Attrition Bias in a Randomized Trial of Health Insurance

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ABSTRACT

We examined the rates of utilization of participants in a randomized study of health insurance to see if there were differences between those who normally completed the study, and those who departed prematurely. Using a negative binomial regression model, we found that the dropouts had significantly different rates of use, but that the magnitude of the difference was small.

KEY WORDS: Attrition; health insurance; cost-sharing; negative binomial.

1. INTRODUCTION

Decisions about the desirability of cost-sharing in health insurance have been based in part on observational studies of individuals with varying insurance coverage. The results from these studies were suspect because individuals could select their own coverage. To the extent that premiums did not accurately reflect individual risks, one would expect that sicker individuals would select better coverage. This adverse selection would imply that the estimates from observational studies could be biased to the extent that individuals differed in unmeasured ways.

To obtain unbiased estimates of the effects of cost-sharing on the use of health services and health status, the federal government sponsored a randomized trial of insurance coverage, The RAND Health Insurance Experiment. Results of that social experiment have been published, using data on the individuals who accepted the offer to participate and stayed in the study. See Newhouse et al. (1981; 1982), Keeler et al. (1982), Duan et al. (1983, 1984), Manning et al. (1984 a,b; 1985; 1987), Leibowitz et al. (1985) for use of health services; and Brook et al. (1983; 1984), Valdez et al. (1985), Bailit et al. (1985), Sloss et al. (1987), Keeler et al. (1987) and Ware et al. (1986) for health status. Some of these studies use some, but not all, of the data on the dropouts. Despite the use of randomization in the original assignment of families o insurance plans, there were two potential threats to the balance across the insurance plans: nonrandom refusal of the offer to participate, and nonrandom attrition from the study. Our primary defense against these threats was to offer substantial financial incentives to enroll and complete the study. Families were given a lump-sum payment equal to their worst-case financial risk associated with the plan; thus, no family was worse off financially for being in the study. The amount of the lump-sum payment was independent of use of health care services. The family's nonexperimental health insurance coverage was maintained for the family by the HIE during the experimental period with the benefits of the policy assigned to the HIE. If the family had no coverage, the HIE purchased a policy on their behalf. Thus, no family could become uninsurable as a result of its participation in the study.

Rates of refusals and sample loss were fairly low, but they were higher on the less generous insurance plans. Refusals varied from 6% on the most generous insurance plan (free care) to 23% on the least generous lan (for five of the sites reported in Brook et al., 1984). Four and one-half percent of the individuals with free care, 10.2% of those with inpatient and outpatient cost sharing, and 10.9% of those with only outpatient cost sharing (free inpatient care) dropped out of the study before the completion of the study; these data exclude the 0.8% who died during the study.

If the families originally assigned to a plan who later refused or dropped out were systematically different from those who stayed, the composition of the plan may change, unbalancing the study. We have used analysis of covariance (ANOCOVA) to adjust the results for the lack of balance in those covariates which are associated with non-completion of the study. Nevertheless, if the plan populations are not balanced on measured characteristics, they may also be unbalanced on unmeasured characteristics (for which we cannot adjust).

As an indirect test for a bias from refusal to join the study or subsequent attrition, we compared whether people completing the experiment are similar across plans, and compared those who refused to enroll or dropped out with those who completed (Brook et al. 1984). There were no statistically significant differences between those completing the experiment on different plans in any of 20 enrollment measures of demographic, health or prior use of health services. Those refusing to participate differed in age and (marginally) in income but not in health status or prior use of medical services. Nevertheless, this imbalance did not vary significantly by plan; that is, those refusing to participate were not differentially older or younger by lan. Dropouts were sicker on average at enrollment than those who stayed, but those quitting the study for health reasons were evenly divided across plans.

In this paper, we study whether those who dropped out differed from those who completed the experiment in terms of their observed use of medical care while they were still in the experiment. We compare the rates of use of both inpatient and outpatient services for those who dropped out while they were still in the experiment with the rates of those who remained. Assuming no anticipation, if the rates differ for the stayers and dropouts, after adjustment for measurable differences, then the selective attrition from the experiment could result in a bias in past estimates based largely on stayer samples. Unfortunately, we do not have data on use after people left the study.

The results presented below suggest that estimates based on the stayers alone are significantly, but not appreciably biased.

2. THE HEALTH INSURANCE EXPERIMENT

The data for this analysis are drawn from the Health Insurance Experiment (HIE). The HIE is a large randomized trial of the effects of different health insurance arrangements on the demand for health services and the health status of individuals, conducted between 1974 and 1982 (Newhouse 1974; Brook et al. 1979). The HIE enrolled families in six sites: Dayton, Ohio; Seattle, Washington; Fitchburg, Massachusetts; Franklin County, Massachusetts; Charleston, South Carolina; and Georgetown County, South Carolina. In each site, families enrolled for either three or five years. The sample is a random sample of each site's non-aged civilian population.

