MEASURING ELIGIBILITY AND PARTICIPATION IN THE HOUSING ASSISTANCE SUPPLY EXPERIMENT

GRACE M. CARTER
STEVEN L. BALCH

R-2780-HUD
SEPTEMBER 1981

HOUSING ASSISTANCE SUPPLY EXPERIMENT

Sponsored by
The Office of Policy Development and Research
U.S. Department of Housing and Urban Development

Rand
SANTA MONICA, CA. 90406

The research reported here was performed pursuant to Contract No. H-1789 with the Office of Policy Development and Research, U.S. Department of Housing and Urban Development.

Published by The Rand Corporation

MEASURING ELIGIBILITY AND PARTICIPATION IN THE HOUSING ASSISTANCE SUPPLY EXPERIMENT

GRACE M. CARTER
STEVEN L. BALCH

R-2780-HUD
SEPTEMBER 1981

HOUSING ASSISTANCE SUPPLY EXPERIMENT

Sponsored by

The Office of Policy Development and Research
U.S. Department of Housing and Urban Development

Rand
SANTA MONICA, CA. 90406

PREFACE

This report was prepared for the Office of Policy Development and Research, U.S. Department of Housing and Urban Development (HUD), as part of Rand's research on eligibility and participation in the Housing Assistance Supply Experiment, conducted in Brown County, Wisconsin, and St. Joseph County, Indiana. It provides estimates of the population eligible for a housing allowance during the first three program years, the rates of participation in the program, and the methodology used to obtain those estimates.

This report is one of several analyzing eligibility and participation; the others are C. Peter Rydell, John E. Mulford, and Lawrence Kozimor, *Dynamics of Participation in a Housing Allowance Program*, N-1137-HUD; Sinclair Coleman, *How Enrollees Respond to Allowances and Housing Standards*, R-2781-HUD; Phyllis Ellickson, *Who Applies for Housing Allowances? Early Lessons from the Housing Assistance Supply Experiment*, R-2632-HUD; and James Wendt, *Why Households Apply for Housing Allowances*, R-2782-HUD. Findings from these reports will be integrated into a single model of participation process in the final report by Grace Carter and James Wendt, *Participation in an Open Enrollment Housing Allowance Program: Evidence from the Housing Assistance Supply Experiment*, R-2783-HUD.

Comments by I. S. Lowry, C. Lance Barnett, Robert Bell, and Allan Abrahamse on earlier drafts of this report were most useful. Daniel A. Relles suggested using a simulation to evaluate the variance of the eligibility estimates. Helen Wagner performed all the programming to retrieve data on enrollees and constructed and maintained the files used for analysis. Jacqueline Bowens typed several drafts of the report, and Dolores Davis typed the production version. Penny Post edited the report and supervised its production.

The report was prepared pursuant to HUD Contract H-1789, Task 2.16.6.

SUMMARY

The Housing Assistance Supply Experiment offered monthly cash payments to low-income households in two metropolitan housing markets to help them afford decent, safe, and sanitary housing. This report estimates the size and composition of the population of households in each community that were eligible to receive such allowances; analyzes changes in that population over time, including flows of individual households into and out of eligibility; and combines the eligibility estimates with counts of enrollees and recipients to estimate rates of participation in the program for various types of households.

Our data on the eligible population come from household surveys administered annually to occupants of a stratified probability sample of housing units. However, the sample was designed to measure the effects of an allowance program on housing supply, whereas we are particularly interested in how well the program is helping certain population subgroups especially likely to be in need of assistance, such as minorities and the elderly. Because the stratification variables are only weakly related to those subgroups, and sampling proportions vary greatly by stratum, the weighted counts of eligible households within the subgroups have extremely large sampling variances.

Instead of such weighted counts, we chose a two-stage estimation procedure that exploits the strong correlation of eligibility with life-cycle stage and minority status. First, we used a maximum-likelihood logit model to estimate the probability that a household is eligible, given its sampling stratum, life-cycle stage, and minority status; the demographic variables contributed greatly to the predictive power of the model. Second, we estimated the distribution of subgroups within the population of each stratum by an empirical Bayes procedure based on the Dirichlet distribution. The combined procedures estimated the number of eligible households in each subgroup with much lower variance than did the weighted counts. For example, the standard deviations of the estimates of eligibles by life-cycle stage were (at the median) 40 percent lower than those of weighted counts. The estimate of the total number of eligibles also showed modest improvement. Our methodology

may be useful to others seeking to estimate population subgroups from an inconveniently stratified sample.

Roughly one-fifth of the population could not afford the standard cost of adequate housing without spending more than one-fourth of their income, and at the same time meet the other eligibility requirements of the allowance program. Two life-cycle groups, single parents and elderly singles, were heavily overrepresented in the eligible pool. In one of the two sites, for instance, the eligibility rate of these two groups exceeded 60 percent. The same site had a substantial racial minority, whose members were twice as likely as non-Hispanic whites to be eligible.

There were slight changes in the size and composition of the eligible population over time, due to demographic change in the population at large, modifications in Social Security payments, and unemployment shifts. However, there were substantial changes in the identity of eligible households. From one year to another less than three-quarters of the eligible households retained their eligibility and continued to be headed by the same person(s). Of the newly eligible households, about one-third were newly formed by marriage or separation.

During the program's first two years the number of enrolled households grew rapidly, but by the third year it reached a steady state at about 40 percent of those eligible. To qualify for payments, an enrollee had to occupy adequate housing. Both enrollment and recipiency rates were highest among the subgroups that also had the highest eligibility rates: renters, single parents, and elderly singles. Minorities were much more likely to enroll than nonminorities, but, probably due to housing problems, those who actually received allowances were only a slightly higher proportion of eligibles (35 percent at the end of the third program year) than were whites (29 percent).

In conclusion, we review the differences between our findings and those of other housing allowance participation studies. Comparison with results from the Housing Allowance Demand Experiment, for instance, suggests that a program that invites individual households to enroll may acquire a demographic mix different from that of an open enrollment program, even if those invited to enroll are a representative sample of eligibles.

CONTENTS

TABLES

FIGURES

I. INTRODUCTION

The Housing Assistance Supply Experiment was undertaken as part of a larger research program to help HUD and the Congress decide whether direct cash assistance to low-income households is a feasible and desirable way to improve their housing and relieve their housing expense burdens. In this experiment, monthly housing allowances were offered to nearly all low-income households in two metropolitan housing markets, on condition that the recipients find adequate housing on the private market and maintain it to program standards. The eligibility standards were similar to those of the long-established public housing program which, however, has never accommodated more than a small percentage of those eligible.

This report deals with three questions about such an open-enrollment program:

- Given the eligibility standards, what is the size and composition of the eligible population?
- Among the eligibles, who chooses to enroll?
- Among the enrollees, who qualifies for payments?

The answers to these questions bear on both program cost and effectiveness. The size and composition of the eligible population determine the potential fiscal liability of such a program. The size and composition of the enrolled population and its turnover determine the administrative cost of the program. The size and composition of the recipient population determine the allowance cost of the program and also its effectiveness in reaching those for whom--according to the eligibility standards--help was intended.* Evidence from the two experimental sites greatly reduces the uncertainties inherent in estimating the cost and effectiveness of a national program.

---

*Other aspects of program effectiveness include how the participants change their housing consumption and how household budgets are affected; for analysis of these, see Mulford et al. (1981). This report deals only with eligibility, enrollment, and receipt of payments.

Although eligibility and participation estimates can be reduced to gross counts of households or dollars of allowance entitlement, program evaluation also requires compositional detail. The program's eligibility standards specify income and asset limits, the former varying with household size. What kinds of households meet those standards is a matter of considerable interest to policymakers. In this report, we distinguish groups that because of their nonfinancial characteristics are of special interest: renters and homeowners, whites and racial minorities, and various family configurations that roughly reflect stages in a household's life cycle. These groups are defined in Table 1.1.

The data collected in the two HASE sites (Brown County, Wisconsin, whose central city is Green Bay, and St. Joseph County, Indiana, whose central city is South Bend) offer unusual opportunities for analyzing eligibility and participation. Annual household surveys provide detailed sample representation of each community's population, including nearly all the household characteristics that bear on eligibility. Consequently, we are able to estimate the size and characteristics of the eligible population with unusual precision. Allowance program records provide corresponding detail on those who enrolled and those who received payments, so we are able to calculate participation rates for various groups of interest.

Moreover, our data are longitudinal as well as cross-sectional. Our household surveys were repeated annually, collecting information from the same housing units each time but not necessarily the same occupants. However, the large proportion of overlap among households from year to year made it possible to trace the movements of many of them into and out of eligibility. Even when a household was only interviewed once, its record includes retrospective data that can be used to estimate changes in eligibility status. Combining survey and program data, we can estimate turnover in both the eligible and enrolled populations.

Rates of turnover are important for several reasons. For one, the administrative work load of the housing allowance offices (HAOs) depends not only on the number of clients in the program at any one time, but on how many new clients join the program each year (see Kingsley

TABLE 1.1

POPULATION SUBGROUPS

Subgroup	Definition
Renters	All renters who pay full or partial rent (including lodgers), and those who live rent-free.
Owners	All who own their homes, occupants of mobile home parks, and, in St. Joseph County, occupants of cooperatives and condominiums.*
Minorities	All households whose head is black, Hispanic, Oriental, or native American.**
Elderly single	Aged 62 or older.
Single parent	Household headed by single man or woman under 62, with at least one member under 18.
Other singles	All persons not living with a spouse or other family member, and under 62; were excluded from the allowance program and therefore assumed to be ineligible.***
Young couple with young children	Head of household aged 45 or younger; at least one child under age 6.
Elderly couple	Household whose head is 62 or older.
Other couples	Both heads under 62, and either no children under 6 or head of household over 45.†

SOURCE: Classification scheme devised by HASE staff for analysis of data from the household surveys.

*Some occupants of mobile homes rent the land on which the homes are located, and a few also rent the homes themselves.

**Brown County has virtually no minority households. The minority population in St. Joseph County is predominantly black.

***The data for this report were collected prior to August 1977; only non-elderly singles who were disabled, handicapped, or displaced were eligil before that date, and the surveys did not identify those circumstances.

†The designation of "Other couples" includes a few (less than 2 percent) three-generational families.

and Schlegel, 1981); most new clients in a mature program are newly eligible households. Second, turnover affects participation rates since most newly eligible households do not apply at once for a housing allowance, but await some stimulus that informs or reminds them that the allowance program is available. Even after they have considered the program and decided to apply, it takes some time to complete the enrollment process. As a result, enrollment rates never reach 100 percent.[*] Third, the length of time a household expects to remain eligible influences the allowance's attractiveness--that is, whether or not the benefits will be worth the effort of applying. Because duration of eligibility is also an indicator of need, lack of participation by the briefly poor may not detract from the program. In related studies, HASE analysts use a cost/benefit framework to explain households' decisions to participate or not, which in turn clarifies the different participation rates observed in this study. (See Coleman, 1981; Wendt, 1981; Carter and Wendt, 1981.)

Much of this document is devoted to the methods we use to estimate the number of eligible households within the general population. Our procedure should be of interest to others working with stratified samples,[**] in view of the increasing practice of sharing data bases.[***] The techniques we employ will be most valuable when a stratified sample collected for one purpose is used for a different estimation problem where the relationship between the stratification variables and the characteristic of interest is not very strong.

Our problem is that while the population characteristics of concern to us (see Table 1.1) are strongly correlated with eligibility, they are only weakly correlated with the variables used for stratifi-

---

[*] Rydell, Mulford, and Kozimor (1981) present a Markov model of the participation process.

[**] A stratified sample is one in which the population is divided into strata or subsets prior to sample selection. Different proportions of each stratum are usually included in the sample.

[***] The HASE household surveys will be available to the research community through HUD. Other examples of shared stratified data bases are the Michigan Panel on Income Dynamics and the Bureau of the Census' Annual Housing Survey.

cation; and sampling rates vary greatly among strata. Weighted counts* of eligible population subgroups are therefore likely to contain a large component of random error. However, eligibility is also somewhat related to the stratification variables. Consequently we have separated the estimation problems into two parts.

The probability that a household is eligible given both its characteristics and sampling stratum is estimated by a maximum likelihood logit model. The distribution of household subgroups within the population of each stratum is estimated by an empirical Bayes procedure based on the Dirichlet distribution. The results are then combined to estimate the number of eligibles in each population subgroup and in total. Our techniques produce a more accurate set of estimates, as demonstrated in the smaller standard deviations presented in Sec. II.

The rest of this introduction will describe the data sources we draw upon, the experimental housing allowance program, and definitions of eligibility. All discussion of statistical methods is in Secs. II and III; readers not interested in those details can proceed directly to Sec. IV, which contains the findings about the eligible population and its participation.

Section II explains the statistical techniques used for estimating eligibility. First, we present estimates of eligibility obtained from weighted counts of survey records, to show why a more efficient approach is desirable. Then we explain the logit estimator of eligibility and the procedure for estimating the distribution of household types; and compare those estimates with the ones from weighted counts. (The stochastic simulation from which the standard deviation of our estimates was computed is described in Appendix A.) Section III contains our method for estimating turnover within the eligible pool. For many households, these estimates are based on retrospective data. However, other households were interviewed repeatedly. Consequently we were able to correct biases in the retrospective data by using evidence ob-

---

*Weighted counts are sums of sample observations, with the weight on each observation from stratum i equal to N_i/n_i, where N_i is the population of stratum i and n_i is the sample size in stratum i.

tained from households for whom we have actual eligibility status in two or more successive years.

Section IV begins by characterizing the eligible population as of 1977 and the changes in that pool since the start of the program. It then identifies household flows into and out of eligibility for each type of household mentioned above. Participation rates are measured both over time and as of the end of the third year in each site, the approximate date of the last household survey available when the analysis was done. (Estimates of the fourth program year's participation rates, based on the assumption that the eligible pool did not change in either size or composition over that year, are provided in Appendix B.)

DATA SOURCES

Our description of the eligible population comes from annual surveys of the occupants of a stratified random sample of residential properties at each site.* Since the surveys are keyed to tax parcels rather than households, the same housing units were surveyed each year whether or not the household was the same. Each sample was divided into 18 strata, and newly constructed properties were added to both panels each year. Although stratum sizes varied, sampling histories allowed us to infer population parameters. (See Corcoran, 1981a; 1981b.)

For each step of estimation and analysis, we have used the maximum number of survey records available and suitable for the task at hand. For instance, our estimate of the distribution of population subgroups is based on all records containing complete household descriptions, while our estimates of eligibility are restricted to the slightly smaller number of households who also provided complete income information. Analysis of eligibility changes among households was further restricted to those who supplied complete and internally consistent employment histories. A comparison of repeatedly surveyed households

---

*Household surveys are documented in a series of codebooks for each site; for example, Boren (1980). For a complete overview of the design of HASE, including the household surveys and other surveys not directly used here, see Lowry (1980).

with both missing and complete employment histories revealed no evidence that the exclusions biased the turnover estimates.

Our data on participants in the housing allowance program come from administrative records maintained by the housing allowance office in each site.* Similarities of household composition and other characteristics recorded by both the HAOs and the household surveys facilitate comparison between HAO clients and the general population.

We use data from the household surveys for all four years in Brown County beginning in 1974, and for the first three years in St. Joseph County from 1975; the surveys were conducted mainly in the first half of each year. The HAO records also included four years of operation, beginning in June 1974 for Brown County and January 1975 for St. Joseph County.** Thus our survey data cover the time from before the program began through about three years of operation in Brown County and a little more than two years in St. Joseph County. (A fourth survey wave in St. Joseph County was not ready to be included in this analysis.)

THE HOUSING ALLOWANCE PROGRAM

The allowance program is open to nearly all families*** in both Brown and St. Joseph counties who cannot afford the standard cost of adequate housing in the local market without spending more than a quarter of their adjusted gross income. Each participating household receives monthly cash payments equal to the difference between a quarter of its income and the standard cost, provided that its housing meets minimum standards of decency, safety, sanitation, and space. (See McDowell, 1979; 1980.)

Whether they are renters or homeowners, participants negotiate the terms of lease or purchase in the open market. They are entirely re-

---

*We used the Client Characteristics File, documented through program year 3 by Wang and Geller (1979).

**The program in St. Joseph County began operation with a period of limited enrollment. Open enrollment began only in April 1975. Thus the end of the fourth program year really corresponds to only 3.75 years of open enrollment.

***The program is now also open to most single adults. They were excluded until August 1977, which is later than our survey information.

sponsible for meeting their obligations to landlords, lenders, or other parties to their housing transactions. As long as their housing continues to pass periodic inspection by the program's officials, they may change tenure or move anywhere within the program's jurisdiction without loss of benefits.

