solution to the raking equations (5), we tend to use the scaled versions of the exponential model for nonresponse adjustment when the over-all response rate  $\rho_o$  is close to one. For our application to the 1987 SIPP data, we used the scaled logistic model to predict longitudinal response propensity. With the overall response rate  $\rho_o = 0.8$ , we were able to obtain convergent solutions for each of our subpopulation models with  $U=2$ , the Bureau's weighting class adjustment bound.

#### 4. Estimating Model Parameters

We began the development of the alternate nonresponse adjustments for the 1987 SIPP by examining marginal response rates across several main effect and low order interaction terms in order to specify an initial set of response propensity predictor variables. During this search we found that testing the statistical significance of the univariate response rate differentials across the levels of the categorical variables was not a useful indication of a variable's predictive power. The combined wave one sample size of 30,767 guaranteed that most of the associated variable specific chi-squared tests would exhibit highly significant results. This power to declare negligible response rate differentials statistically significant persisted when proper account was taken of the SIPP design induced clustering effects. Therefore, for this initial screening, we chose to disregard variables with response rate differentials that fail to exceed a subjective threshold of 10 percentage points. We observed that age group, race/ethnicity, relationship to reference person, living quarters owned indicator, race/ethnicity of the head of

household, household type (levels), wage dollars imputed indicator, the source dollars imputed indicator, covered by Medicare indicator, and covered by General Assistance indicator all exhibited a response rate differentials approaching or exceeding 10 percent.

This initial look at response rate differentials, coupled with some exploratory modeling to estimate the significance of various lower-order interactions, continuous variables and linear spline functions of several variables, led us to eventually build separate models for seven subsamples (nonresponse classes) defined in terms of household income, race/ethnicity, marital status, and Census Region. Within each class, we began with models containing a large number of explanatory variables, and eliminated statistically nonsignificant parameters using a backwards elimination process. The statistical test to determine the significance of the parameters was based on students-T type statistics derived from sample design-based (cluster sampling) variances estimated using the Taylor Series method. The level of significance used for these tests was set at  $\alpha=.10$ . The variables and  $\beta$  coefficients retained in our final models are presented in Table 1.

In order to minimize the effect of these adjustments on the coefficient of variation of the resulting respondent sample weights, a lower bound of 1.0 and an upper bound of 2.0 was set in the scaled logistic models for each nonresponse class. Further, generalized Wald statistics adjusted for design effects were created to test the overall significance of each model. These Wald statistics test the null hypothesis that all slope parameters are zero. The significance probabilities of the Wald statistics are presented at the bottom of Table 1.

While we tested for several two-way interactions among the significant main effects in our models, very few significant interactions were found. One should note however that by virtue of fitting separate models within the seven classes, we have implicitly interacted the nonresponse class variable with all the fitted main effects.

##### **5. Evaluation Of Alternate Weights**

Our new longitudinal sample weights were evaluated in terms of estimated relative bias, relative standard error and the error in 90% confidence interval (CI) coverages. The coverage error in the estimated 90% CIs approximates the probability that a population parameter lies outside the bounds of a

90% CI, assuming our bias and variance estimates are error free and the sampling distribution is normal. These evaluation statistics were computed using the new "RTI-Revised" weight and using the Bureau's "Current" nonresponse-adjusted longitudinal sample weight. Standard errors in this analysis were based on the linearization variance for cluster sample ratio means and proportions which treat the response propensities and second stage, post-stratification adjustments as known without error. Though this assumption may lead to bias in the variance estimates, our experience with variance estimators based on the proper generalized raking variances suggests that these biases are likely to be modest.

To evaluate the RTI-Revised weights, two comparative evaluations were conducted (Table 2):

- 1987 SIPP wave one estimates derived from longitudinal respondent data using the RTI-Revised and the Current weight were compared against benchmark statistics created using the original, 1987 SIPP wave one sample (30,767) with their household and nonresponse adjusted base sample weight. For this analysis, a second stage adjustment was applied to each weight using the scaled exponential model specified in Section 3 with explanatory variables representing all the effects currently used in the Bureau's second stage adjustment as specified in Section 2. No attempt was made to omit nonsignificant terms in the second stage adjustment model.
- 1987 SIPP wave seven estimates using the RTI-Revised weight and the Current weight were compared against benchmark statistics created using wave one data from the 1989 SIPP. Recall from Section 2 that the seventh wave of the 1987 SIPP overlaps the first wave of the independently selected 1989 SIPP. Further, except for births and immigrations added to the 1989 panel, there is no difference in the population coverage of the two panels. To minimize this population coverage difference, those 1989 SIPP respondents who were 0-2 years old in wave one were omitted from the analysis since they had not been born when the 1987 panel was established at wave one.

