Demand for Outpatient Mental Health Services in a Heavily Insured Population: The Case of the Blue Cross and Blue Shield Association's Federal Employees Health Benefits Program

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This article presents the results of a study of the impact of an increase in coinsurance on the demand for outpatient mental health services. The study population was a set of fully employed subscribers enrolled in the Blue Cross and Blue Shield Association's Federal Employees Health Benefits Program at some time during the period 1979 through 1981. A two-part model was used to examine the determinants of both the probability of mental health service use and the level of use. Our results indicate little price sensitivity in either part of the model, but substantial and significant income elasticities. Our results concerning the role of various sociodemographic and environmental variables are also reported.

As public and private third-party payers become increasingly cost-conscious, the impact of the structure of health benefits on consumer demand for health services is receiving considerable attention. The literature on the effects of certain financial incentives such as coinsurance and deductibles on the use of general health services has increased

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rapidly in recent years [1]. Much more limited, however, is the empirical evidence on these issues in the area of mental health services. The need for an empirical understanding of the impact of benefit changes on the cost and utilization of mental health services and of how these effects might differ from those in the general medical care setting is growing as the benefit structure for mental health services continues to change at a rapid pace.

One carrier of mental health services that has undergone many changes in its benefit structure is the Blue Cross and Blue Shield Association. This article presents results from our ongoing study of the impact of changes in the mental health benefit provided to federal workers and their families in the Blue Cross and Blue Shield Association's Federal Employees Health Benefits Plan (FEHBP) from 1979 through 1981. In 1979 and 1980, the high-option association plan had a 20 percent copayment on outpatient mental health benefits and a $100 deductible, while the low-option plan had a $200 deductible and a 25 percent copayment. The deductible applied to total health care expenditures, for both physical and mental health care. Only two members per family needed to satisfy the deductible. There was no copayment nor deductible for inpatient mental health services for either high-option or low-option coverage, although the low-option plan had a 90-day limit per confinement, with a copayment of 40 percent for supplemental coverage thereafter. In January 1981, in response to rising costs, the outpatient copayment on the high-option plan was increased to 30 percent and the deductible was increased from $100 to $150. For low-option coverage, the copayment was increased to 40 percent while the deductible remained at $200.

The empirical results in this paper are presented in the context of a cross-sectional demand model for outpatient psychiatric visits. The model is estimated for the three years in our study period 1979, 1980, and 1981. Following Wells et al. (hereafter referred to as RAND [2]), we estimate a two-part model. First, we relate the probability of using mental health services to a set of explanatory variables using a logistic regression model; and second, we examine the level of outpatient service use for users only. While our demand models include a variety of sociodemographic and geographic explanatory variables, we emphasize the two financial variables of particular policy relevance: price and income.

Our article is organized into five sections. The first section presents an overview of literature on the demand for mental health services with special emphasis on the work of RAND [2] and McGuire [3]. Next we discuss the theoretical context for the demand for mental health care. This is followed by a description of our data and the method of estimation. The fourth section contains our empirical results, and the final section presents conclusions and plans for future work.

LITERATURE OVERVIEW

Very few multivariate demand models have been estimated for mental health care. The two well-known studies in this area which examine price and income elasticities are RAND [2] and McGuire [3].

The RAND database is well known for its richness for demand research purposes [2]. The data are experimental and thus reduce the possibility of selection bias. The database has detailed information on utilization, and a variety of measures of mental as well as physical well being. A major limitation of the RAND study, however, is that the sample of users of mental health care is quite small. Under the most liberal definition of use, RAND finds that only 407 of its sample are users. An immediate practical result is that the estimates of population statistics have relatively large variances; thus, summary statistical descriptions of the population of users are somewhat imprecise.

Other limitations of the database include the exclusion of upper-income families and Medicare recipients from the experiment. Also, the RAND study was based on an insurance plan which excluded benefits for anyone using more than 52 visits annually to a mental health provider. This means that the sample excludes anyone undergoing psychoanalysis, and other heavy users.

Results from the RAND study suggest that, for those in the RAND experimental groups, the response to coinsurance for ambulatory mental health services is similar to that for ambulatory medical services. The price elasticities reported by RAND are less than 1.0 and vary considerably across the different plan levels' copayments and deductibles. The greatest degree of price responsiveness occurs with respect to probability of use. The RAND study reported insignificant income elasticities.

The McGuire study [3] is based on a survey of the last ten patients seeing a given psychiatrist in a day. This data set has very detailed information on those using outpatient services delivered by psychiatrists. It has many sociodemographic variables and detailed diagnostic information, as well as insurance and income measures. McGuire's study has the advantage that the mental health status of the user is represented fairly accurately by the psychiatrist. Further, McGuire's demand model is one of the most fully developed conceptual models in the mental health literature.

A limitation of McGuire's study is the fact that his results are from a nonrandom sample of the population. Data from the last ten visits to psychiatrists will overrepresent heavy users and provide no information on nonusers. Consequently, there is no way to analyze the probability of using mental health services because all of the information concerns those already in therapy. McGuire uses a dependent variable that is somewhat unusual, based as it is on the provider's projection of the patient's use rather than on actual use. This may lead to interpretive and statistical difficulties.