Each household that enrolls is informed of the program's housing requirements and the amount of allowance it is entitled to receive. If its current dwelling fails the initial inspection, the household is informed of the reasons and must then either make the necessary repairs or move to an acceptable dwelling. There is no time limit for action, but benefits are not provided until the housing requirements are met.

DEFINITIONS

A household is eligible to enroll in the allowance program if it meets certain asset, income, and family composition standards. The asset limit in use at both sites during all of the first three years was \$32,500 for an elderly household and \$20,000 for a nonelderly household. Until August 1977, persons under the age of 62 who lived alone or with nonrelatives could enroll only if they were handicapped, disabled, or displaced by a government program. Since the surveys did not report handicaps, disabilities, or displacement, we treated all nonelderly singles as ineligibles. For consistency we also eliminated them from counts of HAO clients.

In order for a household to enroll in the program, its annual allowance entitlement must equal or exceed \$120. The allowance entitlement, B, is calculated from R^*, the standard cost of adequate housing, and Y, adjusted income, as

$$B = R^* - .25Y .$$

R^* is a function of household size. Adjusted income is derived from gross income by subtracting a standard deduction of 5 percent for nonelderly households and 10 percent for elderly ones, and other deductions such as that for dependents (\$300 each) and occupational or medical expenses.

The adjusted income for each household has been calculated as closely as possible to the figure the HAO would have reached if the head of the household had gone to the HAO office the day of the survey. Earnings from employment reflect the then current job and wage rates. Other income is based on the rates* of the preceding year. The survey data enable us to calculate the amount of all deductions except medical and occupational expenses, which we assume to be zero.

The R^* values used by the HAOs were first measured during the screener surveys that preceded baseline in both sites. They are periodically updated to reflect measured inflation in the housing market. However, considerable time elapsed before the first such adjustment in each site. For example, in Brown County the first values of R^*, measured in September 1973, were used by the HAO from the beginning of the program until April 1976 when R^* was increased to the values measured in January 1976. (See Lowry, 1979.)

The values we use in estimating eligibility and participation are slightly different from those used by the HAOs because instead of basing our estimates on the periodic changes in R^*, we assume that its values rose linearly between each pair of measurements. We therefore interpolated measurements at the median point in each survey. Using these interpolated values of R^* (reported in Appendix C) enabled us to estimate how many households lacked enough income to meet the current cost of adequate housing at the time of each survey, even though not all would have been actually eligible to enroll then.

Thus we estimate which households need the allowance program rather than which households would have been allowed to enroll. However, in Sec. IV we show that the alternative estimates were fairly close after the program matured.

---

*If the household heads are either retired or unemployed, retirement income and unemployment insurance from the preceding year were adjusted to annual rates by dividing those amounts by the fraction of that year that the household heads spent retired or unemployed.

II. ESTIMATING ELIGIBILITY FROM A STRATIFIED SAMPLE

Each year we received complete survey responses from about 4.8 percent of Brown County households and 2.4 percent of St. Joseph County households (averaged over all waves).* The surveys told us the type of each household, which sampling stratum it belonged to, and whether it was eligible.

We wanted to estimate the number of eligible households at each survey wave by housing tenure, life-cycle stage, and minority status. The stratification and large size of the sample in each site suggests that weighted sample counts** should be accurate enough to measure such simple quantities as the number of eligible households in each population subgroup. However, the stratification was designed to meet the primary research objective of HASE, i.e., to measure housing supply response to an allowance program. That design is not very good for making eligibility estimates; in fact, it produces more random variance in our estimates than would a sample of the same size drawn at random.

We solved the problem in two stages. First we estimated a logit regression for P_{ikt}, the probability that a household of life-cycle stage and minority status k and occupying a unit in panel stratum i is eligible at wave t.*** Because household type proved to be strongly associated with eligibility in our model even after we controlled for stratum, it improves our predictions.†

In the second stage we estimate θ_{ikt}, the proportion of households in stratum i at wave t who are of household type k. Because household type is only weakly related to stratum, we use an empirical Bayes technique to pool the data across strata to estimate the distribution of household type.

---

*The average sample sizes per wave are 2,249 in Brown County and 1,701 in St. Joseph County.

**This procedure is described under the next heading, Weighted Count Estimates of Eligibility.

***Tenure is completely defined by panel stratum.

†Household type functions as an auxiliary variable in our estimate of P_{ikt}. See, for example, Morris and Rolph (1978), pp. 320-322.

The number of eligibles of type k at wave t, v_{kt}, is then estimated as:

$$v_{kt} = \sum_i N_{it}\theta_{ikt}P_{ikt} \tag{1}$$

where N_{it} is the number of households in stratum i at wave t. Confidence limits on v_{kt} are calculated by a stochastic simulation.

To show why such an elaborate procedure is necessary, we first present weighted counts from survey records, then the estimates of P_{ikt} and θ_{ikt}. We show that the second method reduces the standard deviation of the estimate of eligibles in each subgroup.

WEIGHTED COUNT ESTIMATES OF ELIGIBILITY

Table 2.1 shows the number of households in each stratum, and the percentage who provided enough information to determine eligibility, during the household survey for wave 4 in Brown County and wave 3 in St. Joseph County. The sampling fractions for the other waves are quite similar.* The residential properties in each county were divided into 18 strata at baseline. They consisted of the major categories of housing tenure and rural/urban location, subdivided according to rent and assessed value. Urban rental properties were further subdivided by the number of residential units on them. A new construction sample was added at each wave and we treat them as a single stratum for each tenure group. Mobile homes and lodger units (including additions) are in separate strata.

The number of certain kinds of households--eligible and elderly, for instance--is estimated from a stratified sample by weighted counts from each stratum as

$$\hat{X} = \sum_i \frac{N_i}{n_i} X_i \, , \tag{2}$$

---

*The Brown County sampling fractions at waves 1, 2, and 3 were 0.052, 0.050, and 0.045 with sample sizes of 2,357, 2,327, and 2,129. In St. Joseph County the sampling fractions at waves 1 and 2 were 0.024 and 0.023, corresponding to sample sizes of 1,741 and 1,626. A larger baseline sample was available but was not used, in order to avoid artificial changes in eligibility estimates due to sample changes.

Table 2.1

1977 POPULATIONS AND SAMPLE SIZES BY STRATUM

Stratum	Brown County			St. Joseph County		
	Number of Households	Sample Size	Percent Sampled	Number of Households	Sample Size	Percent Sampled
Urban Renter						
Single-family:						
Low rent	331	74	22.4	1,184	56	4.7
Medium rent	524	121	23.1	2,226	150	6.7
High rent	624	43	6.9	1,787	91	5.1
2-4 units:						
Low rent	2,507	234	9.3	2,709	235	8.7
Medium rent	2,487	320	12.9	1,710	212	12.4
High rent	2,095	91	4.3	663	62	9.4
5+ units:						
Low rent	403	72	17.9	718	92	12.8
Medium rent	1,579	311	19.7	663	58	8.7
High rent	1,424	87	6.1	2,703	95	3.5
Rural Renter						
Low or medium rent	644	127	19.7	592	84	14.2
High rent	247	37	15.0	227	31	13.7
New Construction	1,335	101	7.6	244	20	8.2
Urban Owner						
Low value	5,296	121	2.3	9,834	112	1.1
Medium value	6,931	163	2.4	12,611	124	1.0
High value	13,257	70	0.5	25,444	59	0.2
Rural Owner						
Low value	2,867	70	2.4	1,990	76	3.8
High value	2,808	28	1.0	4,024	37	0.9
New Construction	1,677	52	3.1	1,003	80	8.0
Other Residential						
Rooming house	142	16	11.3	193	9	4.7
Mobile home property	894	42	4.7	1,361	55	4.0
All strata	48,072	2,182	4.5	71,886	1,738	2.4

SOURCE: Tabulated by HASE staff from HAMISH and the household surveys for wave 4 in Brown County and wave 3 for St. Joseph County.

NOTE: Population estimates and sample counts exclude resident landlords, because they were not surveyed, and households occupying subsidized dwellings, because they were excluded from the eligibility estimates presented in this report.

Sample sizes refer only to records providing complete data on income and household characteristics.

where N_i = the number of households in stratum i,

n_i = the number of households sampled in stratum i, and

X_i = the number of households in this sample with the characteristic of interest; in this example, who are eligible and elderly.

We shall use the carat (^) to denote estimates throughout the paper. The variance of $\hat{X}$ is:

$$\mathrm{Var}\ (\hat{X}) = \sum_i N_i(N_i - n_i)P_i(1 - P_i) \tag{3}$$

where P_i = the proportion of stratum i who have the characteristic(s) of interest.

How efficiently Eq. (2) estimates X depends on two elements of the original sampling design. First, do the stratum subpopulations vary with respect to the characteristics of interest? Several stratification variables such as tenure and amount of rent or assessed value are usefully correlated with eligibility; but they do little to differentiate such population subgroups as the elderly and minorities. Second, how is the total sample (n) allocated among the strata? Ideally, n_i would be proportional to population size and also to standard deviation, in which case the ratio of sample sizes in each pair of strata i and j would be expressed:

$$\frac{n_i}{n_j} = \frac{N_i\ \sqrt{P_i(1 - P_i)}}{N_j\ \sqrt{P_j(1 - P_j)}}$$

But the proportion of households sampled within each stratum varies much more than is convenient for making eligibility estimates. The most outstanding example is renters, who were heavily oversampled relative to homeowners.* So our total sample is especially unsatisfactory in this second respect.

---

*Three times as many renters as homeowners were sampled in Brown County, whereas our posterior estimates of eligibility rates indicate that the optimal sample would have been 6 renters for every 10 homeowners. The problem was similar in St. Joseph County, where the sample

Table 2.2 shows the weighted count estimates of St. Joseph County households eligible in 1977 by tenure, life-cycle stage and minority status. The total number of eligible households is estimated at 14,379, with a 95 percent confidence limit of plus or minus about 2,425. Large as that interval looks, it is sufficiently precise to calculate overall participation rates. For instance, a participation rate of 0.5 based on that estimate would allow us 95 percent confidence that the true number fell between 43 and 60 percent.

Even though weighted counts provide sufficiently accurate estimates of total number of eligibles, they are too imprecise for measurement of population subgroups. It is easy to compare the relative precision of subgroup estimates by using the coefficient of variation (the standard deviation of the estimate, calculated from Eq. (3)* and expressed as a percent of the estimate). Because of the larger proportion of renters sampled, estimates for them are more accurate than the ones for homeowners. However, for estimates of subgroups defined by life-cycle stage or minority status the coefficient of variation ranges from 15 to 32, which is quite large.

One manifestation of the large variance is that estimates of the number of households eligible within certain population subgroups varied widely from year to year (See Table 2.2). From one survey wave to another, for instance, the estimated number of eligible young couples with young children varied by almost 50 percent; yet that was not even a statistically significant difference.** Another manifestation of the variance is that estimates of participation rates are very imprecise. For example, among elderly couples, if we estimated a participation rate of 40 percent, the limits of a 95 percent confidence interval

---

included 26 renters for every 10 homeowners, compared to an optimal ratio of 3 renters per 10 homeowners. Though sampling proportions within tenure groups did decline as rent or assessed value increased, those ratios are also far from optimal.

*P_i was estimated at the sample proportion in each stratum. When no cases in a group were interviewed, P_i in Eq. (3) was set to $1/n_i$.

**The differences are not statistically significant even accounting for the difference estimate's lower variance resulting from repeated surveys of the same housing units.

Table 2.2

WEIGHTED COUNT ESTIMATES OF ELIGIBLE HOUSEHOLDS IN ST. JOSEPH COUNTY, BY SELECTED HOUSEHOLD CHARACTERISTICS: 1977

Household Type	Number of Households Eligible	Standard Deviation of Estimate	Coefficient of Variation for Estimate	Annual Change in Eligible Households (%)	
				1975-76	1976-77
Tenure:					
Renter	4,739	214	5	2	9
Owner	9,640	1,219	13	-24	13
Life-cycle stage:					
Elderly couple	2,479	698	28	-41	36
Elderly single	4,639	682	15	-14	1
Single parent	3,874	692	18	25	19
Young couple, young children	1,814	521	29	-48	50
Other	1,573	508	32	-9	-23
Race of head:					
White non-Hispanic	10,614	994	9	-18	0
Minority	3,765	815	22	-12	61
Total	14,379	1,238	9	-17	11

SOURCE: Tabulated by authors from the household surveys for waves 1-3.

NOTE: Households living in government-assisted properties and resident landlords are excluded. Nonelderly single-person households are excluded because almost all were ineligible until August 1977.

would be 26 and 89 percent. If we knew only that somewhere between a quarter and nine-tenths of eligible elderly couples were participating, it would be a rather poor measure of how well the allowance program was serving them.

The Brown County data presented in Table 2.3 is similar to that for St. Joseph County. Because we sampled a larger percentage of the population in Brown County, our estimates there are slightly more accurate; but measurement of subgroups among the eligible population is still poor.

Below we describe in detail the model we constructed to estimate population subgroups more accurately. From the new counts we were much better able to target our estimates of participation, and thus provide the basis for a more precise assessment of the outcome of the allowance program.

LOGIT MODEL OF ELIGIBILITY

Among the households in the sampled dwellings, eligibility and life-cycle stage showed a strong correlation. For instance, over half the elderly single households in the sample, but less than 10 percent of nonelderly couples, were eligible at each wave. This correlation was useful for improving estimates of eligibility. The logit model described in this section estimates the conditional probability, P_{ikt}, that a household is eligible given its panel stratum i, life-cycle stage (and minority status in St. Joseph County) k, and wave t. If Z_{ikt} is a vector of dummy variables describing a particular household, and b is a vector of coefficients, then

$$P_{ikt} = \frac{exp(bZ_{ikt})}{1 + exp(bZ_{ikt})} \tag{4}$$

Sample Selection

A logit model rests on the assumption that each datum point is an independent sample from a binomial process. But the eligibility states of the same household surveyed at two different waves are not independent; so each surveyed household provides only one datum point for the model.

Table 2.3

WEIGHTED COUNT ESTIMATES OF ELIGIBLE HOUSEHOLDS IN BROWN COUNTY, BY TENURE AND LIFE-CYCLE STAGE: 1977

Household Type	Number of Households Eligible	Standard Deviation of Estimate	Coefficient of Variation for Estimate	Annual Increase in Eligible Households (%)		
				1974-75	1975-76	1976-77
Tenure:						
Renter	3,615	162	4	12	5	-4
Owner	4,007	480	12	14	-7	-3
Life-cycle stage:						
Elderly couple	1,108	276	25	8	35	-30
Elderly single	2,016	293	15	39	-12	-6
Single parent	2,116	309	15	28	12	5
Young couple, young children	1,569	280	18	-8	-22	32
Other	813	242	30	-10	-10	-14
Total	7,622	507	7	13	-2	-3

SOURCE: Tabulated by authors from the household surveys for waves 1-4.

NOTE: Households living in government-assisted properties and resident landlords are excluded. Non-elderly single person households are assumed to be ineligible because almost all were categorically excluded from participation until August 1977.

Households interviewed at more than one wave were assigned to a single wave by a random procedure that controlled for household characteristics as follows. Households interviewed at two waves and no more were first grouped according to their eligibility state at each wave, panel stratum, life-cycle stage, and race in St. Joseph County. Whenever all of these variables were the same for any two households, one was assigned to each wave. The remaining households who had been interviewed twice were then sorted into groups defined by eligibility state at each wave. Half of each group was randomly assigned to each wave. The procedure for those interviewed at three or four waves was similar except that three or four households were necessary for systematic rather than random assignment. Over half of those interviewed at more than one wave were assigned to a wave by the systematic rather than random selection procedure.

Nonelderly single-person households were excluded from the sample because most were categorically ineligible until August 1977, and by then the surveys we are using were completed. Brown County occupants of subsidized housing were omitted from the sample* because there are few of them in the population and none were in the survey until wave 3.

The Equation

The independent variables describe the stratum, household type, and wave at which each interview was conducted. Because we also stratified the sample households according to the number of times they were interviewed, another variable representing that number was also included in the equation.** Maximum likelihood was the estimation technique used to fit the model.

Stratum Variables. Dummy variables were used to identify each of the strata shown in Table 2.1; the low-rent tercile of urban properties with 5+ units was the comparison group. Dummy variables were defined for each of the main classes of properties: rural homeowners, urban

---

*St. Joseph County occupants of subsidized housing are included in the sample, but treated separately as described later.

**Interactions of number of waves interviewed with life-cycle stage and tenure were not significant at the 0.1 level.

homeowners, rural renters, and each of the medium- and high-rent terciles of urban renters. Subsidiary variables identified individual strata within these major classes.*

Occupants of government-subsidized housing in St. Joseph County were separated into three extra strata because they were very likely to meet the allowance program's eligibility criteria, but were not able to receive payments while in that housing.** The new strata were defined as 1) subsidized homeowners, 2) subsidized renters initially assigned to the lowest rent tercile stratum, and 3) all other subsidized renters.