Intuitively, in our first analysis of the weights, the bias calculation derived from the 1987 panel wave one data should maximize the performance of our RTI-Revised weights since our generalized raking solution [equation (5)] forces mean equality within

our seven model subpopulations for wave one variables that are included in the model. Despite the fact that these first stage controls are perturbed by the global second stage adjustments, the results in the left side of Table 2 suggest that the bias in the RTI-Revised weighted statistics is considerably less than the bias in the Current-weighted estimates. For example, for the statistics grouped under the "PROPORTION RECEIVING INCOME FROM SOURCE", the median relative bias was .37% for RTI-Revised weighted statistics and 1.16% for the Current weighted statistics. Similarly, the median projected coverage error for a 90% CI was .128 for RTI-Revised weighted statistics and .236 for the Current weighted statistics.

The right half of Table 2 displays results from comparing the 1987 SIPP, wave seven estimates with estimates derived from the 1989, wave one SIPP. This 1989 analysis is clearly a more stringent test of our revised weights. One would expect that our wave one predictors of longitudinal response would be reasonably good predictors of the wave one data values missing from longitudinal nonrespondents records. On the other hand, we would expect these covariates to be less effective predictors of the nonrespondents' unobserved wave seven data.

The data in the right half of Table 2 show mixed results. The median relative bias in the RTI-Revised weighted estimates is slightly less for the statistics displayed under the "MEAN \$ AMOUNT AMONG RECIPIENTS" column and slightly more for the statistics displayed under the "PROPORTION RECEIVING INCOME FROM SOURCE" column. In general, the evaluation relative to the 1989 panel benchmark provides no evidence that the RTI-Revised weight has achieved any bias reduction compared to the Bureau's Current weight. Similar results were found when this analysis was repeated for several subgroups of the population, including groups defined by age group and race/ethnicity. Similar results were also found when the analysis was restricted to comparing only those estimates with significant biases.

## 6. Conclusion

The upcoming switch to a four year longitudinal SIPP with no overlapping panels places a high priority on making the most effective use of the statistical tools available for reducing attrition bias. We believe that inverse response propensity weighting via generalized raking is one of the more promising statistical methods designed for this purpose. Recent

developments in this area permit the imposition of arbitrary bounds on the weight adjustment in order to guard against undue variance inflation. Given the ability of these regression modeling approaches to accommodate large number of control variables, they are an obvious candidate for refining SIPP weight adjustment methods.

Our evaluation of the new method's performance relative to the Bureau's current weighting class approach was mixed. In the somewhat artificial test using the full wave one sample to produce wave one benchmark statistics, the RTI-revised statistics showed notable bias reductions compared to the Bureau's Current adjustments. However, the more realistic test of our method used benchmark statistics derived from wave one of the 1989 panel to compare with wave seven 1987 panel estimates. This test did not demonstrate any clear superiority for the new method. Given the mixed results shown by our two evaluations and indications that a serious attrition bias potential exists, we are still inclined to recommend a switch to a generalized raking approach. We are swayed by the intrinsic merits of the approach as suggested by the potential bias and variance reduction properties that flow from its equivalence to regression imputation (Folsom, 1991).

## References

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**Table 1. Summary of Nonresponse Adjustment Models**