McGuire's study produced a number of important empirical results. He found that while demand increased with income, the increase was less than proportional (i.e., the income elasticity of demand was positive but less than 1.0). An interaction between the income variable and the price variable demonstrated that the positive response of demand to price falls as income levels fall. This suggests that higher-income individuals in the sample were more responsive to changes in the price of mental health services than lower-income individuals.

EMPIRICAL SPECIFICATION

The two-part model we estimate represents the individual's demand for outpatient mental health services. We assume that the price of these services is exogenous so that individuals are price takers. We focus, for this article, on the behavior of full-time, actively employed FEHBP Blue Cross and Blue Shield subscribers who were enrolled for at least 11 months of the study year in the high-option plan. This subsample provides the most homogeneous group of users and allows us to relate use to socioeconomic variables of interest while limiting the potentially confounding impact of many other factors.

PROBABILITY OF USE

The first stage of our empirical analysis focused on the impact of the coinsurance change on the probability that a subscriber became a user of mental health services. To address this question, we estimated a series of logistic regressions on a set of independent variables in each of our three study years (1979, 1980, 1981). The dependent variable, USE, took a value of one if the subscriber had any outpatient use in that year, and zero otherwise. The independent variables in our basic

Model 1 included dummy variables for sex (FEM), race (NON-WHITE), presence of a self-reported nonmental disability (DIS-ABLD), supervisory job category (BOSS), and presence of a second mental health service user in the family (DOTHIND); a continuous age variable (AGE), salary from 1980 federal employment deflated by the regional cost of living for the relevant year (SALARY), and a gross price variable (PHD). Gross prices in each year were estimated from our claims files by taking mean charges for an outpatient mental health visit to a physician within Health Care Financing Administration (HCFA) regions, and deflating by yearly regional cost of living indexes. A description of the deflation procedure is given in the appendix. Initially, we used mean charges by county as our price variable, but small sample problems led us to our final specification by HCFA region. The 222 HCFA regions are the areas used by Medicare in determining physician fees payable under Part B coverage. With the exception of Washington, DC, the sample of outpatient users is almost uniformly distributed across the HCFA regions.

The basic model incorporates the usual patient-specific socioeconomic variables. As noted below, DISABLD and DOTHIND are proxies for certain aspects of mental health status, and BOSS is included as a measure of non-money time costs (and perhaps mental health status as well). In Model 2, we added four county characteristics: 1980 population density in thousands (DENSITY), and a dummy variable equal to one if the county was designated as rural (RURAL). These variables were intended to capture characteristics of the local market environment which might affect demand (through time costs, mental health status, and McGuire's [3] "bandwagon" effect) and supply (through practice patterns and availability of complements and substitutes). Model 3 added the 1978 ratio of psychiatrists to all physicians (PSYCHMD) to the four county characteristics and the variables of the basic model to capture any remaining supply side effects.

The variables NEWEN80, NEWEN81, and NEW8081 in the pooled equation reflect various enrollment status conditions as defined in Figure 1, and are included to account for the manner in which our study sample was constructed and the self-selection effects of new enrollees. Details of the complicated sample construction are given in Jewell et al. [4, 5]. Briefly, we combined stratified independent samples from the enrollment files and each year's use file. The dummy variables NEWEN80, NEWEN81, and NEW8081 identify contracts newly enrolled in 1980 (NEWEN80), contracts newly enrolled in 1981 (NEWEN81), and contracts newly enrolled in 1980 and still enrolled in 1981. These variables attempt to control for the self-selection bias that would result if use in newly enrolled contracts was systematically different from use in continuing contracts. Although we are able in this way to adjust for the biased selection effects on new enrollees, we are still unable to adjust for the effects of individuals who drop out of the insurance plan between years. For example, if mental health users dropped from the high-option Blue Cross and Blue Shield plan in 1981 because of the benefit changes, a decline in mental health usage would be observed in 1981. Such effects are included in the size of the coefficient of the dummy variables D1980 and D1981, which also measure the year effects. This possibility should be taken into account when these variables are interpreted below. However, this type of dropout effect is unlikely, because the Blue Cross and Blue Shield plan continued to provide the most comprehensive coverage for mental health care.

Means and standard deviations for all variables are given in Table 1. In 1979, the value of the dependent variable USE was .278. In 1980, it was .248; and in 1981, it was .244. Variable definitions and data sources are listed in Figure 1.

The coefficients of the three separate cross sections allow us to draw inferences regarding those factors that affect the use of mental health care services; a comparison of the coefficients in the 1979 and 1980 regressions with those in the 1981 equations provides a test of the existence and nature of any structural changes in these relationships brought about by the coinsurance change or other intertemporal changes not otherwise accounted for.