Most of the strata have a smaller proportion of eligibles than the comparison category described above, as is clear from the negative coefficients on those variables in Tables 2.4 and 2.5. The proportion of eligibles declines as rent or assessed value increases. Only among Brown County rural homeowners with assessed property value higher than the median is this not true; and the *t*-statistic of 0.26 is so small that this exception is probably due to random error.

Household Type Variables. For the set of variables explaining household type, the comparison category was all nonelderly couples except young couples with young children. Not surprisingly, each of the other groups was more likely to be eligible, even after controlling for stratum. Elderly homeowners were more likely to be eligible than is indicated by the separate effects of life-cycle stage and panel stratum, although in Brown County that finding was only marginally significant.

Minority households in St. Joseph County were more likely than other households to be eligible even after stratum and life-cycle stage were accounted for. The data also suggest that an interaction between elderly and minority groups may exist; i.e., that elderly minority

---

*The small sample size for lodgers in St. Joseph County dictated pooling this group with a low-rent urban group. Single-family dwellings were chosen.

**Their inability to receive payments was the reason for their exclusion in the estimates of eligibles reported in Sec. IV. They were included in the regression to estimate eligibility, however, since those estimates will be used for other purposes as well.

Table 2.4

LOGIT REGRESSION OF ELIGIBILITY ON STRATUM, LIFE-CYCLE STAGE, AND TIME: BROWN COUNTY

Variable	Coefficient	*t*-statistic
Constant	-1.034	3.44
Stratum Effects		
Urban renters:		
Low rent		
Single-family	-.135	.39
2-4 units	-.189	.63
Medium rent	-.597	-1.98
Single-family	.493	2.73
2-4 units	.137	.89
High rent	-1.130	3.31
Single-family	.634	2.17
2-4 units	-.270	.99
New construction	-.787	2.07
Rural renters	-.864	2.70
High rent	-.402	1.22
Rooming houses	.438	.71
Mobile homes	-.962	2.22
Urban owners	-1.206	3.45
Medium value	-.998	3.77
High value	-2.100	4.17
Rural owners	-1.743	4.38
High value	.126	.26
New construction owners	-3.437	4.38
Household Type Effects		
Elderly couple	1.847	9.51
Elderly single	3.095	12.60
Single parent	2.860	21.63
Young couple, young children	.856	8.21
Elderly homeowners	.458	1.78
Time Effects		
Survey wave	.122	3.12
Wave: Elderly single	-.362	3.18
Number of waves surveyed	-.088	1.93

SOURCE: Estimated by the authors from household survey data for waves 1-4.

NOTE: Sample size = 3,725; χ^2 = 1,217 with 27 degrees of freedom. The stratum variables are formulated in terms of main effects and interactions; low-rent urban households in properties with five or more units are the comparison group.

Table 2.5

LOGIT REGRESSION OF ELIGIBILITY ON STRATUM, LIFE-CYCLE STAGE, MINORITY STATUS, AND TIME: ST. JOSEPH COUNTY

Variable	Coefficient	*t*-statistic
Constant	-.434	1.56
Stratum Effects		
Urban renters:		
Low rent		
Single-family	-.875	2.48
2-4 units	-.443	1.56
Medium rent	-.525	1.27
Single-family	-.427	1.15
2-4 units	.137	.37
High rent	-1.909	5.59
Single-family	.975	3.19
2-4 units	1.185	3.41
New construction	-.174	.20
Rural renters	-1.011	3.28
High rent	-.504	1.25
Subsidized renters	.271	.66
Higher rent	-1.160	2.71
Mobile homes	-.287	.50
Urban owners	-1.477	4.62
Medium value	-.031	.11
High value	-1.272	3.24
Rural owners	-1.411	3.84
High value	-1.286	2.51
New construction owners	-3.334	4.96
Subsidized owners	-2.278	4.06
Household Type Fffects		
Elderly couple	1.198	5.11
Elderly single	2.840	9.94
Single parent	2.266	10.43
Young couple, young children	.526	4.01
Minority	.961	6.96
Elderly homeowner	.711	2.75
Elderly minority	-.532	1.30
Time Effects		
Wave: Elderly single	-.161	.98
Wave 2 or 3:		
Nonelderly	-.172	1.29
Single parent	.567	2.14
Number of waves surveyed	-.131	1.75

SOURCE: Estimated by the authors from household survey data for waves 1-3.

NOTE: Sample size = 2,745; χ^2 = 1,139 with 32 degrees of freedom. The stratum variables are formulated in terms of main effects and interactions; low-rent urban households in properties with five or more units are the comparison group.

households were slightly less likely to be eligible than the additive effects of being elderly and minority would indicate. In order to avoid bias in estimates of eligibility by household type, this interaction was kept in the regression, although there is a 20 percent chance the effect may have been random.

Time-Linked Variables. The third set of coefficients describes changes in eligibility over time. Each household is identified according to when it was surveyed. The year of the survey is related to eligibility for three reasons: fluctuating unemployment, increases in Social Security, and the static nature of the stratum assignments relative to the changes in housing stock over time.

As the unemployment rate goes up, a household whose head is in the labor force may move into eligibility; as the rate goes down, the same household may become ineligible. According to the Bureau of Labor Statistics, at the time of our baseline survey in South Bend the unemployment rate there was 8.5 percent;[*] a year later it had dropped to 5.1 percent, and the following year to 4.1 percent.[**] Survey data on heads of households reporting unemployment during that period showed a similar pattern. While corresponding BLS data are not available for Brown County,[***] our surveys indicate less variation in unemployment rates over time there than in St. Joseph County.

Social Security payments increased faster than the standard cost of housing during this period. Between December 1973 and 1976, approximately the same span as the Brown County survey years, Social Security benefits went up an average of 35 percent, while the standard

---

[*] *Area Trends in Employment and Unemployment*, U.S. Department of Labor, April 1975.

[**] *Area Trends in Employment and Unemployment*, U.S. Department of Labor, May, June, July 1977.

[***] However, while total employment in the non-farm sector of the economy (omitting the military, proprietors, self-employed, domestic workers in private homes and unpaid family workers) actually decreased by 4.2 percent between 1974 and 1975 in the South Bend labor area, it continued to grow at a reduced rate in Brown County.

The employment figures in thousands for 1974, 1975, 1976 and 1977 were 105.0, 100.6, 103.4, and 107.6 in South Bend and 64.8, 65.7, 67.9 and 72.5 in Brown County. *Employment and Earnings, States and Areas: 1939-1978.* U.S. Department of Labor, BLS, 1979, Bulletin 1370-13.

cost of housing increased 24 percent for single-person households, and 20 percent for two-person households.*

Units were permanently assigned to each stratum by category of rent or assessed value when first surveyed.** However, as old units were demolished and newly constructed units were added each year, the number of units within each stratum of the population changed. In Brown County, for instance, the original sample of dwellings represented 100 percent of the occupied rental units at baseline, but only 91 percent of them at wave 4; the remaining 9 percent were represented at wave 4 by the new construction sample.

If the proportion of the population who were eligible remained constant over time, and those who moved into newly constructed dwellings were less likely to be eligible, then the proportions of the population in the original panel strata who were eligible would increase slightly over time. Thus, once we controlled for stratum--as in our regression--the probability of eligibility would also increase over time.

Exploratory analysis with a full model accounting for each of the time-related factors*** described above showed that virtually all of their effects could be expressed in the three variables we used in Tables 2.4 and 2.5.

---

*Average monthly benefits for males in December of 1973 to December 1976 in dollars were 183, 207, 228 and 248. For females the corresponding numbers are 146, 165, 182 and 197. (*Statistical Abstract of the United States*, Bureau of the Census, 1978, p. 341) The increase in the cost of housing over our three survey waves in St. Joseph County was 11 percent for one-person households and 12 percent for two-person households while Social Security payments increased by about 20 percent, based on Bureau of Census data. Our survey data indicate a slightly greater increase in Social Security for survey respondents than the national figures.

**Stratum changes were made in response to tenure changes. Sampling history records allow adjustment of sampling proportions in each stratum each year.

***The full model contains:

1. Year; a positive coefficient would account for the changing composition of our strata.
2. Interaction of year with both of the elderly groups; different negative coefficients were expected for the two groups, because Social Security provides a greater proportion of income for elderly singles than for couples.

The Brown County coefficients show that, as expected, the eligibility of elderly singles declined over time, whereas the probability of all other household types being eligible, given their panel stratum, increased over time. The *t*-statistic on both of these coefficients is greater than 3, underscoring the validity of our conclusions.

In St. Joseph County the effects of time are not statistically significant in the aggregate. However, the signs are plausible; the slight decrease in eligibility for nonelderly between baseline and wave 2 corresponded to the large drop in unemployment, and the decrease in eligibility among elderly singles matched the higher Social Security benefits; the proportion of single parents who were eligible went up slightly after controlling for stratum.

The life-cycle variables and their interactions contribute very heavily to the model's ability to predict eligibility. In fact, life-cycle stage explains more of the variance than all the stratum and time variables together. If we predict eligibility for Brown County on the basis of all stratum and time-linked variables,* and number of waves interviewed, the resulting likelihood ratio is 212, compared to 607 for our final regression with the life-cycle variables; for St. Joseph County the value is 200 compared to 569.

DISTRIBUTION OF HOUSEHOLD TYPES

Our predictions of eligibility are improved by using life-cycle stage and minority status, but we then must estimate the distribution of those characteristics within each stratum. In this section we explain how we went about making those estimates. (Note that minority status was a factor only in St. Joseph County.)

---

3. Interaction of year-specific dummies with each of the nonelderly life-cycle stages, to account for fluctuation in unemployment and its effects on each life-cycle stage.

The χ^2 statistic for the hypothesis that all variables in the full model and not in the shown model are zero is 4.4 with 9 degrees of freedom in Brown County, and 1.28 with 5 degrees of freedom in St. Joseph County. We eliminated coefficients only when the *t*-statistic was less than 1.

*Using dummies for each wave.

We have used an empirical Bayes framework to develop estimates of the distribution of household type within each stratum. The combined data from all the strata tell us much about the distribution of household type in the general population. Empirical Bayes techniques allow us to use this information to construct a prior density for the distribution within a stratum picked at random. The sample for each stratum is then combined with the prior to provide the posterior distributions. The estimate for the mean of the posterior distribution has the form of a James-Stein estimator. Many studies suggest that James-Stein estimates are superior to sample proportions, even when the prior distribution is only approximate.* Among the reported applications are the incidence of urban fires (Carter and Rolph, 1973**), toxoplasmosis in El Salvador (Efron and Morris, 1978), baseball batting averages (Efron and Morris, 1978), and income (Fay and Herriot, 1979).

Theoretical Model

The basis of our model is Bayes Theorem, which can be stated as follows. If we have reason to assume that the vector of random variables, $\tilde{\theta}$, has the density $g(\tilde{\theta})$, and take a sample and observe $\tilde{x}$, then the posterior density is $f(\tilde{\theta})$:

$$f(\tilde{\theta}) = \frac{g\,(\tilde{\theta})Pr\{\tilde{x}/\tilde{\theta}\}}{\int g\ \ (\tilde{z})Pr\{\tilde{x}/\tilde{z}\}dz}$$

In this case we wish to find the posterior distribution for each vector $\tilde{\theta}_i$ where θ_{ik} is the true proportion of households in stratum i who are in population subgroup k at a given wave t. Since we use a cross-sectional model, for simplicity the subscript t is omitted on all variables. Here $\tilde{x}_i$ is the vector, such that x_{ik} is the number of surveyed households of type k in panel stratum i. For any $\tilde{\theta}_i$, $Pr\{\tilde{x}_i/\tilde{\theta}_i\}$ is the multinomial distribution:

---

*A very good overview of why James-Stein estimates work may be found in B. Efron and C. Morris, "Stein's Paradox in Mathematical Statistics," *Scientific American*, May, 1978, pp. 119ff.

**Due to phenomena unique to their data, Carter and Rolph recommended the use of an empirical Bayes model different from Eq. (6).

$$Pr\{\tilde{x}_i/\tilde{\theta}_i\} = \begin{pmatrix} n_i \\ x_{i1}\ x_{i2}\ \cdots \end{pmatrix} \prod_{k=1}^{m} \theta_{ik}^{\ x_{ik}} \tag{5}$$

where n_i = the sample size, $n_i = \sum\limits_{k=1} x_{ik}$, and

m = the number of household types (m = 6 in Brown County, m = 12 in St. Joseph County).

We have chosen the Dirichlet distribution as the form of the prior density because this form is flexible, fits the data, and is computationally convenient for use with multinomial data. A Dirichlet distribution is completely specified by a parameter vector $\tilde{\alpha}$. If we then denote $\sum\limits_{k=1}^{m} \alpha_k$ by τ, the Dirichlet density is:

$$D(\tilde{\theta};\tilde{\alpha}) = \frac{\Gamma(\tau)}{\prod\limits_k \Gamma(\alpha_k)} \prod_{k=1}^{m} \theta_k^{\ \alpha_k - 1},$$

which has the properties:

$$E(\theta_k) = \frac{\alpha_k}{\tau}$$

and

$$\mathrm{Var}(\theta_k) = \frac{\alpha_k(\tau - \alpha_k)}{(\tau + 1)\tau^2}.$$

If we assume that the prior density of $\tilde{\theta}_i$ is of the form $D(\tilde{\theta}_i;\tilde{\alpha})$, the resulting posterior density is $f(\tilde{\theta}_i) = D(\tilde{\theta}_i;\tilde{\alpha} + \tilde{x}_i)$. (Note that $f(\tilde{\theta}_i)$ must be proportional to $\left[\prod\limits_k \theta_{ik}^{\ x_{ik}}\right]\left[\prod\limits_k \theta_{ik}^{\ \alpha_k - 1}\right]$ and that the constant of proportionality must produce a true density function.) Then our posterior estimate of θ_{ik} is the mean of the posterior distribution:

$$\hat{\theta}_{ik} = \frac{x_{ik} + \alpha_k}{n_i + \tau}.$$

An algebraic transformation shows this in the form of a James-Stein estimator:

$$\hat{\theta}_{ik} = \frac{n_i}{n_i + \tau} \cdot \frac{x_{ik}}{n_i} + \frac{\tau}{n_i + \tau} \cdot \frac{\alpha_k}{\tau} , \qquad (6)$$

since the weight on the sample mean is inversely proportional to the expectation over θ_{ik} of the variance of x_{ik} given θ_{ik}; and the weight on the prior mean is inversely proportional to the variance of θ_{ik}.

A rough idea of the improvement to be expected from the estimator of Eq. (6) can be seen as follows. If $g(\tilde{\theta})$ were the true description of $\tilde{\theta}$ across panel strata, then the expected mean square error or loss of the sample mean is:

$$L_1 = E\left[Var\left(\frac{x_{ik}}{n_i} \mid \theta_i\right)\right] \qquad (7)$$

$$= E\left[\frac{1}{n_i}\ \theta_i(1 - \theta_i)\right] = \frac{1}{n_i}\ \frac{\alpha_k(\tau - \alpha_k)}{\tau(\tau + 1)} .$$

The expected loss of $\hat{\theta}_{ik}$ if $g(\tilde{\theta})$ is known is:

$$L_2 = \frac{1}{n_i + \tau}\ \frac{\alpha_k(\tau - \alpha_k)}{\tau(\tau + 1)} . \qquad (8)$$

Thus the expected improvement increases with τ and decreases with n_i.

There is a limit to the flexibility of the Dirichlet in that it cannot accommodate a bimodal marginal distribution.* Because the distribution of life-cycle stage is radically different across tenure, we have estimated separate distributions of $g(\tilde{\theta})$ for renters and homeowners.** The flexibility of the Dirichlet is quite adequate to accommodate the information we have about the distribution of the vectors $\tilde{\theta}_i$.

---

*Except in the degenerate case where the modes are 0 and 1.

**Households in subsidized housing are treated as a separate group in St. Joseph County and omitted in Brown County where data on them were not available until wave 3.

Estimation

Here we describe the method of estimating the vector $\tilde{\alpha}$ that specifies the prior distribution. Separate distributions are estimated for each wave and tenure group. The sample consists of all households who provided complete household information at each wave, and thus includes the small number excluded from the logit estimate because they did not answer questions concerning income.