Most of the families participating in the experiment were assigned to different fee-for-service plans. Some were assigned to two prepaid group practice insurance plans.[1] We focus here on the fee-for-service plans. The plans can be divided into three major groups: (1) the free plan paid for all care (1843 individuals enrolled); (2) the family pay plans had different levels of cost sharing which varied over two dimensions: the coinsurance rate and an income-related upper limit on

[1]Use of medical services on the two prepaid group practice insurance plans are reported in Manning et al. (1984).

out-of-pocket family expenses (2634 enrolled); and (3) the individual deductible plan, which essentially had a $150 limit per person on outpatient out-of-pocket expenses and free inpatient services (1276 enrolled). The coinsurance rates (percentage paid out-of-pocket) were 0, 25, 50, or 95% for health services. Beyond the upper limit, the insurance plan reimbursed all expenses in full. All plans covered the same wide variety of services (Clasquin 1973).

Families were enrolled as a unit with only eligible members participating. No choice of plan was offered; the family could either accept the experimental plan or choose not to participate. Families were assigned to treatments using a balanced randomization based on the Finite Selection Model (Morris 1979).

The Sample. For this paper, we use data on all 5809 individuals initially enrolled in the study on the three groups of plans described above. We exclude newborns and children adopted during the course of the three to five years of the study. We also exclude the individuals who died on the study; death rates are not significantly different by insurance plan, and these individuals are known to differ appreciably in their use of services, especially during their last year of life.

The unit of observation is a person, because the major determinants of the use of health care services are individual (e.g., age, sex, and health status) rather than family (e.g., insurance coverage, and family income).

Dependent Variables. We examine both inpatient and outpatient care. Expenditures on inpatient care account for 59% of the total expenses for medical care. Our measure of inpatient use is the number of hospital admissions that an individual had during the study. An admission is any unique person-hospital encounter without a break in inpatient treatment. Our measure of outpatient care is the number of episodes of outpatient medical treatment. The episode includes all visits and drugs associated with the treatment of a particular bout of illness; see Keeler et al. (1982) for further details.

Independent Variables. We use three groups of independent variables: insurance plan variables, measures of health status, and sociodemographic and economic measures.

For inpatient care, we use indicator variables for the groups of insurance plans described above: family pay, and individual deductible; the free plan is the omitted group. These plan variables are interacted with an indicator variable for being a child (aged less than 18 at enrollment) or an adult. Earlier analysis of HIE data (Newhouse et al. 1981, 1982; Manning et al. 1987) indicated that there was no appreciable or significant differences in inpatient use among the family pay plans, but that there was a statistically significant and appreciable difference in the response of children and adults to cost-sharing.

For outpatient care, we use the square root of the coinsurance rate and an indicator variable for individual deductible plan; the free plan is the omitted group.

We use four measures of health status: (1) a scale for general health perceptions (Ware 1976; Davies and Ware 1981; Eisen et al. 1980); (2) the presence of physical or role limitations,indicators for any limitations due to poor health (based on scales reported in Stewart et al. 1977, 1978, 1981a, 1981b; Eisen et al. 1980); (3) chronic disease status that is a count of the number of 26 common diseases or problems (Manning, Newhouse, and Ware 1982); and (4) a scale for mental health status (Veit and Ware 1984; Ware et al. 1979, 1980; Williams et al. 1981; Eisen et al. 1980). Each of these measures is based on the self-administered Medical History Questionnaire for individuals 14 years or older (Ware et al. 1980; Brook et al. 1984). Measures for children are based on questionnaires filled out by parents (Eisen et al. 1980; Valdez et al., 1985). All of the health status data used here were collected at the beginning of the study.

The model used in our analysis also includes covariates for age, sex, race, family income, and family size. With the exception of family income, the data were collected before or at enrollment in the study. Family income data are from 1975 in Dayton, 1978 for the three-year group in South Carolina, and 1976 for all other participants.

3. NEGATIVE BINOMIAL REGRESSION MODEL

We use a regression model based on the negative binomial distribution to estimate the response of both admissions and of outpatient episodes of treatment to cost sharing. The negative binomial distribution is an appealing model for medical use because it can yield a large proportion of zeros and a skewed distribution of positive use. It is also appealing because of its convolution properties with respect to time observed. It is more flexible than a Poisson regression because it allows for unmeasured characteristics to generate overdispersion; the negative binomial model can be represented as a Poisson model with random effects.