Table 1: Variable Means

	1979		*1980*		*1981*	
	Mean	*S.D.*	*Mean*	*S.D.*	*Mean*	*S.D.*
LOUT	2.91	1.27	2.93	1.22	2.94	1.22
FEM	0.46	0.50	0.44	0.50	0.45	0.50
NONWHITE	0.12	0.32	0.11	0.31	0.08	0.28
AGE	41.09	10.06	40.68	10.11	40.65	10.25
SALARY	25.87	9.95	24.57	10.21	25.97	9.95
PHD	46.91	6.47	44.31	6.35	47.54	6.40
DISABLD	0.32	0.47	0.30	0.46	0.31	0.46
DOTHIND	0.23	0.42	0.27	0.45	0.27	0.44
BOSS	0.24	0.43	0.20	0.40	0.23	0.42
RURAL	0.01	0.11	0.02	0.14	0.03	0.16
PSYCHMD	0.08	0.04	0.08	0.04	0.08	0.04
RPCIN77	6.75	1.30	6.76	1.33	6.79	1.27
UNEMRATE	0.04	0.03	0.04	0.03	0.04	0.03
DENSITY	7.21	11.43	6.72	11.21	6.11	9.27

Figure 1: Variable Definitions

Variable Name	*Description*	*Source*
USE	= 1 if subscriber used stated type of mental health services in a calendar year.	Claims file
LOUT	The log of outpatient mental health care visits to psychiatrists, psychologists, mental health teams, or physicians in a calendar year.	Claims file
FEM	= 1 if female, 0 if male.	OPM file
NONWHITE	= 0 if white, 1 all other groups.	OPM file
AGE	Age of subscriber as of 1/1/80.	Enrollment file
SALARY	Annual deflated federal salary of subscriber, 1980, in thousands of dollars. (See appendix for construction of deflator.)	OPM file
PHD	Average price (deflated) of a physician/psychiatrist visit, by HCFA regions. (See appendix for construction of deflator.)	Claims file
DISABLD	= 1 if physical disability reported on employment record.	OPM file
DOTHIND	= 1 if there was a second user in the family.	Claims file
BOSS	= 1 if the subscriber holds a supervisory or managerial position.	OPM file
UNEMRATE	Percent county unemployment in 1980.	ARF file
RPCIN77	Real county per capita income in 1977.	ARF file
PSYCHMD	Number of psychiatrists in the county divided by the total number of physicians, 1978.	ARF file
D1980	= 1 if subscriber was sampled in 1980.	Enrollment file
D1981	= 1 if subscriber was sampled in 1981.	Enrollment file
NEWEN80	= 1 if subscriber newly enrolled in 1980.	Enrollment file
NEWEN81	= 1 if subscriber newly enrolled in 1981.	Enrollment file
NEW8081	= 1 if subscriber newly enrolled in 1980, still enrolled in 1981.	Enrollment file

We also estimated our three models on a pooled, cross-section, time series sample, which included users and nonusers for all three years. Here the impact of the coinsurance change is captured in the coefficient of the year dummy, D1981. An interaction term between the cross-sectional gross price measure and the year dummy was entered in the specification to test for differential effects of the increased coinsurance rate as a function of price. Its coefficient was everywhere insignificant and so was dropped from the analysis.

LEVEL OF USE

The second part of the empirical model analyzes the impact of the insurance changes on the number of services purchased by users of mental health services. We estimated this model using ordinary least-squares regression techniques for each year. The continuous dependent variable, LOUT, was measured as the logarithm of total outpatient mental health visits in that year by the subscriber to four types of providers (non-psychiatric physician, psychiatrist, psychologist, and mental health team). The independent variables in the basic model included, as before, the dummy variables FEM, NONWHITE, DISABLD, BOSS, and DOTHIND; and the continuous variables AGE, SALARY, and PHD. The second model again added the four county characteristics, UNEMRATE, RPCIN77, DENSITY, and RURAL. The third model added PSYCHMD.

As with the probability of use specification, we estimated separate yearly cross-sectional equations and a pooled cross-section time series equation including all three years' users and yearly dummy variables to capture the impact of the coinsurance change. Again we included a price/year interaction term, but it had no significant statistical effect and was dropped from our analysis.

UNIQUE CHARACTERISTICS OF THE DATA

Our data set and sample selection contain a number of unique characteristics that have important implications for our study, particularly as it compares to previous studies. First, all of the individuals in our study have identical (and deep) insurance coverage for both physical and mental health services during the study years. It is generally agreed that the Blue Cross and Blue Shield Association's FEHBP plan provided the most comprehensive mental health benefits at the time. Thus, many of the usual problems of comparisons across different combinations of benefit packages are eliminated. Second, a deductible change accompanied the coinsurance increase in 1981. The deductible increased from $100 in 1979 and 1980, to $150 in 1981. We have not tried in this study to modify our price variable to reflect the discontinuity in out-of-pocket prices in the deductible range. This means that for some low users (i.e., in the range of two to four visits), net price has been misspecified if the deductible has not otherwise been met. We believe the bias resulting from this omission to be small, because the deductible applies to all health care services (physical as well as mental)—especially since users of mental health care services tend to be higher-than-average users of physical health care services [6].