Let ρ_k = the expected portion of households of type k in a stratum picked at random $(\rho_k = \alpha_k/\tau)$, and

y_{ik} = the fraction of the sample who are type k $(y_{ik} = \frac{x_{ik}}{n_i})$, where $k = 1, 2, \ldots m$; $i = 1, 2, \ldots$ p, and

m = the number of household types (m = 6 in Brown County; m = 12 in St. Joseph County), and

p = the number of strata in the tenure and wave group of interest (p = 12 for baseline renters; p = 13 for renters at later waves; p = 6 for baseline owners; p = 7 for owners at later waves).

We estimate the parameter τ in the prior Dirichlet by analogy with the James-Stein estimator for the normal distribution case: The expected value of the variance of y_{ik} around the true mean θ_{ik} is σ_{ik}^2 where:

$$\sigma_{ik}^2 = \frac{\rho_k(1 - \rho_k)\tau}{n_i(\tau + 1)} .$$

The total variance of y_{ik} including the variance of θ_{ik} around ρ_k is

$$s_{ik} = \frac{\rho_k(1 - \rho_k)(n_i + \tau)}{(\tau + 1)n_i} = \frac{\rho_k(1 - \rho_k)}{\tau + 1} + \sigma_{ik}^2 .$$

As in the normal distribution case the ratio of these variances is the weight on the mean of the prior distribution in the equation for the mean of the posterior distribution. In particular,

$$\frac{\sigma_{ik}}{s_{ik}} = \frac{\tau}{n_i + \tau} ,$$

which is independent of k.

We will obtain an unbiased estimate of $\frac{\sigma_{ik}^2}{s_{ik}}$ by combining estimates of s_{ik} and σ_{ik}^2. First we get an unbiased estimate of $\rho_k(1 - \rho_k)/(\tau + 1)$ by using the following weighted sum of squares. For arbitrary non-negative weights, w_i, we define:

$$S_k(\tilde{w}) = \sum_{i=1}^{p} w_i(y_{ik} - \rho_k(\tilde{w}))^2 \tag{9}$$

where

$$\rho_k(\tilde{w}) = \sum_{i=1}^{p} w_i y_{ik} / \sum_{i=1}^{p} w_i , \tag{10}$$

then

$$E(\rho_k(\tilde{w})) = \rho_k$$

and

$$E(S_k(w)) = \sum_i var(y_{ik})\, w_i(1 - \frac{w_i}{\sum w_j}) .$$

If $w_i = \frac{n_i}{n_i + \tau}$, then $\rho_k(\tilde{w})$ is the minimum variance estimate of ρ_k and $E(S_k(w)) = [p - 1]\,\frac{\rho_k(1 - \rho_k)}{\tau + 1}$.*

---

*For any constant c, if $w_i = \frac{c\, n_i}{n_i + \tau}$, then $\rho_k(\tilde{w})$ is still the minimum variance estimate of ρ_k and $E(S_k\,(\tilde{w})) = c \cdot [p - 1]\, \rho_k\,(1 - \rho_k)/(\tau + 1)$.

If y_i were from a normal distribution and $c = \frac{1}{p} \sum_i \left(\frac{n_i + \tau}{n_i} \right)$, then

We denote $\rho_k(1 - \rho_k)\ \tau/(\tau + 1)$ by σ_k^2, so that $\sigma_{ik}^2 = \frac{1}{n_i} \cdot \sigma_k^2$.

Now for each i, $E(y_{ik}(1 - y_{ik})) = \frac{(n_i - 1)}{n} \sigma_k^2$. Therefore we shall estimate σ_k^2 by a weighted sum of $\frac{n_i}{n_i - 1}\ y_{ik}(1 - y_{ik})$, with the weight on stratum i proportional to n_i:*

$$\hat{\sigma}_k^2 = \sum_i \frac{n_i^2}{(n_i - 1)} \cdot y_{ik}(1 - y_{ik})/\sum_i n_i \tag{11}$$

In the next step we assume that $\hat{\sigma}_k^2$ is independent of $S_k(\tilde{\omega})$, which is only approximately true. (The assumption can be justified on the basis that since y_{ik} comes from the binomial with large n_i, y_{ik} is approximately normal given θ_{ik}.) Assuming independence,**

$$E\left(\frac{S_k(\tilde{\omega})}{\sigma_k^2}\right) = \frac{(p - 1)}{\tau} .$$

Consequently we find the τ that is the solution to

$$\frac{1}{m} \sum_k \frac{1}{\hat{\sigma}_k^2} \sum_i \frac{n_i}{n_i + \tau} (y_{ik} - \rho_k(\tau))^2 = \frac{p - 1}{\tau} \tag{12}$$

$\frac{1}{p - 1} S_k(\tilde{\omega})$ is the minimum variance estimate of $\frac{1}{p} \sum_i s_{ik}$. This is a major reason for our choice of w_i. Since c drops out of the estimating equation, we keep $c = 1$ throughout.

*The expected value over θ_{ik} of the variance of $y_{ik}(1 - y_{ik})$ given θ_{ik} is proportional to $(n_i - 1)^2/n_i^3$ plus a term of order $1/n_i^2$.

**This equation ignores the correction factor in $E(\frac{1}{\hat{\sigma}_k^2})$ since the total sample is so large.

where

$$\rho_k(\tau) = \sum_i \frac{n_i}{n_i + \tau} y_{ik} / \sum_i \frac{n_i}{n_i + \tau} \quad .$$

This procedure yields an unbiased estimate of $\frac{1}{\tau}$ and consequently of $\frac{n_i + \tau}{\tau}$ for each n_i, which is just the inverse of the quantity we seek. We use a correction factor analogous to the normal case, multiplying the τ found from Eq. (12) by $((m - 1)(p - 1) - 2)/(m - 1)(p - 1)$, since $\sum_k S_k$ has $(m - 1)(p - 1)$ degrees of freedom.

Table 2.6 gives the calculated values of τ for each wave, site, and tenure group. Using the loss equations (7) and (8), we can interpret τ as equivalent to the number of samples added to each stratum by the prior. Consequently the table also shows the average sample size per stratum, $\bar{n}$, and gives τ as a percent of $\bar{n} + \tau$. The Dirichlet estimator provides only modest improvement in our estimates for rental units, but quite substantial improvements for homeowners, particularly in St. Joseph County. The distributions of life-cycle stage and minority status categories are quite similar in each of the St. Joseph County homeowner strata, and hence τ is very large relative to sample sizes. (It is fortunate that we get a larger improvement for homeowners than for renters, because as we noted earlier homeowners are under-represented in our sample. Consequently our estimates of them need more improvement.)

Interpreting the ratio of τ to sample size as the percent by which the loss function decreases requires that the Dirichlet accurately describe the distribution of $\tilde{\theta}$ across strata. We do not have independent measures of the actual $\tilde{\theta}_i$ with which to compare the Dirichlet estimates. However, new households were interviewed at each wave. We have used baseline data to predict the distribution of life-cycle stage (and minority status) within each panel stratum for households interviewed

Table 2.6

ESTIMATES OF DIRICHLET PARAMETER τ AND AVERAGE SAMPLE SIZE PER STRATUM

Site and Wave	Renter			Owner		
	τ	Average sample size per stratum ($\bar{n}$)	$\frac{\tau}{\bar{n} + \tau}$ x 100	τ	Average sample size per stratum ($\bar{n}$)	$\frac{\tau}{\bar{n} + \tau}$ x 100
Brown County						
Wave 1	15.3	172.8	8	18.1	107.3	14
Wave 2	14.3	144.5	9	34.8	86.6	29
Wave 3	15.4	132.3	10	22.9	84.1	21
Wave 4	18.9	135.5	12	27.7	84.7	25
St. Joseph County						
Wave 1	29.6	115.8	20	85.6	94.7	47
Wave 2	30.1	109.8	22	86.0	84.7	50
Wave 3	27.0	109.8	20	94.4	87.0	52

SOURCE: Estimated by authors from household surveys at waves 1-4 in Brown County, waves 1-3 in St. Joseph County.

in the last wave in each site,* but not at baseline. We compared the predictions from our procedure with the baseline sample means using the loss function:

$$L_i = \sum_k (\hat{\theta}_{ik} - A_{ik})^2$$

where A_{ik} is the proportion of last-wave households not interviewed at baseline in stratum i who are of type k. The Dirichlet estimate from baseline data had smaller loss than the baseline stratum average in 14 out of 18 strata in Brown County and 16 out of 18 in St. Joseph County. The average reduction in squared error of the estimate was 11 percent in Brown County and 18 percent in St. Joseph County. St. Joseph County homeowner strata showed the greatest average improvement of all four tenure/site categories.

PRECISION OF ELIGIBILITY ESTIMATES

The estimated numbers of eligibles in each tenure and life-cycle group provided by our model show great improvement in precision over the estimates derived from weighted counts. Table 2.7 compares standard deviations for each set of estimates.

In Brown County the standard deviation in each life-cycle group (averaged across tenure) ranged from 27 to 43 percent lower than those for the weighted count estimates. (Change is expressed as a percent of the standard deviation of the weighted count.) In St. Joseph County, only one life-cycle group (elderly singles) showed as little as 10 percent improvement; all the others improved by 40 percent or better. To appreciate the magnitude of this improvement, consider what would happen if the sample size had been doubled. Then the comparable improvement in precision for the weighted counts would have been only 29 percent.

Our estimates of eligibles by tenure likewise showed noticeable improvement. On average they were 10 percent more precise, which could be achieved with weighted counts only if the sample size were increased by 23 percent. The standard deviation for renters improved more than

---

*Wave 4 in Brown County and wave 3 in St. Joseph County.

Table 2.7

COMPARISON OF STANDARD DEVIATIONS OF ESTIMATED NUMBER OF ELIGIBLE HOUSEHOLDS IN 1977 FROM WEIGHTED COUNTS AND THE COMBINED DIRICHLET-LOGIT MODEL BY LIFE-CYCLE STAGE AND TENURE

Life-Cycle Stage	Standard Deviation of Estimate								
	Renters			Owners			Total		
	Weighted Count	Dirichlet-Logit Model	Percent Decrease[a]	Weighted Count	Dirichlet-Logit Model	Percent Decrease[a]	Weighted Count	Dirichlet-Logit Model	Percent Decrease[a]
Brown County									
Elderly couple	52	39	25	271	187	31	276	198	28
Elderly single	75	75	0	283	182	36	293	215	27
Single parent	110	91	17	288	143	50	309	176	43
Young couple, young children	95	67	29	264	151	43	280	179	36
Other	63	50	24	234	112	52	242	137	43
Total	162	149	8	480	436	9	507	501	1
St. Joseph County									
Elderly couple	64	45	30	695	409	41	698	418	40
Elderly single	130	109	16	669	502	13	682	611	10
Single parent	156	126	19	674	383	43	692	413	40
Young couple, young children	90	64	29	513	244	52	521	269	49
Other	85	57	33	501	196	61	508	219	57
Total	214	178	17	1,219	1,058	13	1,238	1,091	12

SOURCE: Standard deviation of weighted counts were tabulated by the authors from the household survey for wave 4 in Brown County and wave 3 in St. Joseph County. The standard deviations of the Dirichlet-logit model are from the simulation of Appendix A applied to the same survey waves.

[a]Improvement in standard deviation expressed as a percent of the weighted count.

that for homeowners, because of the way we constructed the logit model: Households interviewed more than once provided only a single sample point for the logit, and since renters move more frequently than homeowners they provided more points. As a result, the sample on which the logit model is based is skewed even more heavily toward renters than the original survey sample. Consequently our estimates of panel stratum effects are better for renters than for owners, which cancels out the greater power of the prior Dirichlet for owners.

The results presented in Table 2.7 are for 1977, but they are typical of those for other waves as well.* Our estimates of eligibles by tenure for each survey wave showed improvement in precision for every group but one (Brown County homeowners at baseline). For each of the two remaining survey waves in Brown County the estimate of total eligibles was improved by 12 percent. In St. Joseph County, for the two waves not included in the table the improvement was 7 and 13 percent in the estimates of total eligibles. Thus the improvement in estimates for life-cycle groups also shows up in our estimates of total number of eligibles.

The model also produces striking improvements in our estimates of minority eligibles, as shown in Table 2.8. Our precision increased by 49 percent or more at each wave. Again, estimates for homeowners improved more than those for renters.

---

*For example, our model yielded decreases in the standard deviation for the baseline estimates of each life-cycle group in Brown County of 24, 6, 43, 44, and 57 percent; and in St. Joseph County of 46, 23, 35, 55, and 39 percent.

Table 2.8

COMPARISON OF STANDARD DEVIATIONS OF ESTIMATED NUMBER OF ELIGIBLE MINORITY HOUSEHOLDS IN 1977 FROM WEIGHTED COUNTS AND THE COMBINED DIRICHLET-LOGIT MODEL, BY TENURE: ST. JOSEPH COUNTY, 1977

	Standard Deviation of Estimate								
	Renters			Owners			Total		
Year	Weighted Count	Dirichlet-Logit Model	Percent Decrease[a]	Weighted Count	Dirichlet-Logit Model	Percent Decrease[a]	Weighted Count	Dirichlet-Logit Model	Percent Decrease[a]
1975	126	101	20	477	220	54	493	251	49
1976	136	108	21	558	236	58	574	269	53
1977	139	108	22	803	304	62	815	331	59

SOURCE: Standard deviation of weighted counts were tabulated by the authors from the household survey at waves 1-3 in St. Joseph County. The standard deviations of the Dirichlet-Logit model are from the simulation of Appendix A applied to the same survey waves.

[a] Improvement in standard deviation expressed as a percent of the standard deviation of the weighted count.

III. MEASURING TURNOVER WITHIN THE ELIGIBLE POOL

In order to estimate the proportions of households who enter and leave eligibility each year, we examine turnover between each of two pairs of years at each site: 1975-76 and 1976-77. About half of the households interviewed in the second year had been interviewed the previous year as well, so we knew the amount of turnover in that group. However, because the survey waves followed residential properties across time rather than households, for the rest we had to rely on retrospective information in order to estimate prior eligibility. Two data sets gave us what we needed: 1) changes in the marital status of the head(s) of household, which identified newly formed households; and 2) a retrospective employment and income history for each household during the previous year.

To test the frequency and direction of possible errors resulting from using the retrospective information, we use data on households surveyed more than once. The same data allow us to check the accuracy of our predictions of prior wave eligibility. In Sec. IV we have applied these equations to households not interviewed at a previous wave to determine their eligibility the year before. Our assumption is that errors in survey responses are similar for both sets of households, those who were previously interviewed and those who were not; but not that turnover rates are necessarily the same for both groups.

DATA LIMITATIONS

Household formation and dissolution contribute to our turnover estimates, but no record is available of the eligibility status of the household(s) preceding either change. We define a household by the identity of the two spouses in the case of couples and by its head otherwise. When a child leaves home, for instance, the parents' household is assumed to be the same, but if a couple separates, that household dissolves and two new ones are formed.

Our turnover estimates for households not newly formed are based solely on changes in income eligibility; since we exclude changes in eligibility based on assets and family composition, our estimates necessarily understate actual turnover. However, the difference is probably small. In St. Joseph County, among households who were interviewed for two waves and changed eligibility, less than 3 percent did not experience a change in income eligibility.

Most income eligibility turnover is associated with changes in employment. The main source of data for estimating prior wave eligibility of households not interviewed before is the retrospective employment history. However, such data are often inaccurate. When we tested this source of information by using it to predict prior eligibility for cases in which we knew the answer, about 15 percent of our predictions were incorrect. In addition, most of these errors were in a single direction: Cases that were eligible were predicted to be ineligible. However, we can eliminate this bias because the errors in our retrospective data were correlated in a predictable way with other information from the same survey wave. Thus we were able to estimate the probability of error in predicting eligibility from the employment history contained in the survey record and improve on the prediction from the retrospective employment history.

ESTIMATION OF PRIOR WAVE ELIGIBILITY

First we used the retrospective job history of each household head to estimate the household's benefit level at the previous wave, assuming that any other income not job-related was constant over the intervening year. Then we divided the sample of households interviewed over two successive waves into those whose estimated past benefit level was above and below the cutoff for eligibility. Logit regressions for each group, based on data from the later wave, estimated the probability that a household was income eligible at the prior wave.

The independent variables include information about total annual income which we use to estimate the reliability of employment history--that is, how likely it is to be inaccurate or insufficient. They are as follows:

- Amount of unemployment benefits for the year preceding the survey for heads who reported no months of unemployment in that year.*
- The difference between wages calculated from the employment history for the year before the survey and total household wages reported for the same year.* If the total income is higher than the wages, there is a greater likelihood that either wage rates are overstated or the household exaggerated the amount of employment. This value (truncated at zero) is therefore used in the regression on households whose job history indicated lack of eligibility. Negative values (obtained if the total income is less than the wages, and also truncated at zero) are used in the other regression.
- A dummy variable indicating whether the primary head had worked continuously at the same job since the last survey. Since only the last wage is reported for each job, past income from steady employment is overestimated.
- The amount of Social Security income when one or both heads were retired at the previous survey. Since Social Security payments rose faster than the standard cost of housing, this variable should increase the probability that the household had been eligible.
- A dummy variable indicating whether the household is currently income eligible.
- The log of the negative of the prediction of past benefit level plus a constant. This accounts for the higher probability of misprediction for households closer to the borderline of eligibility than for households further away.