The negative binomial model can be formulated as a mixture of Poisson variates. Let the ith individual's admissions (Y_i) be drawn independently from a Poisson distribution with rate λ_i:

$$p(Y_i = n | \lambda_i, T_i) = (\lambda_i T_i)^n \exp(-\lambda_i T_i)/n! \qquad \textbf{(1)}$$

where T_i is the period of time observed for individual i.

Assume that the rate λ_i for different individuals follow a gamma distribution with shape parameter α, and scale parameter β:

$$f_{\alpha,\beta}(\lambda) = [\beta^{\alpha}\Gamma(\alpha)]^{-1}\lambda^{\alpha-1}\exp(-\lambda/\beta), \qquad \textbf{(2)}$$

It follows then that the observed number of admissions follow a negative binomial distribution where

$$rob\ (admits = n) = \begin{pmatrix} \alpha + n - 1 \\ \alpha - 1 \end{pmatrix} \left(\frac{\beta}{T_i + \beta}\right)^{n} \left(\frac{T_i}{T_i + \beta}\right) \qquad \textbf{(3)}$$

(Johnson and Kotz 1969, pp. 122-142).

Model Specification. In the results below, we assume that the parameters α and β are linear combinations of observed individual characteristics. For admissions, the log of the parameter α is specified as a linear combination of indicators for plan, for being a child, and interactions among these indicators. The log of the parameter β is a linear combination of all characteristics mentioned above:

$$\boldsymbol{ln(\beta) = -x_i \ast} \qquad \textbf{(4)}$$

where x_i is row vector of given individual characteristics, including an intercept, and * is a column vector of parameters to be estimated; the plan and child/adult interactions with plan are included here as well.

For outpatient episodes, we specify plan by two variables -- the square root of the coinsurance rate, and an indicator variable for the individual deductible plan. The free plan (zero coinsurance) is the reference group. Both insurance variables are interacted with family income, adjusted for family size. The specification of the model differs from that for inpatient care in that the *l*n(") term is a function of plan, income, gender, and physical limitation.

Fit on Stayers. The model is estimated by maximum likelihood on the stayer population. To test goodness of fit, we use the estimated parameter values to predict the probability that each stayer will have no admissions, one admission, etc., using the formula above for the pdf. These predictions were summed over individuals, and contrasted with the observed pdf for admissions using a P^2 statistic. We followed the same process for episodes of outpatient treatment.

These " P^2 " statistics are not necessarily distributed as a P^2, because we tested the goodness of fit on the data set used to fit the parameters. As a result, we would expect the model to fit better than if the parameters were known a priori or fitted on another data set. To test how important this problem was, we conducted a single split sample analysis for the overall goodness-of-fit test for inpatient admissions. We gave each person a random number from a uniform distribution, and assigned each individual with value of 0.5 or more to the estimation sample. Using parameter esimates from this training sample, we predicted the distribution of admissions for the other half sample, and calculated the P^2 statistic. The goodness-of-fit test statistic on the other half sample was $\chi^2(7) = 3.12$. Because the qualitative result was qualitatively the same (i.e., a relatively good fit), we proceeded with the uncorrected chi-square approach in the results below.

Test for Attrition Bias. To test for an attrition bias, we compare the actual hospitalization and outpatient rates for the dropouts before they dropped out with the predictions for the dropouts based on the negative binomial model fitted to the stayers. That is, we predict the distribution of admission or episodes of outpatient treatment for each dropout (the probability that each person would have no admissions, the probability of one admission, etc.). These predictions were summed over individuals, and compared with the actual distribution of admissions and episodes.

This approach is predicated on the assumption that the response to cost-sharing is stable over time. If there were transitory surges in utilization as a function of plan, this formulation could erroneously attribute such transitory effects to selection. A priori, we would expect that those facing better coverage on the experiment than they had before the experiment could have a surge in demand as they had health problems treated (e.g., vision examinations or discretionary admissions). Similarly, at the end of the experiment, individuals on generous plans might rush to have deferrable problems treated before the end of the study, when they would return to less generous non-experimental health insurance.

For inpatient care, this assumption is found to be empirically valid; we did not see any significant transitory effects. Using first differences in admissions for the stayer population (admit(t) - admit(t + 1)), we examined both the difference in the response to insurance plan between the first two and the last two years to capture transitory surges in use at the beginning and end of the study. We found no significant evidence of transitory behavior ($\chi^2(2) = 2.57$). For those who stayed five years, we also examined the middle years (years 2, 3, and 4) and found no significant differences ($\chi^2(2) = 0.80$). Although the number of outpatient episodes exhibits a small transitory surge during the first three months of the study (Keeler et al 1982), we have not corrected our estimates for this. Because most dropouts leave early in the experiment, this omission would give the appearance of higher use rates for dropouts when their true rate is equal. Thus, this omission increases our chance of observing an "attrition bias," because the dropouts could have higher rates of use due to the surge during their short stay on the study rather than due to any real selection effect.