Variable	Average HH Income: <\$1,200	Average HH Income: \$1,200 - \$1,400					Average HH Income: >\$4,000
		Hispanic/Black Householder	White, Not Married HH	White, Married Householder			
				North East	South	North Central, West	
<b>SELECTED MODEL COEFFICIENTS</b>							
Intercept Term	-1.52983	0.04111	-3.45717	-1.44214	-0.95339	-0.42018	-2.90626
0-14 Year Old Indicator	...	-0.02878	1.59519	1.28601	1.25917	...	0.72501
15-30 Year Old Indicator	...	-0.71952	-0.38678	0.08777	-0.03305	...	0.14529
31-64 Year Old Indicator	...	-0.43718	-0.27720	0.03763	0.16870	...	0.31214
Splined Age	**	...	...	...	...	**	...
Hispanic Indicator	-0.25147	...	...	...	...	...	-0.71238
Black Indicator	-0.62689	...	...	...	...	...	-0.89493
Hispanic Household Indicator	...	0.38886	...	...	...	...	...
Widowed, Divorced, Separated Indicator	...	...	-0.26265	...	...	...	...
0-11 Years of Education	-0.45158	...	...	...	-0.33810	...	-0.40952
12-15 Years of Education	-0.32393	...	...	...	-0.36870	...	-0.43848
Census Region	...	...	**	...	...	**	**
# Adults in Household Indicator	**	...	**	...	...	**	**
Female Household, No Spouse Present	...	...	0.53918	...	...	...	...
Married Household	0.87507	...	...	...	...	...	...
Person is Household Reference Person	...	...	0.83465	0.59289	0.56874	...	2.47107
Person is Spouse Reference Person	...	...	...	0.44904	0.43029	...	2.41419
Person is Child/Relative to Reference Person	...	...	0.75732	...	...	...	1.81897
In Labor Force Indicator	...	...	...	...	...	...	-0.76688
Government Worker Indicator	...	...	0.67267	...	0.99177	...	...
Self Employment Indicator	...	...	...	...	...	...	-0.44274
Bonds Indicator	...	...	...	...	...	0.74748	...
Bonds x North Central Interaction	...	...	...	...	...	-0.75619	...
Disability Indicator	...	...	...	...	...	...	-0.93429
HH Gets Govt Energy Assistance	0.50300	...	...	...	...	...	...
Anyone in HH Get SSI Indicator	0.65596	...	...	...	...	...	...
Living Quarters Owned Indicator	0.70684	...	0.38527	0.73702	0.37361	...	0.86928
% of Reported Assets Imputed	**	**	**	**	...	...	**
Any Reported Wage/Self \$ Imputed	...	...	...	-1.76532	...	-1.15562	...
Reported Wage/Self \$ Not Imputed	...	...	...	...	...	...	0.77846
Any Income Source \$ Imputed	...	...	0.58213	...	...	...	...
Personal Income Quartiles	...	...	...	...	**	...	...
Spline Personal Income	...	...	**	...	**	**	**
Region x Personal Income Interaction	...	...	...	...	...	**	...
Covered By Food Stamps Indicator	0.48852	...	...	...	...	...	...
Covered by Social Security, 0-64 Yrs. Old, Indicator	0.42471	...	...	...	...	...	...
<b>MODEL SUMMARY STATISTICS</b>							
Original (1+CV <sup>2</sup> <sub>weights</sub> )	1.023	1.018	1.020	1.018	1.020	1.015	1.011
Final (1+CV <sup>2</sup> <sub>weights</sub> )	1.046	1.029	1.049	1.029	1.026	1.026	1.040
Max. Wt. Adjustment Obtained	1.973	1.747	1.960	1.901	1.542	1.628	1.991
Min. Wt. Adjustment Obtained	1.006	1.117	1.008	1.056	1.035	1.035	1.020
Wald Statistic Significance Prob.	.000	.003	.003	.001	.000	.000	.000

\*\* Several Categories of Variables Included in Final Model. Beta Coefficient Not Presented in Table.