Third, our sample population for this article includes only subscribers, which means necessarily that all individuals are employed. Further, we have eliminated all but full-time employees. The latter fact has three implications for our study. First, it truncates our age distribution at both ends, eliminating both children and retired persons. Second, full-time employed individuals are less likely to be severely disabled. Thus, we also have a smaller range of physical and mental impairment than would be found in the general population. Finally, the criterion of full-time employment means that we have few very low-income individuals in our sample. Mean income (adjusted for regional variations in cost of living) is about $25,000 with a minimum of $7,800. Unlike the RAND study, however, we have no upper restriction on income. As a result, we are able to investigate the behavior of high-income people, who are likely to be relatively high users of outpatient mental services.

Fourth, because we have employment information only for the subscriber, we have an incomplete measure of family income. In particular, our measure is biased downward, especially for married female subscribers who are likely to have working spouses. The impact of this measurement problem is to bias our estimate of the coefficient on the income variable.

In spite of these limitations, our data set provides a unique opportunity to study the behavior over three years and a benefit change of a large, uniformly well insured population whose use accounts for a substantial share of mental health expenditures. Since the claims data are collected for purposes of bill payment, they are both detailed and accurate, allowing us to measure our dependent variable with more precision than is afforded in data sets collected by survey.

EMPIRICAL RESULTS

Since there were no significant structural differences in the estimated equations across years, we have reported only our pooled results here, in Tables 2 and 3.

SOCIODEMOGRAPHIC VARIABLES

In our insured population, the three sociodemographic variables, FEM, NONWHITE, and AGE, are predictors of both the probability of use of outpatient mental health services, and the amount of use. As

Table 2: Probability of Use Pooled All Three Years

Equation Number	*(1)*	*(2)*	*(3)*
Dependent Variable: Use/nonuse			
Independent Variables:			
INTERCEPT*	-4.05†	-4.60†	-4.81†
	(.38)‡	(.49)	(.49)
FEM	0.98†	0.91†	0.91†
	(.10)	(.11)	(.11)
NONWHITE	-0.92†	-1.04†	-1.05†
	(.13)	(.13)	(.13)
AGE	-0.03†	-0.03†	-0.03†
	(.004)	(.004)	(.004)
SALARY	0.04†	0.03†	0.03†
	(.006)	(.007)	(.007)
PHD	-0.0005	0.01	0.01
	(.007)	(.007)	(.007)
DISABLED	0.33†	0.30†	0.29†
	(.10)	(.10)	(.11)
DOTHIND	5.02†	5.02†	5.00†
	(.59)	(.59)	(.59)
BOSS	-0.38†	-0.41†	-0.41†
	(.13)	(.14)	(.14)
RURAL		-0.57†	-0.51
		(.29)	(.29)
UNEMRATE		-7.38†	-2.92
		(1.70)	(2.29)
RPCIN77		0.05	-0.03
		(.04)	(.05)
DENSITY		0.03†	0.02†
		(.006)	(.007)
PSYCHMD			6.73†
			(2.33)
D1980*	0.21	0.27†	0.27†
	(.12)	(.12)	(.12)
D1981*	0.22†	0.24†	0.24†
	(.11)	(.12)	(.12)
NEWEN80*	-0.96†	-0.99†	-1.02†
	(.43)	(.43)	(.43)
NEWEN81*	-0.24	-0.23	-0.24
	(.27)	(.27)	(.27)
NEW8081*	0.05	0.0001	0.0006
	(.33)	(.34)	(.34)
	χ^2 = 791.75	χ^2 = 848.92	χ^2 = 855.34
	N = 2,766	*N* = 2,766	*N* = 2,766

*Coefficient adjusted to reflect stratified sample design. See [4] for details.
†Significant at 5 percent level.
‡Asymptotic standard errors in parentheses.

Table 3: Outpatient Visits, Pooled All Three Years

Equation Number	*(1)*	*(2)*	*(3)*
Dependent Variable: Use/nonuse			
Independent Variables:			
INTERCEPT	4.15*	2.83*	2.73*
	(.28)†	(.34)	(.34)
FEM	0.27*	0.24*	0.24*
	(.07)	(.07)	(.07)
NONWHITE	-0.20*	-0.29*	-0.29*
	(.11)	(.10)	(.10)
AGE	-0.03*	-0.03*	-0.03*
	(.003)	(.003)	(.003)
SALARY	0.03*	0.03*	0.03*
	(.004)	(.004)	(.004)
PHD	-0.02	-0.01	-0.01
	(.005)	(.005)	(.005)
DISABLD	0.06	0.03	0.03
	(.07)	(.069)	(.069)
DOTHIND	0.22*	0.25*	0.24*
	(.079)	(.076)	(.077)
BOSS	-0.20*	-0.23*	-0.22*
	(.09)	(.088)	(.088)
D1980	-0.008	0.02	0.02
	(.08)	(.078)	(.078)
D1981	-0.004	0.01	0.01
	(.079)	(.076)	(.076)
RURAL		-0.15	-0.15
		(.22)	(.22)
UNEMRATE		-2.00	0.79
		(1.07)	(1.56)
RPCIN77		0.12*	0.07*
		(.026)	(.032)
DENSITY		0.02*	0.02*
		(.003)	(.004)
PSYCHMD			4.04*
			(1.66)
	$\bar{R}^2$ = .112	$\bar{R}^2$ = .166	$\bar{R}^2$ = .169
	F = 17.46	F = 19.55	F = 18.72
	N = 1,308	N = 1,308	N = 1,308