4. RESULTS

Inpatient Use -- Stayer Sample. We estimated the rate of admissions for the subpopulation that stayed to the end of the study. Among the stayers, individuals facing cost sharing for both inpatient and outpatient care on the family pay plans had 20% fewer admissions than those with free medical care ($t = -3.04$, $p < .01$). For those with free inpatient care but outpatient cost sharing, the reduction was a statistically insignificant 11% ($t = -1.47$). See Table 1 for the results based on the parameter estimates from negative binomial regression. The model provides a reasonably good fit to the actual distribution of admissions over the three to five year period; Table 2 compares the actual and predicted density function; we have $\chi^2(7) = 3.73$.

There is similar agreement between predicted and observed if we break the sample by length of enrollment (five versus three years, $\chi^2(7) = 1.62$ and 5.60 respectively), or by insurance plan (free versus family pay plans versus individual deductible; $\chi^2(7) = 5.05$, 5.59, and 4.05 respectively).

This agreement could occur if our covariates explained little or none of the variation in outpatient or inpatient use. In such a case, it would be difficult to detect meaningful differences in rates of use between stayers and dropouts. Among stayers, we could explain 13.7 percent of the variation in inpatient use. Figure 1 plots actual versus predicted number of admissions; each point corresponds to a tenth of the sample, with averages by tenths ranked by the predicted scores. Between the lowest and highest tenths of the sample, based on predicted use, there was 12 fold difference in inpatient use.

Inpatient Use -- Attrition Bias. Using the estimates from the stayer population, we made forecasts to the individuals who died or dropped out. As expected, the deaths had a much higher rate of use, adjusting for observed differences, than those who survived the experiment (not shown). Dropouts on the free plan had 87.8% more admissions than predicted, while the corresponding numbers were 34.5% less on the family pay plans and 20.0% more on the individual deductible plan. If we look at the average residuals (actual admissions minus predicted admissions, based on the model's parameter estimates based on the stayer sample), we find that the free plan has more admissions than predicted (t = 1.76), the family pay plans too few (t = -2.61), and the individual deductible plan insignificantly too many (t = 0.59).

An alternative way of looking at the differences between dropouts and stayers is to estimate the negative binomial regression model on the pooled sample, including both stayers and dropouts, and a set of interactions between plan and being a dropout. The qualitative conclusion is unchanged: the free plan dropouts have a 69% higher mean admission rate than the stayers (t = 2.11), the family pay dropouts have a 33% lower admission rate (t = -1.96), and the individual deductible plan dropouts have a 16% higher admission rate (t = 0.52). If we pool the stayers and the dropouts and include interaction terms between plan and being a dropout, we can estimate the differences among the plans as if the dropouts had stayed in the study. Table 3 presents the mean annual admission rates based on stayers only and based on dropouts pooled with stayers rates. These results indicate that excluding the dropouts understated the response to plan. The magnitude of the understatement is small for the individual deductible plan. With the attrition cases included, use admission rates are 11 percent lower than on the free plan. With the attrition cases excluded, the admission rate is 10 percent lower than on the free plan. The magnitude of the understatement is moderate for the family pay plan. With the attrition cases included, use admission rates are 25 percent lower than on the free plan. With the attrition cases excluded the admission rate is 19 percent lower than on the free plan. In both cases, the difference is due to attrition bias on the free plan, not on the family pay or individual deductible plans.

The dropouts other than the deaths also had a different distribution of admissions than predicted, after adjusting for observed characteristics and time on study; $P^2(7) = 23.14$. The dropouts had a greater likelihood of no admissions or of a large number of admissions; see Table 4.

The misfit was even more pronounced by plan. The dropouts on the free plan had fewer cases without hospitalization and more with a large number of hospitalizations; $P^2(7) = 44.83$. The family pay plan and individual deductible plans' predictions fit the observed distribution reasonably well; $P^2(7) = 3.75$, $P^2(7) = 6.58$, respectively. We get a similar picture if we group the outcomes into three categories: zero, one, and two plus admissions. For all plans, $P^2(2)$ is 1.69, while it is 4.25 for the free plan, 3.21 for the family pay plan, and 2.70 for the individual deductible. Although this alternative test solves the empty cell problem in the P^2 calculation, it does not recognize the differences between predicted and observed in the right tail.

Outpatient Use -- Stayer Sample. We have conducted a similar analysis for the use of outpatient care; covering the period of the first three years of the study. Among the stayers, individuals facing a coinsurance rate of 95% for outpatient and inpatient care, subject to an upper limit on out-of-pocket expenses, had about one third fewer episodes of treatment per year as those with free care ($t = -11.57$, $p < 0.0001$). For those with cost sharing for outpatient care, but free inpatient care (i.e., the Individual Deductible plan), there were 30% fewer episodes of outpatient treatment than with free care ($t = -9.66$, $p < 0.0001$). See Table 5 for the results based on the stayer population parameter estimates.