---

*Both unemployment benefits and the wage differential are divided by the difference between predicted past benefit level and the eligibility cutoff ($120).

The results of the regressions are presented in Table 3.1. Each of them shows a statistically significant improvement over the estimate of eligibility based solely on retrospective employment history ($P < .0001$).

Table 3.2 shows the distribution of cases by predicted and actual status at the prior wave. The prediction classifies 89 percent of the cases correctly; a prediction of .5 or greater classifies a household as eligible. In addition, most predictions are at the extreme values of the range; only 11 percent of them lie between .25 and .75. An examination of accurate classification by life-cycle group showed only modest variation; in St. Joseph County it ranged from 84 to 91 percent. The regressions improve the number of cases classified correctly by only 4 percent over those predicted from their job histories. However, the regressions avoid the bias against prior wave eligibility found in the uncorrected retrospective data and reproduce exactly the amount of turnover in the sample.

To obtain the population estimates reported in Sec. IV, the equations in Table 3.1 were applied to households that were not interviewed at an earlier wave, but existed at that wave. The sum of the probabilities over this sample, added to the count of those known from the stayer sample to be eligible at the previous wave, yielded the amount of turnover in each life-cycle and tenure group. Aggregations across groups were made using our previously described estimates of the number of households in each eligibility, life-cycle, and tenure group.

Table 3.1

EFFECTS OF INCOME FACTORS ON LIKELIHOOD OF HOUSEHOLD'S ELIGIBILITY LAST YEAR, GIVEN STATUS FROM HEADS' JOB HISTORIES

Households Eligible Last Year According to Job Histories

Variable	Brown County		St. Joseph County	
	Coefficient	*t*-statistic	Coefficient	*t*-statistic
Constant	19.343	2.2	15.530	1.5
Wage shortfall from job history estimate	-0.442	2.8	-0.727	2.6
Income eligible now	0.604	1.0	1.587	2.0
Log of negative benefit level	-2.263	2.0	-1.881	1.4
Chi-square	23		21	
Sample size	365		283	

Households Ineligible Last Year According to Job Histories

Variable	Brown County		St. Joseph County	
	Coefficient	*t*-statistic	Coefficient	*t*-statistic
Constant	24.990	4.3	31.197	4.7
Wages in excess of job history estimate	0.497	2.9	0.383	1.9
Unemployment benefits	0.797	1.9	1.351	1.6
Income eligible now	1.002	3.0	1.262	3.0
Social Security income	0.0002	2.3	-0.00004	.4
Same job both waves	-0.848	2.8	-0.617	1.7
Log of negative benefit level	-3.230	4.7	-3.990	4.9
Chi-square	144		84	
Sample size	787		483	

SOURCE: Estimated by authors from household survey records for 1975-77.

NOTE: Sample includes only households interviewed at two successive waves. See text for detailed definition of independent variables.

Table 3.2

COMPARISON OF PREDICTED AND ACTUAL INCOME ELIGIBILITY
IN THE PREVIOUS YEAR

Predicted Probability of Eligibility in Previous Year (%)	Status of Household at Previous Year					
	Brown County			St. Joseph County		
	Ineligible	Eligible	Total	Ineligible	Eligible	Total
0 - 50	623	47	670	377	33	410
25 - 50	62	30	92	32	19	51
50 - 75	13	27	40	15	14	29
75 - 100	39	311	350	22	254	276

SOURCE: Tabulated by authors from household survey records for 1975-77.
NOTE: Sample includes only households interviewed at two successive waves.

IV. ESTIMATES OF ELIGIBILITY AND PARTICIPATION

Using the methods described in Secs. II and III, we estimated the size and composition of the household populations eligible for housing allowances in Brown and St. Joseph counties for each year covered by our annual household surveys: 1974-77 for Brown County and 1975-77 for St. Joseph County.* By combining those estimates with administrative counts of enrollees and allowance recipients for corresponding dates, we also estimated participation rates for the eligible population of each site and for specific household types.

This section reports both the eligibility estimates and the corresponding participation rates, and comments on their programmatic implications. It compares the HASE results on participation rates with those of the Housing Allowance Demand Experiment, and explains briefly how the findings here reported will be used in further HASE research.

THE ELIGIBLE POPULATION

As noted in the Introduction, the size and composition of the household population eligible for housing allowances should be interesting to policy analysts for two reasons. First, the number of eligibles, together with their average entitlement, sets an upper limit on the fiscal liability of the program. Second, the eligibility rules only partly characterize the eligible population, primarily with respect to income and household size. Policy analysts are also interested in other characteristics, such as housing tenure, race, and household composition, of those who are eligible. By applying the eligibility rules to the actual populations of two metropolitan housing markets, we learn much about the probable composition of the population that would be eligible for a national program operating under the same or similar rules.

---

*The fourth survey wave in St. Joseph County was conducted in 1978, but the data were not ready for analysis when this study was done.

Eligibility in 1977

One fact established by our research is that the population of eligibles is constantly changing as to membership, composition, and size. Consequently, it is not adequately described by its makeup on any specific date. However, such a cross-sectional view is a useful first approximation, and is also appropriate for calculating eligibility rates. Here we present estimates of the populations eligible for housing allowances in 1977, classifying them by housing tenure, life-cycle stage, and (in St. Joseph County) race of head.

Not all those whom we count as eligible in 1977 would have been permitted to enroll in the program. As was explained in Sec. I, the income limit for enrollment is tied to the standard cost of adequate housing, which is updated periodically by the HAOs to reflect market-wide inflation in housing costs. We interpolated housing costs and income limits between the HAOs' updates in order to determine which households were unable to afford adequate housing under the one-fourth-of-income rule at the time of each household survey. The effect on eligibility of periodic rather than continuous updating of income limits is discussed later in this section.

Table 4.1 shows our eligibility estimates for 1977. As explained in Sec. II, we estimated the distribution of households in each stratum of each county by type of household with the results summed for each county shown in the table under "Population." We also estimated the probability that a household of a specific type in each panel stratum would be eligible. Multiplying this probability by the estimated number of such households and summing over panel strata yields the number of eligible households shown in the second column of the table. The ratio of these numbers yields the percent eligible: 16.5 percent of all Brown County households (about 7,900 households) and 21.4 percent of St. Joseph County households (about 15,400 households).

As one might expect, eligibility rates among renters are much higher than among homeowners; the latter are typically more prosperous. However, because homeowners are also much more numerous in the population, their low eligibility rates nonetheless yield a majority of the eligibles in both sites. About 47 percent of the eligibles in

Table 4.1

NUMBER OF HOUSEHOLDS AND FRACTION ELIGIBLE, BY TENURE, LIFE-CYCLE STAGE, AND RACE OF HEAD: BROWN AND ST. JOSEPH COUNTY, 1977

Household Type	Brown County			St. Joseph County		
	Number of Households			Number of Households		
	Total[a]	Eligible	Percent Eligible	Total[a]	Eligible	Percent Eligible
Renters						
Life-cycle stage:						
Elderly couple	489	266	54.4	747	277	37.0
Elderly single	1,226	771	62.9	1,797	1,271	70.7
Single parent	1,693	1,296	76.6	2,370	1,925	81.2
Young couple, young children	2,732	845	30.9	2,298	642	27.9
Other couple	2,855	511	17.9	2,935	529	18.0
Nonelderly single	5,347	0[b]	0.0[b]	5,472	0[b]	0.0[b]
All stages	14,342	3,689	25.7	15,619	4,644	29.7
Race of head:						
White, non-Hispanic	(c)	(c)	(c)	12,889	3,211	24.9
Other	(c)	(c)	(c)	2,730	1,433	52.5
All races	14,342	3,689	25.7	15,619	4,644	29.7
Owners						
Life-cycle stage:						
Elderly couple	3,390	1,087	32.1	7,989	2,399	30.0
Elderly single	2,615	1,007	38.5	6,792	4,193	61.7
Single parent	1,778	677	38.1	3,772	1,876	49.7
Young couple, young children	9,612	858	8.9	12,981	1,236	9.5
Other couple	13,603	594	4.4	17,699	1,052	5.9
Nonelderly single	2,732	0[b]	0.0[b]	7,036	0[b]	0.0[b]
All stages	33,730	4,223	12.5	56,267	10,755	19.1
Race of head:						
White, non-Hispanic	(c)	(c)	(c)	51,446	9,297	18.0
Other	(c)	(c)	(c)	4,821	1,458	30.3
All races	33,730	4,223	12.5	56,267	10,755	19.1
Total						
Life-cycle stage:						
Elderly couple	3,879	1,353	34.9	8,736	2,676	30.6
Elderly single	3,841	1,778	46.3	8,589	5,464	63.6
Single parent	3,471	1,973	56.9	6,141	3,801	61.9
Young couple, young children	12,344	1,703	13.8	15,279	1,878	12.3
Other couple	16,458	1,105	6.7	20,633	1,580	7.7
Nonelderly single	8,079	0[b]	0.0[b]	12,508	0[b]	0.0[b]
All stages	48,072	7,912	16.5	71,886	15,399	21.4
Race of head:						
White, non-Hispanic	(c)	(c)	(c)	64,336	12,508	19.4
Other	(c)	(c)	(c)	7,550	2,891	38.3
All races	48,072	7,912	16.5	71,886	15,399	21.4

SOURCE: Estimated by the authors from the household surveys for wave 4 in Brown County and wave 3 in St. Joseph County.

NOTE: Eligibility rates by household type were estimated from 1977 data by means of a logit model whose parameters were estimated from a multiwave sample. The number of households of each type in each county was estimated by means of a Dirichlet model. Both models are described in Sec. II. The estimated number of eligibles is the product of the eligibility rate and the number of households "at risk" in 1977. Distributions may not add exactly to totals because of rounding.

[a] Resident landlords and residents of subsidized housing are excluded.

[b] Single persons under 62 are classified as ineligible, although a few were eligible due to handicaps, disabilities, or residential displacement.

[c] Not estimated for Brown County, where nearly all residents are non-Hispanic whites.

Brown County and 30 percent in St. Joseph County are renters. The *Third Annual Report of the Housing Assistance Supply Experiment* (1977) shows that the market value of homes is higher in Brown County than in St. Joseph County (Fig. 4.6 of that report) but that rents do not differ much between sites (Fig. 4.8). Consequently, home purchase is more of a bargain in St. Joseph than in Brown County and more low-income households are owners there.*

Elderly singles and single parents have much higher eligibility rates than other life-cycle groups. Over 70 percent of St. Joseph County renters (and only a slightly smaller proportion in Brown County) in these two life-cycle groups are eligible. Among the elderly, the difference in eligibility rates between singles and couples is due to both a greater amount of non-Social Security income and larger Social Security payments relative to R^* for couples than for singles. Examining differences between sites and combining tenure groups, we find that each life-cycle group has roughly similar eligibility rates in both sites. The exception is elderly singles, who are more likely to be eligible in St. Joseph County. This was also true at each of the surveys prior to 1977, but the difference did grow slightly over time. Finally, in St. Joseph County, we note that households headed by nonwhites are twice as likely to be eligible as those headed by whites.

Table 4.2 offers another perspective on the pool of households that were eligible in 1977, showing the composition of each county's pool by life-cycle stage and (for St. Joseph County) race of head. This perspective best displays the consequences of the eligibility rules in determining what kinds of households are offered help by the allowance program. There are broad similarities of composition between the two eligible populations, but also some sharp differences. Single parents account for about a fourth of all eligibles in each site. Households headed by elderly persons account for 40 percent in Brown

---

*Another reason for the disparity between counties in the percent of eligibles who rent is that we have excluded households in subsidized housing from our estimates of eligibles. St. Joseph County contained about 2,000 subsidized renter and 600 subsidized owner households, most of whom meet program eligibility standards. If we had included these, 35 percent of St. Joseph County eligibles would have been renters.

Table 4.2

COMPOSITION OF THE ELIGIBLE POPULATIONS IN BROWN AND ST. JOSEPH COUNTIES BY TENURE, LIFE-CYCLE STAGE, AND RACE OF HEAD: 1977

Household Type	Distribution of Eligibles (%)					
	Brown County			St. Joseph County		
	Renters	Owners	All Eligibles	Renters	Owners	All Eligibles
Life-cycle stage:						
Elderly couple	7.2	25.7	17.1	6.0	22.3	17.4
Elderly single	20.9	23.8	22.5	27.4	39.0	35.5
Single parent	35.1	16.0	24.9	41.5	17.4	24.7
Young couple, young children	22.9	20.3	21.5	13.8	11.5	12.2
Other couple	13.9	14.1	14.0	11.4	9.8	10.3
All stages	100.0	100.0	100.0	100.0	100.0	100.0
Race of head:						
White, non-Hispanic	*(a)*	*(a)*	*(a)*	69.1	86.4	81.2
Other	*(a)*	*(a)*	*(a)*	30.9	13.6	18.8
All races	100.0	100.0	100.0	100.0	100.0	100.0

SOURCE: Computed by authors from entries in Table 4.1.

NOTE: Distributions exclude resident landlords, residents of subsidized housing, and eligible single persons under 62. Distributions may not add exactly to 100.0 because of rounding.

[a] Not estimated for Brown County, where nearly all residents are non-Hispanic whites.

County but 53 percent in St. Joseph County. For the remaining categories, nonelderly couples, the relative frequencies are reversed: 36 percent in Brown County and 22 percent in St. Joseph County. That St. Joseph County's population is older accounts for part of the intercounty difference; the rest is due to the especially high eligibility rate of elderly single persons there.

The Changing Pool of Eligibles

During the several years covered by our surveys, the number of eligible households changed very little in either site. In Brown County, where the population was growing, the number of eligibles increased by 7 percent, 1974-77; in St. Joseph County, where the population was not growing, the number of eligibles did not change significantly over the shorter interval, 1975-77 (see Tables 4.3 and 4.4).

However, in both sites the composition of the eligible pool shifted measurably. The number of eligible single parents increased by at least a third in each site and the number of eligible elderly dropped by a fourth in Brown County only. The increase of eligible single parents reflects a similar growth in their numbers in the population at large, and an essentially static eligibility rate. The decrease of eligible elderly singles in Brown County reflected increased income, because Social Security benefits were rising faster than our interpolated income limits for this group. A similar decline in the eligibility rate was observed for all other elderly groups, but was offset by growth in the numbers of the elderly population.

Turnover in Eligibility Status

The changes in the total size and composition of the eligible pools were the net results of many individual changes in eligibility status, due to changes in household income, marital status, household formation and dissolution, and a small amount of migration into and out of each site. Our methods for measuring turnover, and their limitations, were explained in Sec. III; here, we show the pattern of flows into and out of eligibility in a typical year for each site (Figs. 1 and 2).

Table 4.3

ESTIMATED CHANGES IN BROWN COUNTY'S ELIGIBLE POPULATION
BY LIFE-CYCLE STAGE AND TENURE: 1974-77

	Eligible Households			Annual Increase (%)		
Household Type	1974	1977	t-statistic	1974-75	1975-76	1976-77
Tenure:						
Renter	3,313	3,689	1.98	8	2	1
Owner	4,090	4,223	.35	3	2	-2
Life-cycle stage:						
Elderly couple	1,325	1,353	.14	2	3	-3
Elderly single	2,380	1,778	1.83	-9	-5	-12
Single parent	1,423	1,973	2.82	26	9	1
Young couple, young children	1,411	1,703	1.72	6	7	7
Other	865	1,105	2.00	13	-3	16
Total	7,403	7,912	1.20	5	2	0

SOURCE: Estimates of number of eligible households are derived from the Dirichlet and logit models of Sec. II as applied to the household surveys of waves 1-4. The t-statistic uses the standard deviation of the estimate from the simulation described in Appendix A.

NOTE: Households living in subsidized dwellings and resident landlords are excluded. Nonelderly single-person households are excluded because almost all were ineligible until August 1977.