*Significant at 5 percent level.
†Standard errors in parentheses.

reported in the RAND study [2], females are more likely to use services than males. However, while RAND, McGuire, and Taube [2,3,7] report no relationship between gender and level of use, our results indicate that not only are females more likely to use services than are males, user females purchase roughly 7 to 8 percent (2.5 visits per year) more outpatient visits than male users. This is an important finding given that our sample includes only females who are working full-time and have insurance coverage equivalent to that of males in the sample. Thus, the differences cannot be attributed to employment status and/or insurance coverage between males and females.

AGE is negatively related to both probability of use and level of use, a consistently statistically significant finding. In an earlier model not shown, we entered the square of age to capture the nonlinearities in the age/level of use relationship reported in the literature [8], but its impact was uniformly insignificant. This finding is likely to be due to the absence of children is our sample (by virtue of our full-time-employed criterion), since it is over the very young age ranges that use appears to increase with age. In an earlier paper [9] in which we included use of the dependents of our subscribers, our bivariate tests indicated that the overall effect of age took on the reported inverted "J" shape.

As in McGuire's study [3], our results for the race variable NONWHITE suggest that employed whites are more likely to use, and that white users actually use more outpatient mental health services than employed nonwhites (roughly 6–8 percent more). The coefficient of BOSS is negative and significant for both dependent variables. This result is consistent either with the hypothesis that supervisory personnel tend to have fewer mental health problems (holding income levels constant) or that consuming mental health services is relatively more costly for these individuals when the value of time lost from work is counted.

MENTAL HEALTH STATUS

Both RAND [2] and McGuire [3] report the important influence of measures of mental health status on demand for mental health care. Unfortunately, we have no direct measure of this variable. As a weak proxy, we have entered DISABLD, a dummy variable indicating that the subscriber reported one of a list of physical handicaps to the Office of Personnel Management upon employment, and DOTHIND to indicate other family use of mental health services. The impact of DISABLD on probability of use is positive and significant, but it performs poorly in the level of use equations, where its coefficient is reasonably unstable across years and never significant. DOTHIND is consistently related positively to probability of use and level of use, suggesting that mental health problems tend to have (or be caused by) family effects. In earlier models, we attempted to use a set of county-wide measures of the "environment" (e.g., divorce rate, suicide rate, murder rate) as alternative proxies, but with little success. Clearly, mental health status indicators measured at the individual level are most appropriate.

REGIONAL VARIABLES AND THE BANDWAGON EFFECT

The county variables, RURAL, UNEMRATE, RPCIN77, and DENSITY are entered to capture geographic differences in practice style (relevant primarily in the level of use equations) and, perhaps, acceptance of mental health treatment (a gross view of McGuire's [3] bandwagon effect). They may also capture some of the "environmental" factors noted above that influence demand.

The impact of DENSITY on outpatient use is the most stable and significant of these factors. The coefficient of this variable is positive and significant across all models and all years for both parts of our demand model. This result could reflect something of the negative "environmental" effect on mental health status noted above, or a negative (and thus demand-increasing) impact on travel costs. The coefficient of RPCIN77 is less stable. Never significant in the probability of use equations, it is positive and generally significant in the level of use formulation. Per capita income is among the bandwagon measures suggested by McGuire [3]. It performs weakly here, but in the expected direction.

The inclusion of the supply variable, PSYCHMD, was intended to capture a number of demand-influencing factors. First, a higher ratio of psychiatrists to all physicians may reflect (or create) a greater acceptance of mental health treatment. Further, as PSYCHMD increases, more people are likely to seek care from mental health specialists rather than from general medical providers. This has two implications for our analysis. First, to the extent that treatment patterns differ between specialty mental health providers and their general medical counterparts, PSYCHMD should capture this effect in the level of use equations. Second, our claims data include only those claim codes with mental health diagnoses. For a variety of reasons, general medical providers are less likely to use diagnosis codes in this range for patients with nonacute mental health problems, particularly when the patient is seeking assistance with physical conditions simultaneously. Thus, particularly in counties with low values of PSYCHMD, we may have measurement problems with our dependent variables due to the data sampling procedure.

We had several reasons for selecting the number of psychiatrists as a percentage of all physicians rather than the per capita specification frequently used in demand studies to capture supply effects. First, the number of psychiatrists per capita alone confounds the sector-specific effects we are trying to capture with factors that influence the number of physicians generally. That is, if one area has more psychiatrists per capita than another area, this may only reflect the fact that this particular area is an attractive location in which to practice any type of medicine. Thus, the area may have high per capita figures for all physicians, including psychiatrists.