The model for outpatient care does not provide as good a fit to the actual distribution of the outpatient episodes as we obtained with inpatient admissions. Table 6 compares the actual and predicted distributions; $P^2(8) = 16.76$. The actual distribution has more cases with no use, and fewer with over 30 episodes over three years than the estimated model predicts. Among stayers, we could explain 30.1 percent of the variation in outpatient use. Figure 2 plots actual versus predicted number of outpatient episodes; each point corresponds to a tenth of the sample, with averages by tenths ranked by the predicted scores. Between the lowest and highest tenths of the sample, based on predicted use, there was a six-fold difference in outpatient use.

Outpatient Use -- Attrition Bias. Using the parameter estimates from the stayer population, we have made forecasts to the population that dropped out, other than the deaths. The dropouts had fewer episodes than predicted. The dropouts on the free plan had 5% fewer outpatient episodes than predicted based on the stayer population. The family pay plan participants had 20% fewer episodes than predicted, while the Individual Deductible plan participants were 33% lower.

An alternative way of looking at the differences between dropouts and stayers is to include an indicator for being a dropout, and a set of interactions between the plan variables and being a dropout. The Wald test statistic for a differential bias by plan is $P^2(2) = 4.27$, where the test is whether the estimated plan/dropout interaction terms coefficients are significantly different from zero. The Wald test for any bias (all plans considered) is $P^2(3) = 11.32$, where the test is for the main effect of dropping out, and the interactions with the plan variables. Both test results are driven by a much smaller outpatient episode rate for the individual deductible plan dropouts than predicted based on the stayers. The free plan dropouts are not significantly different than the free stayers ($t = -0.31$), nor is there a systematic pattern for the family pay plans ($t = -0.87$).

If we pool the stayers and the dropouts and include interactions between the plan variables and being a dropout, we can estimate the differences among the plans as if the dropouts had stayed in the study. Table 7 presents the mean annual episode rates based on stayers only and on dropouts and stayers combined. Excluding the dropouts understated the response to plan, but the magnitude of the understatement is quite small.

The dropouts were more likely to have no episodes, and less likely to have a large number, than if they had behaved as the stayers did. As Table 8 indicates, this pattern can be seen in all of the insurance plans.

5. DISCUSSION

The data from the Health Insurance Experiment indicate that individuals who dropped out behaved differently (in ways that measured characteristics and time on the study could not explain) than those who stayed to the end of the study. Thus, past inferences based largely on stayers are biased. For outpatient care, the bias was quite small. But for inpatient care, there was a moderate bias which understated the effect of family pay plans with catastrophic coverage.

The direction of the bias surprised us. We expected that high users on the pay plans would drop out so that they could use their more generous nonexperimental coverage. To avoid this, the HIE made a worst case payment so that no family was financially worse off for being in the study. We expected no bias on the free plan, because its coverage and breadth of benefits was much better than most existing private or public insurance packages (except for the limit of 52 psychotherapy visits per person per year, which very few people reached; see Wells et al. 1982). Instead, we found that it was high users of inpatient services on the free plan and low users of both inpatient and outpatient care on the family pay plans who dropped out.

If anything, then, our previously reported estimates of plan effects appear understated. However, the quantitative extent of the understatement is modest. The small overall bias is due not so much to the absence of a difference in use rates between stayers and dropouts, but more to the small proportion of dropouts.

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Table 1

PREDICTED ANNUAL ADMISSIONS RATES/1000 PERSONS
BASED ON STAYER POPULATION

Plan	Mean (std. err.)	t
Free	122.9 (6.2)	--
Family Pay	98.7 (4.9)	-3.04
Individual deductible	108.9 (7.2)	-1.47

NOTE: Parameter estimates and predictions are for the population present at enrollment who completed the study normally. t is for the contrast with the free plan rate.

Table 2

PREDICTED AND OBSERVED ADMISSIONS
FOR STAYERS

Number of Admits	Predicted Percentage	Observed Percentage
0	75.45	75.43
1	16.38	16.33
2	4.91	5.08
3	1.78	1.84
4	0.73	0.64
5	0.34	0.28
6-10	0.37	0.32
11+	0.04	0.08

NOTE: Predictions based on model fitted to stayers only. N = 5278.