**Table 2. SIPP Estimates Computed Using RTI-Revised Weight and Current SIPP Sample Weight**

Independent Variable	January, 1987 Estimates from 1987 SIPP Wave 1 Sample <sup>1</sup>					January, 1989 Estimates from 1987 SIPP Wave 7/1989 SIPP Wave 1 Sample <sup>1</sup>								
	Estimates		% Abs Rel Bias		90% CI Coverage	Estimates		% Abs Rel Bias		90% CI Coverage				
	Current <sup>2</sup>	RTI-Revised <sup>2</sup>	Bench-Mark	Current	RTI-Revised	Current <sup>2</sup>	RTI-Revised <sup>2</sup>	Bench-Mark	Current	RTI-Revised				
<b>MEAN \$ AMOUNT AMONG RECIPIENTS</b>														
Personal Income	1,384	1,395	1,393	0.64	0.17	0.181	0.105	1,456	1,461	1,448	0.55	0.88	0.157	0.235
Personal Earnings	1,677	1,691	1,688	0.60	0.17	0.155	0.104	1,723	1,733	1,716	0.42	0.97	0.132	0.248
Family Income	2,779	2,781	2,782	0.11	0.01	0.102	0.100	2,984 <sup>b</sup>	2,982 <sup>b</sup>	2,903	2.77	2.71	0.819	0.796
Family Earnings	2,853	2,847	2,845	0.27	0.06	0.108	0.100	2,991 <sup>c</sup>	2,985	2,919	2.45	2.24	0.679	0.589
HH Income	2,858	2,865	2,868	0.34	0.11	0.116	0.102	3,043	3,043	3,001	1.41	1.38	0.390	0.374
HH Earnings	2,917	2,914	2,915	0.07	0.01	0.101	0.100	3,043	3,039	2,997	1.55	1.41	0.400	0.338
Self Employment Income	2,171	2,231	2,210	1.77	0.93	0.121	0.105	2,026	2,061	2,088	2.99	1.33	0.204	0.119
Wages	1,597	1,603	1,603	0.33	0.03	0.121	0.100	1,656	1,661	1,646	0.65	0.95	0.182	0.259
Unemployment	495	497	493	0.36	0.75	0.102	0.109	519 <sup>b</sup>	522 <sup>b</sup>	462	12.42	13.05	0.835	0.864
Welfare	287	292	291	1.41	0.33	0.136	0.102	302	303	303	1.66	1.69	0.102	0.159
Food Stamps	121	123	124	1.94	0.34	0.156	0.102	130	133	128	1.66	3.78	0.132	0.255
Social Security	453 <sup>c</sup>	454	456	0.59	0.32	0.178	0.125	489	489	493	0.92	0.78	0.286	0.238
Federal SSI	234	241	238	1.75	0.99	0.139	0.113	266 <sup>b</sup>	272 <sup>a</sup>	240	11.14	13.43	0.883	0.967
AFDC	357	361	363	1.71	0.64	0.128	0.104	353	357	363	2.63	1.48	0.149	0.116
Median Statistic --->				0.60	0.25	0.125	0.103				1.61	1.45	0.245	0.257
Proportion With Smaller Errors <sup>3</sup> --->				7.1%	92.9%	7.1%	92.9%				50.0%	50.0%	50.0%	50.0%
<b>PROPORTION RECEIVING INCOME FROM SOURCE</b>														
Personal Income	70.4 <sup>a</sup>	70.2	70.0	0.62	0.24	0.315	0.132	73.3	73.2	72.9	0.55	0.34	0.276	0.164
Personal Earnings	45.5 <sup>b</sup>	45.3	45.1	0.99	0.45	0.270	0.135	48.2	48.0	48.1	0.29	0.18	0.119	0.107
Family Income	99.1 <sup>a</sup>	99.1 <sup>a</sup>	98.8	0.26	0.23	0.881	0.758	98.2 <sup>a</sup>	98.1 <sup>a</sup>	98.9	0.69	0.75	0.997	0.998
Family Earnings	79.6	79.7	79.4	0.25	0.30	0.126	0.139	80.1	80.1	80.6	0.59	0.61	0.247	0.254
HH Income	99.5 <sup>a</sup>	99.5 <sup>a</sup>	99.3	0.19	0.19	0.807	0.791	98.9 <sup>a</sup>	98.9 <sup>a</sup>	99.4	0.53	0.55	0.987	0.987
HH Earnings	80.6	80.7	80.5	0.13	0.25	0.108	0.129	80.7 <sup>c</sup>	80.6 <sup>c</sup>	81.8	1.35	1.38	0.691	0.708
Self Employment Income	4.7	4.8	4.8	1.66	0.76	0.131	0.106	5.1	5.2	5.1	0.50	2.75	0.104	0.199
Wages	41.4 <sup>a</sup>	41.1	40.9	1.33	0.43	0.345	0.126	43.9	43.6	43.6	0.62	0.15	0.173	0.104
Unemployment	1.3	1.3	1.3	1.74	1.29	0.113	0.108	1.0	1.0	0.9	6.12	7.29	0.185	0.211
Welfare	4.6 <sup>b</sup>	4.4	4.3	4.96	0.86	0.256	0.105	4.4	4.3	4.4	1.88	1.30	0.125	0.113
Food Stamps	2.8 <sup>a</sup>	2.7	2.7	6.09	0.28	0.310	0.100	2.6	2.5	2.6	1.29	2.70	0.109	0.142
Social Security	14.2	14.1	14.1	0.31	0.21	0.104	0.102	14.9	14.9	14.4	3.62	3.49	0.491	0.472
Federal SSI	1.6 <sup>b</sup>	1.5	1.5	7.09	1.76	0.216	0.108	1.7	1.6	1.6	4.02	0.52	0.146	0.101
AFDC	1.3 <sup>c</sup>	1.2	1.2	6.34	3.96	0.214	0.143	1.1 <sup>c</sup>	1.1 <sup>c</sup>	1.3	15.23	15.57	0.698	0.684
Median Statistic --->				1.16	0.37	0.236	0.128				0.99	1.03	0.216	0.205
Proportion With Smaller Errors <sup>3</sup> --->				14.3%	78.6%	14.3%	85.7%				57.1%	42.9%	42.9%	50.0%

Note: 1. To account for differences in the target population between the 1987 and 1989 SIPPs, 1989 SIPP respondents who were 0-2 years-old in wave one were omitted from this analysis since they would not have been born at wave one of the 1987 SIPP.  
 2. Letters Beside Estimates Indicate Statistical Significance Between Estimate and Benchmark Estimate: 'a' indicates significant at .01 level, 'b' indicates .05 significance, 'c' indicates .10 significance.  
 3. Refers to the proportion of statistics in the group that are smaller when comparing the RTI-Revised statistics with the Current Statistics. For example, 92.9% of the RTI-Revised, % Absolute Relative Bias estimates under the "MEAN \$ AMOUNT AMONG RECIPIENTS" group are smaller than the Current estimates.