Further, although our supply variable is measured in a year prior to our study years, any remaining simultaneity problems with respect to the price variable are likely to be more severe with a per capita measure than with PSYCHMD. Since we are estimating individual rather than market demand curves, we feel that the likelihood of simultaneity problems is small.

As expected, the coefficient of PSYCHMD is positive, significant, and relatively stable over all years for both probability of use and level of use. The elasticity of the level of use with respect to PSYCHMD, calculated at the mean, is .43. Thus, outpatient visits increase as the percentage of physicians who are psychiatrists increases—but not proportionally.

INCOME

The impact of income on the demand for mental health services is of particular interest in this study since we are dealing with individuals who are well insured with identical coverage. In previous studies, such as McGuire [3], it has been difficult to identify the positive impact of income apart from its influence on the selection of insurance coverage. In the RAND study [2], this issue does not arise because individuals were randomly assigned to differing insurance plans; but the income distribution is truncated at higher levels by the experimental design. This may in part account for the fact that RAND estimates of the income coefficient are never statistically significant.

As shown in Tables 2 and 3, we found consistently positive and significant income effects among high-option users on both probability of use and level of use. Thus, not only are high-income individuals more likely to use outpatient mental health services; they also purchase more services than lower-income individuals who also use.

The estimated real-income elasticity (which measures the percentage change in outpatient visits resulting from a given percentage change in income), calculated at mean income levels for the level of use, is 0.75.[1] The elasticity is calculated by multiplying the income coefficient estimate in Table 3 by the sample mean income: -(.03)*25 = 0.75. Therefore, a 10 percent increase in income is associated with a 7.5 percent increase in outpatient use. Our results thus indicate that higher-income individuals use more outpatient mental health services than lower-income individuals even within a heavily insured population. Further, as noted in a previous section, our estimates of this coefficient are biased toward zero, since we have measured only the subscriber's contribution to family income.

PRICE AND THE IMPACT OF INSURANCE CHANGES

As outlined in the previous section, we capture the effect of price on mental health service demand in two ways. The impact of cross-sectional variations in gross price is measured by the coefficient of PHD in Table 2 for probability of use, and Table 3 for level of use. The figures in these tables indicate little cross-sectional price sensitivity over this range with respect to either probability or level of use. The low level of significance achieved in the basic equation for level of use disappears as soon as areawide characteristics are entered. Cross-sectional gross price is never a significant factor in the probability of use equations.

The change in the coinsurance rate in 1981 increased net (i.e., out-of-pocket) prices to consumers at given levels of gross price. In our specification, the impact of this change is measured in the cross-sectional equations (not shown) by intercept changes from 1980 to 1981, and in the pooled sample equations (Tables 2 and 3) by the coefficient of the 1981 year dummy variable, D1981. Again, price changes appear to have little impact on consumer demand for outpatient mental health services over the range for which we have observations. The Chow tests we performed to measure structural differences in the equations across years indicated no significant differences. For the probability of use equation, both the dummy variables D1980 and D1981 were marginally significant (Table 2), indicating higher rates of use in both of these years compared to 1979. However, the coefficients indicate no difference in use rates between 1980 and 1981. One interpretation is that the underlying increase in use of outpatient mental health services noted from 1979 to 1980 was arrested in 1981. This interpretation must be treated cautiously, however, since it relies heavily on the observed change from 1979 to 1980 as indicative of a general trend. There was no significant difference in level of use among users over the three years (Table 3). As noted above, interaction terms combining the impact of cross-sectional and over time price changes were tried in earlier specifications, but their coefficients were always insignificant.

Demand insensitivity to price in the probability of use equations is particularly interesting given the RAND [2] results, which indicate a negative relationship between price and probability of use. A number of possibilities for these results exist. The first relates to measurement problems associated with the cross-sectional price measure, PHD. PHD was calculated from our claims data by taking average fees for outpatient physicians visits over HCFA regions and deflating. While we omitted regions in which there were fewer than 30 claims, we may still have an imprecise measure of undeflated cross-sectional price. Our coefficient estimates using observations from regions with large numbers of claims were significantly improved over those using all observations, lending support to the measurement error explanation of our weak results. In an attempt to increase our precision, we also estimated our model using county-specific price measures and prices estimated from fee regressions. None of these approaches changed our results. Further, although we have used the most reliable and disaggregated cost of living deflator readily available, it may also be measured with error. If this is the case, it biases the coefficient of the price variable toward zero.

Variations in quality across visits for which we cannot adjust may provide another partial explanation. However, we feel this is less of a problem in our work than elsewhere given that we used fees for a relatively homogeneous service (a one-hour visit to a psychiatrist) to construct our variable.

Additionally, we have not accounted for the presence of the deductible nor for its increase in 1981. We believe the bias resulting from this omission is small, but this is a testable proposition on which we will focus future work.