Table 3

ALTERNATIVE PREDICTIONS OF ANNUAL ADMISSION RATES

Plan	Based on Stayers		Based on Non-Deaths	
	Mean (std.err.)	t vs. Free	Mean (std.err.)	t vs. Free
Free	0.124 (.00630)	--	0.129 (.00651)	--
Family Pay	0.101 (.00510)	-2.85	0.097 (.00475)	-3.93
Individual Deductible	0.112 (.00747)	-1.22	0.115 (.00807)	-1.35

NOTE: Predictions based on parameter estimates fitted to the population given at the top of the column, but predicted to the enrollment population. Column 1 differs from corresponding column of Table 2 in that prediction population is all entrants, not just stayers.

Table 4

PREDICTED AND OBSERVED ADMISSIONS FOR DROPOUTS

	All Plans		Free Plan		Family Pay		Ind. Ded.	
Number of Admits	Pred. %	Obs. %	Pred. %	Obs. %	Pred. %	Obs. %	Pred. %	Obs. %
0	86.86	88.19	83.25	80.95	87.83	90.33	87.16	88.41
1	9.83	8.15	11.91	9.52	9.31	9.55	9.56	6.52
2	2.17	2.04	2.96	3.57	1.96	1.12	1.20	2.90
3	0.66	0.61	1.02	1.19	0.56	0.00	0.64	1.45
4	0.25	0.20	0.42	1.19	0.20	0.00	0.25	0.00
5	0.11	0.00	0.20	0.00	0.08	0.00	0.12	0.00
6-10	0.11	0.81	0.22	3.57	0.06	0.00	0.16	0.72
11+	0.01	0.00	0.02	0.00	0.00	0.00	0.02	0.00
N	490		84		269		138	

NOTE: Predictions based on model fitted to stayers only, adjusted for time on the study.

Table 5

PREDICTED ANNUAL EPISODE RATES

Plan	Mean (std.err.)	t vs. Free
Free	3.66 (.0695)	--
25% Coinsurance*	2.91 (.0522)	-10.48
50% Coinsurance*	2.86 (.0736)	-8.00
95% Coinsurance*	2.33 (.0793)	-11.57
Individual Deductible	2.55 (.0909)	-9.66

NOTE: Parameter estimates and predictions are for the population present at enrollment and who completed the study normally.

*Subject to upper limit on out-of-pocket payment of at most $1,000 per year.

Table 6

PREDICTED AND OBSERVED OUTPATIENT EPISODES
OF TREATMENT FOR STAYERS

Number of Episodes	Predicted Percentage	Observed Percentage
0	7.23	7.71
1-3	21.68	21.77
4-6	19.75	18.34
7-9	15.46	15.02
10-12	11.21	11.31
13-15	7.82	8.36
16-18	5.35	5.67
19-30	8.88	9.61
31+	2.60	2.22

NOTE: Predictions based on model estimated for stayers only, adjusted for time on the study.

Table 7

ALTERNATIVE PREDICTIONS OF ANNUAL OUTPATIENT EPISODES

Plan	Based on Stayers		Based on Non-Deaths	
	Mean (std.err.)	t vs. Free	Mean (std.err.)	t vs. Free
Free	3.65 (.070)	--	3.66 (.070)	--
25%	2.89 (.052)	-10.55	2.88 (.051)	-10.86
50%	2.88 (.076)	-7.60	2.85 (.073)	-8.11
95%	2.31 (.079)	-11.68	2.26 (.075)	-12.51
Individual Deductible	2.52 (.089)	-9.98	2.44 (.085)	-10.96

NOTE: Predictions based on parameter estimates fitted to the population given at the top of the column, but predicted to the enrollment population. Column 1 differs from corresponding column of Table 6 in that prediction population is all entrants, not just stayers.

Table 8

PREDICTED AND OBSERVED EPISODES FOR DROPOUTS

	All Plans		Free Plan		Family Plan		Ind. Ded.	
Number of Episodes	Pred. %	Obs. %	Pred. %	Obs. %	Pred. %	Obs. %	Pred. %	Obs. %
0	29.96	39.48	14.36	24.00	30.91	39.37	36.98	48.48
1-3	36.29	32.75	40.38	25.33	38.02	33.07	36.19	36.36
4-6	15.57	13.45	21.08	18.67	15.38	15.75	12.82	6.06
7-9	7.90	5.64	13.25	13.33	7.22	4.33	6.18	3.79
10-12	4.25	4.56	7.91	8.00	3.68	4.72	3.27	2.27
13-15	2.37	1.30	4.69	1.33	1.98	1.18	1.81	1.51
16-18	1.37	0.87	2.83	1.33	1.11	0.79	1.04	0.76
19-30	1.85	1.74	4.12	6.67	1.42	0.79	1.40	0.76
31+	0.43	0.22	1.19	1.33	0.28	0.00	0.31	0.00
N	461		75		354		132	

NOTE: Predictions based on model estimated for stayers only, adjusted adjusted for time on the study.