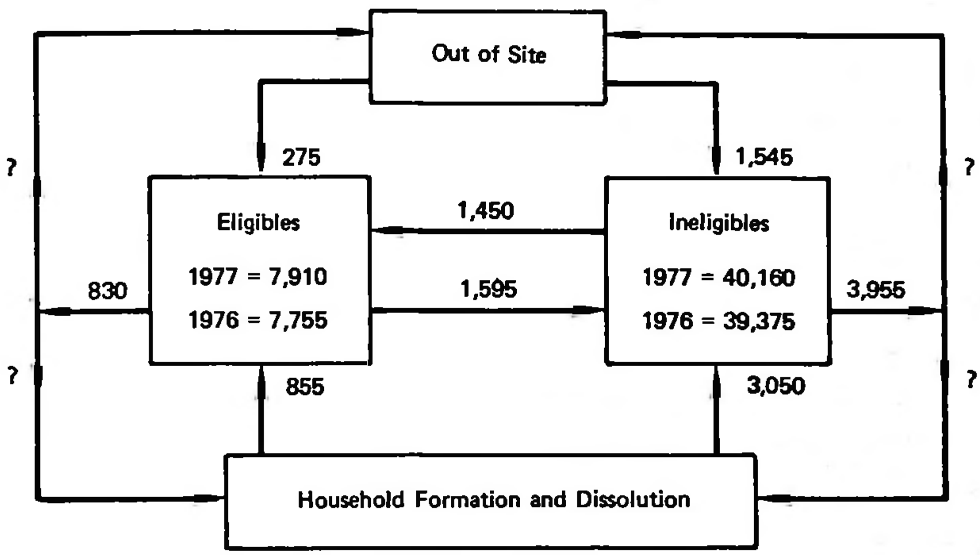
Table 4.4

ESTIMATED CHANGES IN ST. JOSEPH COUNTY'S ELIGIBLE POPULATION, BY LIFE-CYCLE STAGE, TENURE, AND RACE OF HEAD, 1975-77

Household Type	Eligible Households		t-statistic	Annual Increase (%)	
	1975	1977		1975-76	1976-77
Tenure:					
Renter	4,400	4,644	1.02	-1	7
Owner	10,707	10,755	.05	-2	2
Life-cycle stage:					
Elderly couple	2,491	2,676	.53	-3	11
Elderly single	5,487	5,464	.03	1	-1
Single parent	2,828	3,801	2.00	23	9
Young couple, young children	2,041	1,878	.61	-14	7
Other	2,260	1,580	2.62	-24	-8
Race of head:					
Minority	2,439	2,891	1.22	4	14
White non-Hispanic	12,668	12,507	.18	-2	1
Total	15,107	15,399	.32	-1	3

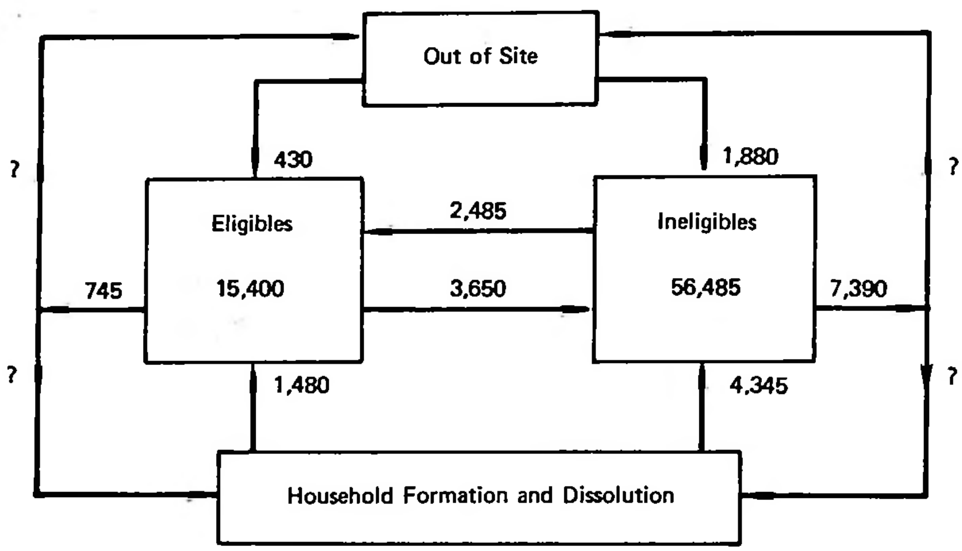
SOURCE: Estimates of eligible households are derived from the Dirichlet and logit models of Sec. II as applied to the household surveys of waves 1-3. The t-statistic uses the standard deviation of the estimate from the simulation described in Appendix A.

NOTE: Resident landlords and households living in subsidized dwellings are excluded. Nonelderly single-person households are excluded because almost all were ineligible until August 1977.



SOURCE: Numbers of households eligible and ineligible in 1977 are based on estimates in Table 4.1. Numbers for 1976 show the typical 2 percent annual growth rate. Flow rates are based on averages observed between 1975 and 1977.

Fig. 1—Annual changes in eligibility status of households in Brown County, 1975-77



SOURCE: Numbers of households eligible and ineligible based on estimates in Table 4.1. Constant number of eligible households assumed. Flow rates are based on averages observed between 1975 and 1977.

Fig. 2—Annual changes in eligibility status of households in St. Joseph County, 1975-77

In Brown County (Fig. 1), of nearly 8,000 eligible households, about 20 percent became ineligible in one year. Another 10 percent dissolved or moved away (our data do not distinguish between these two effects), leaving just under 70 percent of the 1976 eligible pool still eligible a year later. Nearly as many households already in the county became eligible as became ineligible, and newly formed or newly arrived eligible households more than offset the losses to dissolution and outmigration. The net result of all these changes was only a 2 percent increase in the number of eligibles.

The flows for St. Joseph County (Fig. 2) were similar. About 24 percent of the eligible households became ineligible in the course of a year and about 5 percent dissolved or moved away. These losses were replaced by newly eligible households, including some previously existing, some newly formed, and some inmigrants.

Changes within the eligible population are documented by tenure in Table 4.5 and by life-cycle stage in Table 4.6. The significant amount of turnover among elderly couples was often due to marginal eligibility; a small change in income or even none, given the rising cost of housing, could result in newly eligible status. Some of the elderly group experienced a drop in income when they went from an earned salary to retirement benefits. About 30 percent of the younger eligible couples also entered the pool each year because of income changes. But single parents, like elderly singles, were much more likely than other groups to have been eligible the previous year as well. Households in both those categories were the most consistently needy; once their incomes dropped to the level of eligibility, they were likely to remain eligible for several years. Migration into each site added about 3 percent to the number of eligibles each year, about the same as the inmigration rate for ineligibles.

PARTICIPATION

During the first two years of the program, the number of households enrolled grew rapidly, but by the third year the number of new enrollees was almost matched by those who terminated from the program. As shown in Table 4.7, enrollment then stabilized at about 40 percent of the eli-

Table 4.5

ELIGIBLE RENTER AND HOMEOWNER HOUSEHOLDS CLASSIFIED BY THEIR ELIGIBILITY STATUS A YEAR EARLIER

Tenure	Percent Of Currently Eligible Households by Earlier Status							
	Brown County				St. Joseph County			
	Eligible	Ineligible	Didn't Exist	Total	Eligible	Ineligible	Didn't Exist	Total
Renter	69.3	15.7	15.0	100.0	74.2	11.5	14.3	100.0
Homeowner	70.9	12.9	7.2	100.0	73.5	18.8	7.7	100.0
All	70.1	19.1	10.8	100.0	73.8	16.6	9.6	100.0

SOURCE: Estimated by authors from the averages for households eligible at waves 3 and 4 in Brown County, and waves 2 and 3 at St. Joseph County. (See Sec. III for methodology.)

Table 4.6

ELIGIBLE HOUSEHOLDS BY LIFE-CYCLE STAGE, CLASSIFIED BY ELIGIBILITY STATUS A YEAR EARLIER

	Percent of Currently Eligible Households by Earlier Status							
	Brown County				St. Joseph County			
Life-Cycle Stage	Eligible	Ineligible	Didn't Exist	Total	Eligible	Ineligible	Didn't Exist	Total
Single head with children	74.1	8.4	17.5	100.0	70.8	7.7	21.5	100.0
Elderly couple	73.2	26.6	.2	100.0	79.9	19.8	.3	100.0
Elderly single	90.5	7.6	1.9	100.0	87.0	11.7	1.3	100.0
Young couple, young children	52.4	30.1	17.5	100.0	50.1	31.8	18.1	100.0
Other	48.7	33.1	18.2	100.0	50.4	31.9	17.7	100.0

SOURCE: Estimated by authors from the averages for households eligible at waves 3 and 4 in Brown County, and waves 2 and 3 in St. Joseph County. (See Sec. III for methodology.)

Table 4.7

ENROLLMENT RATES BY TENURE AND SITE: 1974-78

Date	Renters		Homeowners		Total	
	Number Enrolled	Percent of Eligible Renters	Number Enrolled	Percent of Eligible Owners	Number Enrolled	Percent of Eligible Households
Brown County						
December 1974	813	23.2	557	13.3	1,370	17.8
June 1975	1,523	42.4	1,098	26.0	2,621	33.5
December 1975	1,534	42.3	1,097	25.6	2,631	33.3
June 1976	1,829	50.1	1,320	30.8	3,149	39.7
December 1976	1,925	52.3	1,387	32.7	3,312	41.8
June 1977	2,035	55.2	1,317	31.2	3,352	42.4
June 1978	2,122	57.5	1,144	27.1	3,266	41.3
St. Joseph County						
June 1975	321	7.3	503	4.7	824	5.5
December 1975	1,614	37.0	1,745	16.5	3,359	22.5
June 1976	2,193	49.5	2,466	23.3	4,659	31.0
December 1976	2,323	50.8	2,699	25.2	5,022	32.9
June 1977	2,653	57.1	3,332	31.0	5,985	38.9
December 1977	2,544	54.8	3,269	30.4	5,813	37.7
December 1978	2,504	53.9	3,177	29.5	5,681	36.9

SOURCE: Compiled by authors. Enrollees are from HAO administrative records; estimates of eligible households were derived by linear interpolation between survey waves for dates earlier than June 1977, and from the 1977 survey for June of that year and later.

NOTE: In order to maintain consistency over time, nonelderly single-person households are excluded because most of them were not eligible to join the program until August 1977. Resident landlords and households living in government-assisted properties are excluded from eligibility estimates.

gible households. In this section we present statistics on participation as of the end of the third program year, a date that approximately coincides with the last available survey results. Since enrollment had stabilized by then, these participation rates may be considered typical of the remainder of the experiment.

Not every enrollee receives an allowance. After a household has enrolled in the program, its dwelling is inspected. If it passes, the household will begin to receive payments immediately. Otherwise the household must repair the dwelling or move to an acceptable dwelling before receiving payments, and some choose to leave the program instead. Thus, although most enrollees eventually receive payments, the best measure of how well the program is serving various groups is the proportion of eligible households that currently receive payments.

Counts of enrollees and recipients are presented by housing tenure and life-cycle stage in Tables 4.8 and 4.9. Enrollment rates are higher than recipiency rates, of course. The difference between enrollment and recipiency rates is smaller for the elderly groups than for nonelderly, both because the elderly have fewer housing problems and because they have longer periods of recipiency (due to their longer periods of eligibility). Consequently, a smaller fraction of enrollees are so new to the program that they have not yet qualified for payments.

The overall pattern of participation by life-cycle stage and tenure is similar whether one uses enrollment or recipiency rates. The groups with the highest participation rates are the same as those with the highest eligibility rates: renters (vs. owners), single parents, and elderly singles. Elderly couples participate at a rate close to that of other couples in Brown County and are closer to them than to elderly singles in St. Joseph County; but in the latter instance the variance in the estimated participation rate is so large that the apparent difference between St. Joseph County elderly groups may be due to random error. Fig. 3 graphs the recipiency rate for each group within a 95 percent confidence limit, based on the assumption that our estimates of eligibles have a normal distribution.*

---

*Table A.1 in Appendix A shows the estimated standard deviation for the number of households eligible and rates of enrollment and recipiency, all by tenure, life-cycle stage, and minority status.

Table 4.8

BROWN COUNTY PARTICIPATION RATES AT END OF THIRD PROGRAM YEAR BY LIFE-CYCLE STAGE AND TENURE

Life-Cycle Stage	Number of Households			Participation Rates (%)	
	Eligible	Enrolled	Receiving Payments	Enrolled	Receiving Payments
Renters					
Elderly couple	266	91	84	34	32
Elderly single	771	528	496	68	64
Single parent	1,296	902	700	70	54
Young couple, young children	845	331	233	39	28
Other	511	183	145	36	28
Total	3,689	2,035	1,658	55	45
Owners					
Elderly couple	1,087	244	232	22	21
Elderly single	1,007	510	492	51	49
Single parent	677	252	216	37	32
Young couple, young children	858	147	117	17	14
Other	594	163	134	27	23
Total	4,223	1,316	1,191	31	28
Total					
Elderly couple	1,353	335	316	25	23
Elderly single	1,778	1,038	988	58	56
Single parent	1,973	1,154	916	58	46
Young couple, young children	1,703	478	350	28	21
Other	1,105	346	279	31	25
Total	7,912	3,351	2,849	42	36

SOURCE: Estimates of eligibles from wave 4 household survey (conducted mainly during the first half of 1977) are based on the methodology described in Sec. II. Counts of enrolled clients and those receiving payments are from HAO administrative records for June 1977.

NOTE: Nonelderly single-person households are excluded because most of them were not eligible to join the program until August 1977. Resident landlords and households living in government-assisted properties are excluded from eligibility estimates.

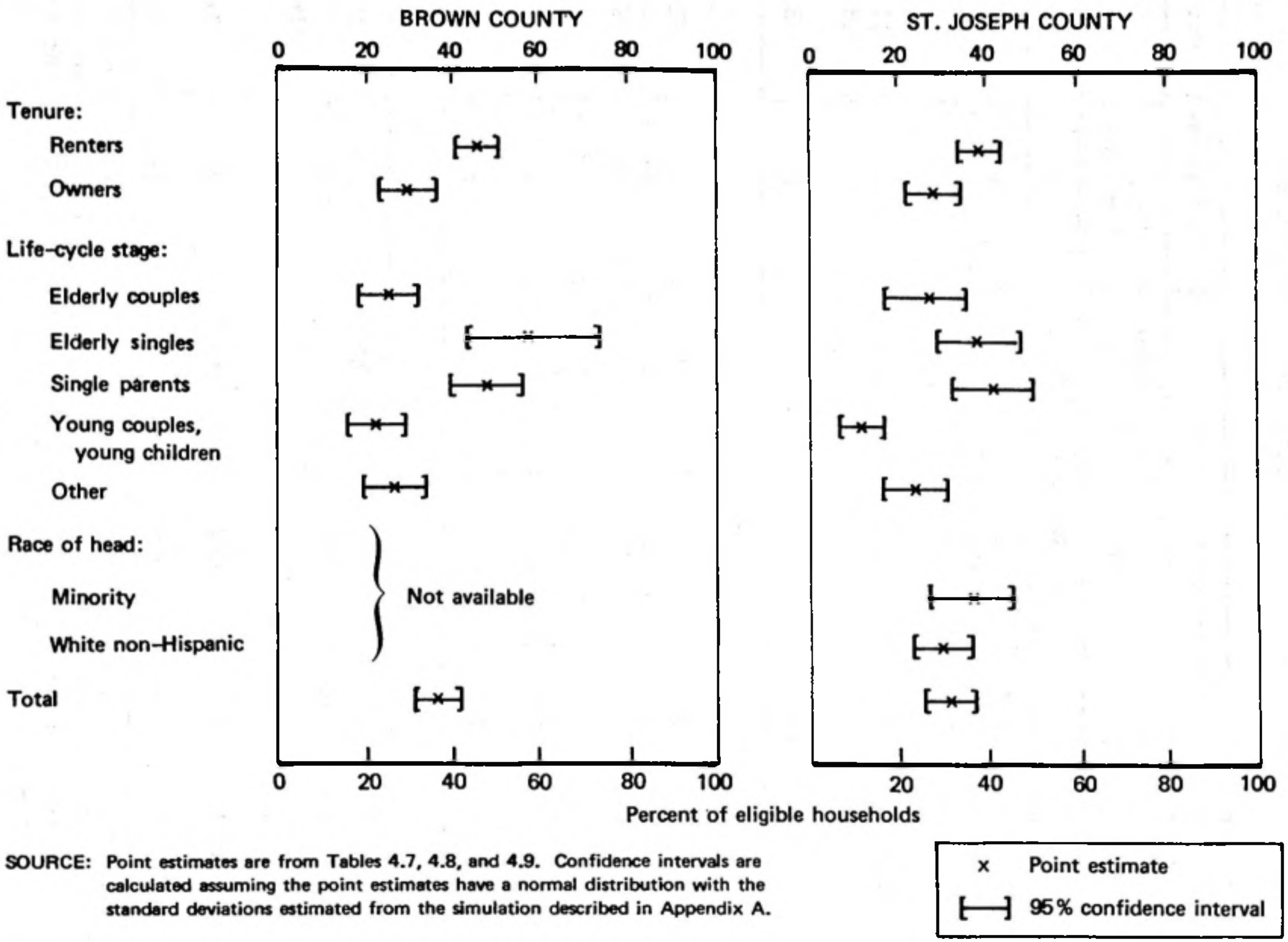
Table 4.9

ST. JOSEPH COUNTY PARTICIPATION RATES AT END OF THIRD PROGRAM YEAR, BY LIFE-CYCLE STAGE AND TENURE

Life-Cycle Stage	Number of Households			Participation Rates (%)	
	Eligible	Enrolled	Receiving Payments	Enrolled	Receiving Payments
Renters					
Elderly couple	277	71	65	26	23
Elderly single	1,271	559	505	44	40
Single parent	1,925	1,477	947	77	49
Young couple, young children	642	256	133	40	21
Other	529	186	120	35	23
Total	4,644	2,549	1,770	55	38
Owners					
Elderly couple	2,399	618	585	26	22
Elderly single	4,193	1,591	1,514	38	36
Single parent	1,876	672	535	36	29
Young couple, young children	1,236	133	91	11	7
Other	1,052	257	223	24	21
Total	10,755	3,271	2,948	30	27
Total					
Elderly couple	2,676	689	650	26	24
Elderly single	5,464	2,150	2,019	39	37
Single parent	3,801	2,149	1,482	57	39
Young couple, young children	1,878	389	224	21	12
Other	1,580	443	343	28	22
Total	15,399	5,820	4,718	38	31

SOURCE: Estimates of eligibles from wave 3 household survey (conducted mainly during the first half of 1977) are based on the methodology described in Sec. II. Counts of enrolled clients and those receiving payments are from HAO administrative records for December 1977.