Price sensitivity may also be small due to the limited variation in out-of-pocket price observed in this study. Based on our calculations from the claims data, most of the cross-sectional net out-of-pocket price variation is within $3.00 per visit in each of the study years. The change in coinsurance caused mean net price to increase by only $5.40 per visit, reflecting the fact that real gross prices were constant or actually falling in some areas between 1980 and 1981, perhaps in response to the coinsurance change. While market prices should have been unaffected by the FEHBP benefit change, psychiatrists aware of the change may have altered fees charged to FEHBP patients in therapy to minimize its impact. Given that PHD was constructed using only fees charged to Blue Cross and Blue Shield Association FEHBP patients, it would capture this effect. Further, mental health providers may have altered their collection procedures for the patients' copayment in 1981 in order to minimize the impact of the price change on their visit volume.

Finally, we have restricted our focus in this study to changes in outpatient use as a result of the benefit change. However, it is possible that federal workers responded by altering their inpatient use (inpatient services remained fully covered throughout the period) or switching health care plans (e.g., from high option to low option; Blue Cross and Blue Shield to Aetna, etc.). The interaction with inpatient use may be of particular importance in comparing our results to those of RAND, since the RAND experiment offered less generous coverage of hospital services than did the Blue Cross and Blue Shield Association FEHBP plan [2].

PRICE INTERACTIONS

To investigate the possibility that the effect of price could be observed through its impact on other independent variables, we interacted price in the pooled sample with the female indicator, FEM, and (following McGuire [3]) with two income category dummies, SAL30 for subscribers with federal incomes between $18,000 and $30,000, and SAL50 for those with federal incomes greater than $30,000. The results for FEM were insignificant, indicating that working males and females exhibit a similar response to cross-sectional price variations with respect to level of outpatient use. The results for the income category variables are given in Table 4. The coefficients of the interaction term PSAL30 are negative and significant. Thus, price elasticity for outpatient mental health services tends to be higher for middle-income than for low-income individuals in our sample, perhaps due to better access to information on prices and quality or value of service. For the high-income category, price elasticity is insignificantly different from that for the low-income segment of the population. The latter finding may relate to higher search costs or greater likelihood of entering less flexible treatment regimens (e.g., psychoanalysis).

SIMULATIONS OF THE OVERALL IMPACT ON OUTPATIENT SERVICES

The overall impact of the change in insurance coverage on use of outpatient services by the entire population (i.e., users and nonusers) can be viewed through the simulated predictions in Table 5. Probability of use and mean use for three types of individuals were calculated

Table 4: Outpatient Visits, Pooled All Three Years Interactions

Equation Number	*(1)*	*(2)*
Dependent Variable: Log (outpatient visits)		
Independent Variables:		
INTERCEPT	3.05*	2.28*
	(.401)†	(.464)
FEM	-0.43	0.22*
	(.44)	(.073)
NONWHITE	-0.29*	-0.31*
	(.105)	(.106)
AGE	-0.03*	-0.03*
	(.003)	(.003)
PHD	-0.01	0.009
	(.005)	(.009)
DISABLD	0.02	0.029
	(.057)	(.069)
DOTHIND	0.24*	0.25*
	(.076)	(.073)
BOSS	-0.22*	-0.12
	(.088)	(.090)
RURAL	-0.16	-0.13
	(.216)	(.228)
UNEMRATE	0.83	0.45
	(1.56)	(1.60)
RPCIN77	0.07*	0.07*
	(.033)	(.032)
DENSITY	0.02*	0.02*
	(.004)	(.005)
PSYCHMD	4.11*	4.22*
	(1.66)	(1.66)
D1980	0.02	0.03
	(.083)	(.083)
D1981	0.008	0.02
	(.072)	(.076)
SAL30		1.36*
		(.519)
SAL50		1.36*
		(.632)
PSAL30		-0.02*
		(.01)
PSAL50		-0.01
		(.007)
PFEM		0.01
		(.006)
	$\bar{R}^2$ = .17	$\bar{R}^2$ = 16
	F = 17.71	F = 15.24
	N = 1,308	N = 1,308

*Significant at 5 percent level.

†Standard errors in parentheses.

Table 5: Predicted Effects of Coverage Changes

	1979	*1980*	*1981*
Person A			
(1) Probability of use	.0010	.0009	.0010
(2) Mean use	2.88	4.90	5.79
(3) Expected use rate = (1) × (2)	.003	.004	.006
Person B			
(1) Probability of use	.0139	.0197	.0247
(2) Mean use	14.92	19.36	10.83
(3) Expected use rate = (1) × (2)	.208	.381	.267
Person C			
(1) Probability of use	.0217	.0303	.0388
(2) Mean use	21.05	21.64	17.21
(3) Expected use rate = (1) × (2)	.456	.656	.669

—Mean values of other variables are assumed.
—Equation 3 coefficients are employed.
—These calculations are based on the antilog of visits. Since the coefficients of our regressions were estimated using the log of the dependent variables, these calculations will be biased [2].

A: 41-year-old nonwhite male in rural area, earning $15,000.
B: 41-year-old white male in urban area, earning $23,000.
C: 41-year-old white male in urban area, earning $30,000.

from Equation 3 coefficients for each of the study years. These figures were then combined to produce expected use rates for the three years for individuals with the specified characteristics (and mean values of the other variables) in the FEHBP population. The characteristics of the three types of individuals represented in Table 5 were chosen to reflect those of a very low user group (column A), a very high user group (column C), and the "mean" user group (column B).