APPENDIX

Table A.1

INDEPENDENT VARIABLES

INDICATOR VARIABLES (If not otherwise defined, variables equal 1 if right hand side condition holds, otherwise zero)

Insurance Plan [a]	
IDP	Individual Deductible
PAY	Family Coinsurance Plans
Other Subexperiments	
TERM3	Three year enrollment [b]
TOOKPHYS	Entry physical examination [c]
NOHR	Did not file health diary [d]
WKLY	Filed health diary weekly [d]
Year and Site [e]	
SEA	Enrolled in Seattle
FIT	Enrolled in Fitchburg
FRA	Enrolled in Franklin Co., Mass.
CHA	Enrolled in Charleston
GEO	Enrolled in Georgetown Co., S.C.
Other variables	
BLACK	Black
CHILD	Age < 18 on first day of year
FEMALE	Female
MISINC	Income data missing
PHYSLM	Physically or role limited
GHINMIS	General health measure missing
FLAG1	Health data from infant form (ages 0 - 4+) [f]
FLAG2	Health data from pediatric form [f]
Continuous Variables	
MA1	(1 - FEMALE)*(AGE - 30)
MA2	(1 - FEMALE)*(LN(AGE) - LN(30))
FA1	FEMALE * (AGE - 30)
FA2	FEMALE * (AGE - 30)**2
FA3	FEMALE * (AGE - 30)**3
FA4	FEMALE * (AGE - 30)**4

(cont.)

Table A.1 (cont.)

FA5	FEMALE * (AGE - 30)**5
XXGHI	Predicted part of General Health Index [g]
XXGHI2	Predicted part of (General Health Index squared) [g]
GHINDX	GHINNM * (Residual part of General Health Index) [g]
MHI	Mental Health Index [h]
LINC	log family income [j]
LINC2	LINC squared
XAFDC,AFDC	Received Aid to Families with Dependent Children
LFAM	Log of family size at enrollment
Interactions	
ADIDP	IDP * (1 - CHILD)
ADPAY	PAY * (1 - CHILD)
CHIDP	IDP * CHILD
CHPAY	PAY * CHILD

[a] The free care plan (no out-of-pocket cost) is the omitted group.

[b] The five year enrollment group is the omitted group.

[c] The no entry examination group is the omitted group.

[d] The group that filed health diaries biweekly is the omitted group.

[e] Dayton is the omitted group.

[f] The adult (ages 14+) version of the health questionnaire is the omitted group.

[g] Due to a missing general health index for all Dayton participants at enrollment, we imputed missing value replacements for all participants. For those with GHINNM = 1, we regressed the general health index (or its square) on socio-economic and demographic variables, and pre-experimental health as measured by Excellent/Good/Fair/Poor, Pain, and Worry, which were available in all sites. We used the predicted part for everyone, and the residuals for the non-Dayton sites as our general health measure. A prefix of x indicates the predicted part; the absence of a prefix of x indicates the residual part (for those with complete data, 0 otherwise).

[h] Not available for the infant form (ages 0-4+) of the health questionnaire.

Table A.2

NEGATIVE BINOMIAL REGRESSION ESTIMATES
FOR HOSPITALIZATION

PARAMETER LN(1/BETA)

VARIABLE	COEFFICIENT	SD(COEFF)	T
INTERCEP	-3.6037E+00	2.966E+00	-1.21
SEA	1.3612E-01	2.186E-01	0.62
FIT	3.2871E-01	2.189E-01	1.50
FRA	2.9139E-02	2.128E-01	0.14
CHA	1.8408E-01	2.271E-01	0.81
GEO	2.8880E-01	2.198E-01	1.31
CHILD	-4.6042E-01	5.261E-01	-0.88
ADPAY	2.7895E-01	2.721E-01	1.03
ADIDP	-6.0610E-01	3.495E-01	-1.73
CHPAY	5.2182E-02	5.875E-01	0.09
CHIDP	1.0920E+00	5.976E-01	1.83
NOHR	-8.6769E-02	1.108E-01	-0.78
WEEKLY	-2.3994E-01	1.387E-01	-1.73
TERM3	-2.8819E-03	7.548E-02	-0.04
TOOKPHYS	5.6135E-02	6.531E-02	0.86
BLACK	-2.2149E-01	1.311E-01	-1.69
MA1	2.2519E-02	6.378E-03	3.53
MA2	-4.5952E-01	1.586E-01	-2.90
FA1	-4.1136E-02	9.333E-03	-4.41
FA2	-2.2448E-03	6.208E-04	-3.62
FA3	1.6824E-04	4.711E-05	3.57
FA4	2.0716E-06	8.160E-07	2.54
FA5	-1.4983E-07	4.806E-08	-3.12
FEMALE	7.6918E-01	9.310E-02	8.26
LINC	1.2212E+00	6.570E-01	1.86
LINC2	-6.9826E-02	4.000E-02	-1.75
LFAM	1.4995E-02	7.214E-02	0.21
AFDC	9.3001E-02	1.342E-01	0.69
GHINDX	-8.5210E-03	2.353E-03	-3.62
FLAG1	-1.2544E-01	2.944E-01	-0.43
FLAG2	-2.2611E-01	1.924E-01	-1.18
MISINC	1.6737E-01	1.075E-01	1.56
GHINMIS	-2.9970E-01	2.444E-01	-1.23
XXGHI	-8.7966E-02	3.787E-02	-2.32
PHYSLM	3.6730E-01	8.158E-02	4.50
MHI	2.8851E-03	2.468E-03	1.17
XXGHI2	4.5120E-04	3.037E-04	1.49