NOTE: Most nonelderly single-person households were not eligible to join the program until August 1977. Because they had such a short time to join, their participation rates would not represent steady state; therefore we excluded them. Resident landlords and households living in government-assisted properties are excluded from eligibility estimates.



SOURCE: Point estimates are from Tables 4.7, 4.8, and 4.9. Confidence intervals are calculated assuming the point estimates have a normal distribution with the standard deviations estimated from the simulation described in Appendix A.

Fig. 3—Percent of eligible households receiving payments at end of year 3, by selected characteristics

The variation in enrollment rates among demographic groups is probably due in large part to the same sociological factors that affect participation rates in all income transfer programs. For instance, of all the programs surveyed by Bendick in his review of public assistance programs (1979), Aid to Families with Dependent Children (AFDC) had the highest rate (87 percent) of participation; its clientele, single parents, also have the highest enrollment rate in the allowance program. By contrast, a similar program for families with an unemployed head (AFDC-UF) achieved a participation rate of only 15 to 30 percent (Hosek, 1979). The target group of that program, two-parent families with children, has correspondingly lower enrollment rates in the housing allowance program as well.

Minorities are much more likely to enroll than nonminorities, regardless of tenure, as can be seen in Table 4.10. However, because of housing problems, a smaller percentage of enrolled minorities actually receive allowances than do other groups, so their recipiency rate is not much higher.

Table 4.11 shows the composition of the recipient population in each county and tenure group by life-cycle stage and (for St. Joseph County) race of head. Over two-thirds of recipient households are headed by a single person (67 percent in Brown County and 74 percent in St. Joseph County). The elderly, most of whom are singles, constitute 46 percent of the recipients in Brown County and 56 percent in St. Joseph County. Nonelderly couples are a greater percentage of recipients in Brown County than in St. Joseph County due to their larger share of the eligible population there, but still constitute a reasonably small fraction of recipients (22 percent in Brown County, 12 percent in St. Joseph County). The composition of the enrolled population is described in Table 4.12.

Participation Rates and the Changing Cost of Housing

The value of R^* automatically sets the income limits for eligibility because of the benefit formula. Our estimates of the number of eligible households, and therefore of participation rates, are based on the smooth inflationary rise in the cost of housing explained earlier.

Table 4.10

ST. JOSEPH COUNTY PARTICIPATION RATES AT END OF THIRD PROGRAM YEAR, BY RACE OF HEAD AND TENURE

Race of Head	Number of Households			Participation Rates (%)	
	Eligible	Enrolled	Receiving Payments	Enrolled	Receiving Payments
Renters:					
Minority	1,433	904	543	63	38
White non-Hispanic	3,211	1,645	1,227	51	38
Owners:					
Minority	1,458	597	487	41	33
White non-Hispanic	9,297	2,674	2,461	29	26
Both tenures:					
Minority	2,891	1,501	1,030	52	36
White non-Hispanic	12,508	4,319	3,688	35	29
Total	15,399	5,820	4,718	38	31

SOURCE: Estimates of eligibles from wave 3 household survey (conducted mainly during the first half of 1977) are based on the methodology described in Sec. II. Counts of enrolled clients and those receiving payments are from HAO administrative records for December 1977.

NOTE: Because most nonelderly single-person households were not eligible to join the program until August 1977, their participation rates would not represent steady state; therefore we excluded them. Resident landlords and households living in subsidized dwellings are excluded from eligibility estimates.

Table 4.11

COMPOSITION OF RECIPIENT POPULATIONS BY TENURE, LIFE-CYCLE STAGE, AND RACE OF HEAD: BROWN AND ST. JOSEPH COUNTIES, END OF THIRD PROGRAM YEAR

	Distribution of Recipients (%)					
	Brown County			St. Joseph County		
Household Type	Renters	Owners	All Recipients	Renters	Owners	All Recipients
Life-cycle stage:						
Elderly couple	5.1	19.5	11.1	3.7	19.8	13.8
Elderly single	29.9	41.3	34.7	28.5	51.4	42.8
Single parent	42.2	18.1	32.2	53.5	18.1	31.4
Young couple, young children	14.1	9.8	12.3	7.5	3.1	4.7
Other couple	8.7	11.3	9.8	6.8	7.6	7.3
All stages	100.0	100.0	100.0	100.0	100.0	100.0
Race of head:						
White, non-Hispanic	(a)	(a)	(a)	69.3	83.5	78.2
Other	(a)	(a)	(a)	30.7	16.5	21.8
All races	100.0	100.0	100.0	100.0	100.0	100.0

SOURCE: HAO administrative records for the end of year 3 (June 1977 in Brown County and December 1977 for St. Joseph County).

NOTE: Nonelderly single-person households are excluded.

[a]Not estimated for Brown County, where nearly all residents are non-Hispanic whites.

Table 4.12

COMPOSITION OF ENROLLED POPULATIONS BY TENURE, LIFE-CYCLE STAGE, AND RACE OF HEAD: BROWN AND ST. JOSEPH COUNTIES, END OF THIRD PROGRAM YEAR

Household Type	Distribution of Enrollees (%)					
	Brown County			St. Joseph County		
	Renters	Owners	All Enrollees	Renters	Owners	All Enrollees
Life-cycle stage:						
Elderly couple	4.5	18.5	10.0	2.8	18.9	11.8
Elderly single	25.9	38.8	31.0	21.9	48.6	36.9
Single parent	44.3	19.1	34.4	57.9	20.5	36.9
Young couple, young children	16.3	11.2	14.3	10.0	4.1	6.7
Other couple	9.0	12.4	10.3	7.3	7.9	7.6
All stages	100.0	100.0	100.0	100.0	100.0	100.0
Race of head:						
White, non-Hispanic	(a)	(a)	(a)	64.5	81.7	74.2
Other	(a)	(a)	(a)	35.5	18.3	25.8
All races	100.0	100.0	100.0	100.0	100.0	100.0

SOURCE: HAO administrative records for the end of year 3 (June 1977 in Brown County and December 1977 for St. Joseph County).

NOTE: Nonelderly single-person households are excluded.

[a] Not estimated for Brown County, where nearly all residents are non-Hispanic whites.

But because the HAOs only remeasure the cost of housing periodically, a gap develops between the R^* they officially recognize and the actual cost of housing. Since actual cost is the appropriate measure of need, that is what we use in estimating eligibility. But did the discrepancy between the actual cost of housing and the standard cost used by the HAOs significantly affect participation? How large is the resulting difference between the percentage of participants among the truly needy and among those eligible by HAO standards?

R^*, the standard cost of housing, was first measured before the program began. These values were not updated for 30 months in Brown County, or for 24 months in St. Joseph County. During these intervals, the incomes of initially eligible households increased, making them ineligible, even though they would have remained eligible if R^* and therefore the *income limits* had been continuously updated to reflect housing cost inflation. Other households who had been ineligible at baseline came to need program assistance long before they could enroll. We estimate that immediately prior to the first Brown County change in R^*, fully 20 percent of those who would have been eligible if housing costs had been updated were unable to enroll in it.

The delay had a harsher effect on some demographic groups than others. By our estimates, 15 percent among renters in need of assistance were unable to enroll, and 25 percent of needy homeowners. Among life-cycle groups, elderly couples were hurt the most; nearly half of those we counted as eligible could not qualify.

As Table 4.7 shows, the obsolescence in R^* caused a marked slowdown in the rate at which households joined the program. Enrollment in Brown County was essentially static from June 1975 until the R^* increase in April 1976, when it again began to climb. As the program matured, changes in R^* were made annually, cutting the effect of inflation on the number of needy but ineligible households by more than half. The increases in enrollment following the second R^* changes in May 1977 in Brown County and September 1977 in St. Joseph County are not nearly as noticeable.

Persons in charge of program management may be interested in participation rates measured according to the number of households eligible to join the program, rather than according to *all* households whose adjusted income is less than one-quarter of the standard cost of housing. In a time of continuing inflation, the needy and the eligible will never be identical because of the necessary administrative delay between measuring and implementing a change in the official HAO standard cost of housing (R^*). They are very close at the time when a new R^* goes into effect, but then diverge further and further until the next change occurs.

Assuming that just before a change in the rules the old R^* is 15 months old (allowing 12 months for an annual measurement and 3 months' delay for data analysis and implementation of the new standard), we estimate that, on average, 9 percent of the needy households could not qualify at any given time for participation. When the gap is widest between needy and eligible households, the participation rate of those able to apply is 11 percent higher than the participation rate of the needy. Thus our apparently steady-state enrollment rate of 40 percent of needy households corresponds to an enrollment rate of 44.4 percent of those who could apply at any given time during the year.

Comparison to Other Participation Studies

To date, ours is the most detailed study of participation rates in a housing allowance program. Others have reported on the subject, however, and in some instances their findings are somewhat different from ours. Below we compare each of them to the results of our investigation.

Phyllis Ellickson (1981) modeled households' decision to enroll in the housing allowance program, using data from St. Joseph County during the first year of the program. Her analysis shows that elderly singles were both less likely to know about the program and less likely to enroll. Our finding was quite different: Elderly singles were among the most likely to enroll. The discrepancy is probably due to the different intervals encompassed by the studies. In both sites the elderly entered the program rather slowly, but continued to enroll for a much longer period than did other groups. In St. Joseph County the number of enrolled elderly singles increased by 161 percent between the end of the first program year, when Ellickson collected her data, and the end of the third program year. In that same interval the number of other enrollees increased only 44 percent.

In a study of renter participation in the Demand Experiment carried out under the Experimental Housing Allowance Program, Kennedy and Macmillan (1980) found that nonminority households participated in the Housing Gap plan* at a rate 1.43 times the rate of black households. Our results for renters show that equal fractions of eligible minorities (primarily black households) and nonminorities are receiving payments in St. Joseph County,** the only site with a significant minority population.

Aside from the possibility that behavior differs between sites, three differences between the two experiments might explain the apparent contradiction in findings. First, enrollment in the Supply Experiment is open, while the Demand Experiment was limited to those invited to join. Second, the Supply Experiment promised that its allowances would continue for ten years, compared to the three years guaranteed by the Demand Experiment. Third, the Demand Experiment's standards for acceptable housing were stricter in the sense that more dwellings failed.

The two kinds of enrollment process suggest that self-selection may be more important in the case of the Supply Experiment; i.e., that white households may be less likely to join an open-enrollment program than one in which they were specifically invited to take part. The different proportions of eligible renter minority (63 percent) and nonminority (51 percent) households enrolled in the Supply Experiment supports this observation.***

At the time of the enrollment decision, a household applying to either program rarely knew whether or not its dwelling would pass or

---

*The Housing Gap plan included programs similar to the Supply Experiment's allowance program, as well as others in which adequate housing was determined by the amount of rent paid by tenants.

**Including owners, we get a slightly higher recipiency rate for minorities (36 percent) than for nonminorities (29 percent), but the difference is not statistically significant.

***Analysis of HADE acceptance rates showed that "rejection of the enrollment offer was based on a variety of household concerns, with no strong causal links to demographic characteristics" (Kennedy and Macmillan, 1980, p.S-5).

fail the housing evaluation. Enrolled minorities in the Demand Experiment received payments at a lower rate than nonminorities because the former were more likely to occupy inadequate housing, and either could not or would not obtain better housing order to participate. If the shorter period of the Demand Experiment's offer was a deterrent to making the changes necessary to participate, then minorities would be more often affected because their dwellings failed more frequently. In the Supply Experiment it was also true that minority enrollees were more likely to encounter difficulty in meeting the standards: Among renter enrollees, 74 percent of nonminority households eventually received payments versus 68 percent of black households.* However, the ratio of 1.09 is much smaller than the Demand Experiment figure (1.43), because of differences in either the standards or the incentives provided by the length of the payment period.

A study that examined the effect of turnover on participation rates in the housing allowance program was conducted by Rydell, Mulford, and Kozimor (1981). Their conclusion was that the overall participation rate for a welfare program like the Supply Experiment can never reach 100 percent because of the nature of turnover in eligibility; newly eligible households do not apply immediately for assistance, nor do they receive an allowance at once. In the meantime other households are terminating eligibility, predominantly because of changes in income. Using crude data for the first three program years in each site, Rydell et al. estimated turnover parameters and used them to predict a steady state in which 50 percent of all eligibles would be currently enrolled. Turnover does partly explain participation rates; but the model for it is incomplete in that it does not account for the likelihood that some eligibles may choose not to participate. A significant proportion of those who were eligible for a long time, such as elderly singles and single parents, never enrolled.

This paper has described what participation rates are in the Experimental Housing Allowance Program: Roughly 40 percent of eligible households are enrolled, roughly one-third are receiving an allowance.

---

*Figures are for St. Joseph County only.

Our current research* is aimed at determining why participation rates are so low by relating household decisions to the benefits of and impediments to participation. Phyllis Ellickson (1981) began this task with her behavioral model, which shows that lack of information about the housing allowance program was extremely important in the first program year. She also showed that both income and housing costs played a role in determining whether or not a household joined the allowance program. James Wendt (1981) has extended the analysis of the decision to apply to later program years when program knowledge approached steady state in the two communities. Sinclair Coleman's research (1981) shows how the housing standards interacted with household characteristics to determine whether enrollees would receive an allowance. An understanding of the reasons for the observed participation patterns should enable us to determine the extent to which the participation rates measured in the experimental program can be generalized to a national program.

---

*Reported in Carter and Wendt (1981).

Appendix A
VARIANCE COMPUTATIONS

POINT ESTIMATES

The variance of estimates of the number of eligible households in the population and its subgroups are calculated from a stochastic simulation. The assumptions of the simulation are:

1. P_{ikt} is defined as the probability that a household of type k in stratum i is eligible at wave t. The coefficients, $\tilde{b}$, of the logit equation for P_{ikt} have the normal distribution. The mean of the distribution, $\bar{b}$, is the maximum likelihood estimate (the equation in Tables 2.4 and 2.5), and the covariance matrix is estimated from the second-order derivatives of the likelihood function at $\bar{b}$.
2. θ_{ikt}, the expected proportion of households in stratum i at wave t who are of type k, has the posterior Dirichlet distribution with parameter vector $\tilde{\alpha}_{it}$.
3. Given $\tilde{\theta}_{it}$, the actual number of all unobserved households in stratum i at wave t who are of type k has the multinomial distribution with parameters $\tilde{\theta}_{it}$ and $N_{it} - n_{it}$ where N_{it} is the population of stratum i at wave t and n_{it} is the sample size of stratum i at wave t.