As the figures demonstrate, the low-income nonwhite male of mean age (41 years) located in a low-density area represents the very low user group. Here, increased mean use over time combines with a nearly constant probability of use to produce a modest increase, from .003 to .006 in expected use from 1979 to 1981. The 1981 increase in coinsurance appears to have had little impact on use trends for this very low user group. The very high user group is represented in column C by a white male of mean age living in a relatively high-density area with a high ratio of psychiatrists to total physicians, earning $30,000. For this individual, the increase in probability of use over time is offset by a decrease in mean use from 1980 to 1981, so that the expected use rate rises very little from .66 in 1980 to .67 in 1981.

It is only for the "mean" user group (column B), that expected use appears to be dampened in 1981, the year of the copayment change. The "mean" user is characterized as a white male of mean age (41) and mean annual income ($23,000) living in an urban area of moderately high density. While probability of use for this type of individual steadily increased over the study period, mean use in 1981 fell substantially to below 1979 levels. The result is a decrease in expected use from .381 in 1980 to .267 in 1981.

Thus, it is clear that the overall impact of the benefit change is determined by the impact of the change on both probability of use and mean use. Further, these effects vary in direction and magnitude according to the socioeconomic characteristics of the population.

CONCLUSION

This article has examined the impact of a coinsurance change on the use of outpatient mental health services by fully employed Blue Cross and Blue Shield Association subscribers in the FEHBP from 1979 to 1981. The coinsurance change under study was relatively small in magnitude and was only observed for one year. Our findings are an important addition to the small econometric literature in this area because our results are not confounded by differential insurance coverage within our sample. Further, our results confirm that the two-part model of demand employed initially by RAND is a useful approach in the mental health area. Our estimate for both probability of use and level of use are robust and stable across study years.

Our study results suggest that small changes in out-of-pocket price do not trigger significant alterations in mental health service use, either with respect to probability of use or level of use. Given the magnitude of the price change, this result is not unexpected for a heavily insured population. Further, the increase in net price caused by the coinsurance change may have been mitigated by changes in provider collection practices for the self-pay portion of the bill which we could not observe. Thus, the differences between our findings and those of other researchers may not be contradictory once account is taken of the different study settings.

Our results with respect to income elasticities are significant. While RAND, Horgan, and Taube [2, 7, 10] find no significant relationship between income and use of services, we find a stable and positive relationship between income and demand, both for probability of use and level of use of outpatient mental health services. This is of particular consequence given the uniformity of insurance coverage.

Thus, even among individuals with identical insurance coverage, increases in income are associated with increased demand for outpatient mental health visits.

Our results indicate that middle-income subscribers demonstrate significantly greater price sensitivity than low-income subscribers, but that demand responsiveness does not differ between high-income and low-income groups. This finding has important policy implications as it suggests that middle-income individuals benefit most from policies that promote lower outpatient prices.

We found that females are more likely to use outpatient mental health services than males. Additionally, we found that female users purchase more services than male users. Since our sample is restricted to fully employed subscribers with identical mental health benefits, these differences cannot relate to employment status or insurance coverage.

Our simulations combining the two parts of the demand model suggest that future work needs to address the differential response to benefit changes by different segments of the population. Further work would also include an investigation of the impact of price changes on the substitution of inpatient care (which remained fully covered during the study period) for outpatient care. Finally, we will also examine the role of the benefit change on choice of plan (i.e., high option versus low option).

APPENDIX

ESTIMATION OF THE COST OF LIVING DEFLATOR

From the Bureau of Labor Statistics (BLS), we obtained 1979 annual budget cost indexes for high-income families for 24 metropolitan areas throughout the country. We omitted Honolulu, Hawaii and Anchorage, Alaska from the set and regressed the cost of living index (COL79) for the remaining 22 cities per capita income of the SMSA (PCINC); density in thousands (DENS); population in thousands (POP); and three regional dummy variables: North Central (NC), South (S), and West (W). The resulting equation was:

$$COL79 = \underset{(5.66)}{80.21} + \underset{(1.64)}{.004\ PCINC} + \underset{(4.09)}{6.00\ DENS} + \underset{(-2.07)}{.002\ POP}$$

$$- \underset{(-1.77)}{5.16\ NC} - \underset{(-2.67)}{8.60\ S} - \underset{(-1.55)}{5.86\ W}$$

$\bar{R}^2 = .640$
$F = 7.509$
(t-values in parentheses)

We then used these coefficients to estimate a cost of living index for all U.S. metropolitan areas for 1979. These were inflated over time by the Bureau of Labor Statistics Consumer Price Index for four city size-codes within each of four regions.

The region-specific nonmetropolitan area indexes from the BLS SMSA budget tables were used for rural counties, and were inflated over time using the CPI figures for the smallest city size (less than 75,000) for each region.

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