(cont.)

Table A.2 (cont.)

PARAMETER LN(ALPHA)

VARIABLE	COEFFICIENT	SD(COEFF)	T
INTERCEP	7.8045E-03	2.048E-01	0.04
CHILD	-1.3208E-01	4.813E-01	-0.27
ADPAY	-4.7282E-01	2.438E-01	-1.94
ADIDP	4.2045E-01	3.287E-01	1.28
CHPAY	-1.2123E-01	5.508E-01	-0.22
CHIDP	-9.5653E-01	5.400E-01	-1.77

APPENDIX II

OUTPATIENT CARE

We have conducted a similar analysis for the use of outpatient care. In that case, the dependent variable is the number of episodes of outpatient medical treatment. The episode includes all visits and drugs associated with the treatment of a particular bout of illness; see Keeler et al (1982) for further details.

Methods. The data cover the period of the first three years of the study. In the case of the group enrolled for five years in South Carolina, we have the last three years; attrition during the first two years in this group is ignored here.

The specification of insurance coverage differs from that for inpatient care. We specify plan by two variables -- the square root of the coinsurance rate, and an indicator variable for the individual deductible plan. The free plan (zero coinsurance) is the reference group. Both insurance variables are interacted with family income, adjusted for family size.

The specification of the model differs from that for inpatient care in that the ln(") term is a function of plan, income, gender, and health status. Although the number of outpatient episodes exhibits a transitory surge during the first three months of the study (Keeler et al, 1982), we have not corrected our estimates for this.

Results. As with inpatient care, we estimated the rate of outpatient episodes for the subpopulation that stayed to the end of the study. The negative binomial regression estimates can be found in Appendix I, Table A.3. The model for outpatient care does not provide as good a fit to the actual distribution of the outpatient episodes as we obtained with inpatient admissions. Table 5 compares the actual and predicted distributions; the $\chi^2(8) = 16.76$. The actual distribution has more cases with no use, and fewer with over 30 episodes over three years than the estimated model predicts. The same phenomenon occurs if we limit the explanatory variables to insurance plan variables.

Among the stayers, individuals facing a coinsurance rate of 95% for outpatient and inpatient care, subject to an upper limit on out-of-pocket expenses, had about one third fewer episodes of treatment per year as those with free care ($t = -11.57$, $p < .0001$). For those with cost sharing for outpatient care, but free inpatient care (i.e., the Individual Deductible plan), there were 30% fewer episodes of outpatient treatment than with free care ($t = -9.66$, $p < .0001$). See Table 6 for the results based on the stayer population parameter estimates.

Using the parameter estimates from the stayer population, we have made forecasts to the population that dropped out, other than the deaths. The dropouts had fewer episodes than predicted. The dropouts were more likely to have no episodes, and less likely to have a large number, than if they had behaved as the stayers did. As Table 7 indicates, this pattern can be seen in all of the insurance plans. The dropouts on the free plan had 0.5% fewer outpatient episodes than predicted based on the stayer population. The family pay plan participants had 20% fewer episodes than predicted, while the Individual Deductible plan participants were 33% lower.

An alternative way of looking at the differences between dropouts and stayers is to include an indicator for being a dropout, and a set of interactions between the plan variables and being a dropout. The test statistic for a differential bias by plan is $P^2(2) = 4.27$, while that for any bias (all plans considered) is $P^2\ (3) = 11.32$. Both test results are driven by a much smaller outpatient episode rate for the individual deductible plan dropouts than predicted based on the stayers. The free plan dropouts are not significantly different than the free stayers ($t = -0.31$), nor is there a systematic pattern for the family pay plans ($t = -.87$ for drop*square root of coinsurance).

If we pool the stayers and the dropouts, we can estimate the differences among the plans if the dropouts had stayed in the study. Table 8 presents the mean annual episode rates based on stayer only and on stayer/dropout rates. These results indicate that excluding the dropouts understated the response to plan, but the magnitude of the understatement is quite small.