The algorithm consists of independent replications of the following steps. First a random vector $\tilde{b}$ is drawn from the multivariate normal for the logit coefficients and P_{ikt} given $\tilde{b}$ is calculated for each population group, stratum and wave. Let Y_{ikt} be the number of unobserved eligible households of type k in stratum i at wave t. The following equations are evaluated:

$$E(Y_{ikt}|\tilde{b}) = \frac{\alpha_{ikt}}{\tau} P_{ikt}(N_{it} - n_{it}), \qquad \text{(A.1)}$$

$$E(Y_{ikt}^{2}|\tilde{b}) = \frac{\alpha_{ikt}(\alpha_{ikt}+1)}{\tau(\tau+1)} P_{ikt}^{2}(N_{it}-n_{it})(N_{it}-n_{it}-1) + \frac{\alpha_{ikt}}{\tau} P_{ikt}(N_{it}-n_{it}) \qquad (A.2)$$

and

$$E(Y_{ikt}Y_{ijt}|\tilde{b}) = \frac{\alpha_{ikt}\alpha_{ijt}}{\tau(\tau+1)} \cdot P_{ikt}P_{ijt}(N_{it}-n_{it})(N_{it}-n_{it}-1). \qquad (A.3)$$

For each relevant group, S, (e.g., one life-cycle stage, owners, minorities, etc.) we then calculate the expected number of unobserved eligibles and its square:

$$E(V_{St}|b) = \sum_{i}\sum_{k\varepsilon S} E(Y_{ikt}|b) \qquad (A.4)$$

and

$$E(V_{St}^{2}|b) = \left(\sum_{i}\sum_{k\varepsilon S} E(Y_{ikt}|b)\right)^{2} \qquad (A.5)$$

Equation (A.5) is calculated from equations (A.2), (A.3), and the fact that assumptions 2 and 3 imply $Y_{ikt}|b$ is independent of $Y_{ujt}|b$ for $i \neq u$.

Each replication provides one observation of V_{St} and its square. The average over replications is used to provide unconditional estimates of $E(V_{St})$ and $E(V_{St}^{2})$ for calculating the standard deviation of V_{St}, which is also the standard deviation of the population estimate.

The simulation was run for a total of 300 replications per site and wave. We analyzed the simulation output in groups of 50 replications each, and determined that the coefficient of variation of the estimated standard deviation was roughly 0.02 for renters, 0.04 for owners and less than 0.04 for each life-cycle group.

Table A.1 presents the standard deviation of the estimated number of eligibles in 1977 by life-cycle stage, tenure, and minority status from the simulation. Bounds on participation rates have been calculated from these standard deviations as follows. If $\hat{V}_S$ is the estimated

Table A.1

STANDARD DEVIATIONS FOR ESTIMATED NUMBER OF ELIGIBLE HOUSEHOLDS AND ONE-STANDARD-DEVIATION RANGES FOR ENROLLMENT AND PARTICIPATION RATES, BY SELECTED HOUSEHOLD CHARACTERISTICS: YEAR 3, BROWN AND ST. JOSEPH COUNTIES

Household Type	Brown County			St. Joseph County		
	Standard Deviation For Estimate of Eligibles	One-Standard Deviation Range[a]		Standard Deviation For Estimate of Eligibles	One-Standard-Deviation Range[a]	
		For Percent Enrolled	For Percent Receiving Payments		For Percent Enrolled	For Percent Receiving Payments
Tenure:						
Renter	149	53-57	43-47	178	53-57	37-40
Owner	436	28-35	26-31	1,058	28-34	25-30
Life-cycle stage:						
Elderly couple	198	22-29	20-27	418	22-31	21-29
Elderly single	215	52-66	50-63	611	35-44	33-42
Single parent	176	54-64	43-51	413	51-63	35-44
Young couple, young children	179	25-31	19-23	269	18-24	10-14
Other	137	28-36	22-29	219	25-33	19-25
Race of head:						
Minority	(b)	(b)	(b)	331	47-59	32-40
White non-Hispanic	(b)	(b)	(b)	985	32-37	27-32
Total	501	40-45	34-38	1,091	35-41	29-33

SOURCE: Estimated by authors from the wave 4 household survey in Brown County, the wave 3 household survey in St. Joseph County, and the year 3 HAO administrative files in each site.

[a]Ranges are obtained by dividing the number of households enrolled or receiving payments by the number of eligible households plus one standard deviation and then by the same number less one standard deviation.

[b]Not estimated for Brown County, where nearly all residents are non-Hispanic whites.

number of eligibles in group S, $\hat{\sigma}$ is the standard deviation of $\hat{V}_S$, and W_S is the number of participants (i.e., either enrolled or receiving payments), then the point that is one standard deviation below the estimated participation rate is $\frac{W_S}{\hat{V}_S + \hat{\sigma}}$, and the upper bound on the one-standard deviation range is

$$\frac{W_S}{\hat{V}_S - \hat{\sigma}}.$$

CHANGE ESTIMATES

Because we survey some of the same households repeatedly, we should be able to measure changes in size and composition of the eligible pool more accurately than we can estimate the size and composition themselves at a single point in time. Our logit model of eligibility did not use the information on the changes in eligibility status of individual households. However, our estimates of $\tilde{\theta}_{it}$ are based on repeated sampling of many of the same households. Since our estimate of $\tilde{\theta}_{it}$ is positively related to our estimate of $\tilde{\theta}_{ir}$ for wave r, the correlation of $\tilde{\theta}_{it}$ over waves results in correlated eligibility estimates. In calculating variances we have assumed that $\tilde{\theta}_{it}$ is independent of $\tilde{\theta}_{ir}$. Thus the estimate of the standard deviation of the change in the number of eligibles is conservative.

Appendix B

FOURTH YEAR PARTICIPATION STATISTICS

In the tables presented in this appendix we have used the estimates of eligible households from the 1977 surveys. The numbers of households who were enrolled and participating at the end of the fourth program year are shown by life-cycle stage and tenure for Brown County in Table B.1, and for St. Joseph County in Table B.2. The enrollment statistics are approximately 15 months later in Brown County, and 21 months later in St. Joseph County, than the date for the estimated number of eligible households. For consistency, we have removed non-elderly singles from client counts as well as from the estimates of eligible households, as we have done throughout this document.

Table B.1

BROWN COUNTY PARTICIPATION RATES AT END OF FOURTH PROGRAM YEAR, BY LIFE-CYCLE STAGE AND TENURE

Life-Cycle Stage	Number of Households			Participation Rates	
	Eligible	Enrolled	Receiving Payments	Enrolled	Receiving Payments
Renters					
Elderly couple	266	97	89	36	33
Elderly single	771	605	583	78	76
Single parent	1,296	1,018	809	79	62
Young couple, young children	845	250	169	30	20
Other	511	152	125	30	24
Total	3,689	2,122	1,775	58	48
Owners					
Elderly couple	1,087	211	209	19	19
Elderly single	1,007	483	473	48	47
Single parent	677	224	188	33	28
Young couple, young children	858	112	91	13	11
Other	594	114	99	19	17
Total	4,223	1,144	1,060	27	25
Total					
Elderly couple	1,353	308	298	23	22
Elderly single	1,778	1,088	1,056	61	59
Single parent	1,973	1,242	997	63	51
Young couple, young children	1,703	362	260	21	15
Other	1,105	266	224	24	20
Total	7,912	3,266	2,835	41	36

SOURCE: Estimates of eligibles from wave 4 household survey (conducted mainly during the first half of 1977) are based on the methodology described in Sec. II. Counts of enrolled clients and those receiving payments are from HAO administrative records for June 1978.

NOTE: Nonelderly single-person households are excluded because most of them were not eligible to join the program until August 1977. Resident landlords and households living in government-assisted properties are excluded from eligibility estimates.

Table B.2

ST. JOSEPH COUNTY PARTICIPATION RATES AT END OF FOURTH PROGRAM YEAR, BY LIFE-CYCLE STAGE AND TENURE

Life-Cycle Stage	Number of Households			Participation Rates (%)	
	Eligible	Enrolled	Receiving Payments	Enrolled	Receiving Payments
Renters					
Elderly couple	277	72	68	26	25
Elderly single	1,271	640	583	50	46
Single parent	1,925	1,430	945	74	49
Young couple, young children	642	193	104	30	16
Other	529	169	122	32	23
Total	4,644	2,504	1,822	54	39
Owners					
Elderly couple	2,399	639	616	27	25
Elderly single	4,193	1,634	1,565	39	37
Single parent	1,876	609	513	32	27
Young couple, young children	1,236	93	68	8	6
Other	1,052	202	168	19	16
Total	10,755	3,177	2,930	30	27
Total					
Elderly couple	2,676	711	684	27	26
Elderly single	5,464	2,274	2,148	42	39
Single parent	3,801	2,039	1,458	54	38
Young couple, young children	1,878	286	172	15	9
Other	1,580	371	290	23	18
Total	15,399	5,681	4,752	37	31

SOURCE: Estimates of eligibles from wave 3 household survey (conducted mainly during the first half of 1977) are based on the methodology described in Sec. II. Counts of enrolled clients and those receiving payments are from HAO administrative records for December 1978.

NOTE: Most nonelderly single-person households were not eligible to join the program until August 1977. Because they had such a short time to join, their participation rates would not represent steady state; therefore we excluded them. Resident landlords and households living in government-assisted properties are excluded from eligibility estimates.

Table B.3

ST. JOSEPH COUNTY PARTICIPATION RATES AT END OF FOURTH PROGRAM YEAR, BY RACE OF HEAD AND TENURE

Race of Head	Number of Households: Eligible	Number of Households: Enrolled	Number of Households: Receiving Payments	Participation Rates (%): Enrolled	Participation Rates (%): Receiving Payments
Renters					
Minority	1,433	867	543	61	38
White non-Hispanic	3,211	1,637	1,279	51	40
Total	4,644	2,504	1,822	54	39
Owners					
Minority	1,458	520	458	36	31
White non-Hispanic	9,297	2,657	2,472	29	27
Total	10,755	3,177	2,930	30	27
Total					
Minority	2,891	1,387	1,001	48	35
White non-Hispanic	12,508	4,294	3,751	34	30
Total	15,399	5,681	4,752	37	31

SOURCE: Estimates of eligibles from wave 3 household survey (conducted mainly during the first half of 1977) are based on the methodology described in Sec. II. Counts of enrolled clients and those receiving payments are from HAO administrative records for December 1978.

NOTE: Most nonelderly single-person households were not eligible to join the program until August 1977. Because they had such a short time to join, their participation rates would not represent steady state; therefore we excluded them. Resident landlords and households living in government-assisted properties are excluded from eligibility estimates.

Appendix C

THE STANDARD COST OF ADEQUATE HOUSING USED FOR ELIGIBILITY ESTIMATES

Table C.1

COMPARISON OF THE STANDARD COST OF ADEQUATE HOUSING (R*) AS SET BY THE HAOS AND AS ESTIMATED DURING EACH SURVEY WAVE, BY SIZE OF HOUSEHOLD: BROWN COUNTY

Number of Persons	Measured R^*			Estimated R^*			
	Sept. 1973	Jan. 1976	Jan. 1977	Wave 1	Wave 2	Wave 3	Wave 4
1	1,200	1,500	1,560	1,260	1,390	1,500	1,560
2	1,500	1,740	1,860	1,550	1,650	1,740	1,860
3-4	1,860	2,100	2,220	1,910	2,010	2,100	2,220
5-6	2,040	2,340	2,460	2,100	2,230	2,340	2,460
7-8	2,280	2,520	2,640	2,330	2,430	2,520	2,640
9+	2,640	2,760	2,940	2,665	2,715	2,760	2,940

SOURCE: FPOG policy memoranda 141 and 186 for HAO values. Interpolations by authors.

NOTE: The dates indicate when the HAO values were measured; they were implemented several months later. The interpolation points are approximately at the median survey dates.

Table C.2

COMPARISON OF THE STANDARD COST OF ADEQUATE HOUSING (R*) AS SET BY THE HAOS AND AS ESTIMATED DURING EACH SURVEY WAVE, BY SIZE OF HOUSEHOLD: ST. JOSEPH COUNTY

Number of Persons	Measured R^*			Estimated R^*		
	Aug. 1974	Jul. 1976	Aug. 1977	Wave 1	Wave 2	Wave 3
1	1,200	1,380	1,440	1,270	1,340	1,415
2	1,500	1,680	1,800	1,570	1,640	1,755
3-4	1,740	1,920	2,100	1,810	1,880	2,030
5-6	1,920	2,100	2,220	1,990	2,060	2,175
7+	2,040	2,220	2,280	2,110	2,180	2,255

SOURCE: FPOG policy memoranda 158 and 193 for HAO values. Interpolations by the authors.

NOTE: The dates indicate when the HAO values were measured; they were implemented several months later. The interpolation points are approximately at the median survey dates.

REFERENCES

Bendick, Mark, Jr., "Why Do Persons Eligible for Public Assistance Fail to Enroll?," The Urban Institute, Washington, D.C., W.P. 0819-02, August 1979.

Boren, Patricia, *Codebook for the Household Survey: Site I, Wave 4*, The Rand Corporation, N-1358-HUD, March 1980.

Bureau of Labor Statistics, *Employment and Earnings, State and Areas: 1939-1978*, U.S. Department of Labor, Washington, D.C., Bulletin 1370-13, 1979.

Carter, Grace, and John Rolph, *New York City Fire Alarm Prediction Models: 1. Box-Reported Serious Fires*, The Rand Corporation, R-1214-NYC, May 1973.

Carter, Grace M., and James C. Wendt, *Participation in an Open Enrollment Housing Allowance Program*, The Rand Corporation, R-2783-HUD, forthcoming.

Coleman, Sinclair, *How Enrollees Respond to Allowances and Housing Standards*, The Rand Corporation, R-2781-HUD, forthcoming.

Corcoran, Timothy, *Selecting the Permanent Panel of Residential Properties, Site I*, The Rand Corporation, N-1107-HUD, April 1981. (First issued as WN-9575-HUD, April 1978.)

Corcoran, Timothy, *Selecting the Permanent Panel for Residential Properties: Site II*, The Rand Corporation, N-1109-HUD, April 1981. (First issued as WN-9577-HUD, April 1977.)

Corcoran, Timothy, *Survey Sample Design for Site I*, The Rand Corporation, N-1065-HUD, February 1981. (First issued as WN-8640-HUD, March 1974.)

Efron, B., and Carl Morris, "Stein's Paradox in Mathematical Statistics," *Scientific American*, May 1978.

Ellickson, Phyllis, *Who Applies for Housing Allowances? Early Lessons from the Housing Assistance Supply Experiment*, The Rand Corporation, R-2632-HUD, forthcoming.

Employment and Training Administration, *Area Trends in Employment and Unemployment*, U.S. Department of Labor, Washington, D.C., April 1975.

Employment and Training Administration, *Area Trends in Employment and Unemployment*, U.S. Department of Labor, Washington, D.C., May, June, July 1977.

Fay, Robert E., III, and Roger Herriot, "Estimates of Income for Small Places: An Application of James-Stein Procedures to Census Data," *Journal of the American Statistical Association*, 74 (No. 366), June 1979.

HASE Staff, *Completing the Supply Experiment*, The Rand Corporation, N-1138-HUD, May 1981. (First issued as WN-10223-HUD, June 1978.)

Hosek, James R., *The AFDC-Unemployed Fathers Program and Welfare Reform*, The Rand Corporation, R-2471-HEW, August 1979.

Kennedy, Stephen D., and Jean MacMillan, *Participation Under Alternative Housing Allowance Programs: Evidence from the Housing Allowance Demand Experiment*, Abt Associates, Cambridge, Mass., June 1980.

Kingsley, G. Thomas, and Priscilla Schlegel, *Housing Allowances and Administrative Efficiency,* The Rand Corporation, N-1741-HUD, forthcoming.

Lowry, Ira S. (ed.), *Design of the Housing Assistance Supply Experiment*, The Rand Corporation, R-2630-HUD, June 1980.

Lowry, Ira S., *Inflation in the Standard Cost of Adequate Housing: Site I, 1973-1976*, The Rand Corporation, N-1102-HUD, October 1979. (First issued as WN-9430-HUD, March 1976.)

McDowell, James L., *Effects of the HAO Lead-Based Paint Standards*, The Rand Corporation, N-1306-HUD, June 1980.

McDowell, James L., *Housing Allowances and Housing Improvement: Early Findings*, The Rand Corporation, N-1198-HUD, September 1979.

Morris, Carl, and John E. Rolph, *Introduction to Data Analysis and Statistical Inference*, The Rand Corporation, P-5819, June 1978.

Rydell, C. Peter, John Mulford, and Lawrence Kozimor, *The Dynamics of Participation in a Housing Allowance Program*, The Rand Corporation, N-1137-HUD, February 1981. (First issued as WN-10200-HUD, June 1978.)

U.S. Bureau of the Census, *Statistical Abstract of the U.S.: 1978* (99th Ed.), Washington, D.C., September 1978.

Wang, Ann, and Leslie Geller, *Codebook for the HAO Client Characteristics File: Site I, Year 3*, The Rand Corporation, N-1146-HUD, May 1979.

Wendt, James, *Why Households Apply for Housing Allowances*, The Rand Corporation, R-2782-HUD, forthcoming.

RAND/R-2780-HUD