COINSURANCE AND THE DEMAND FOR MEDICAL SERVICES

A REPORT PREPARED UNDER GRANTS FROM THE OFFICE
OF ECONOMIC OPPORTUNITY AND THE NATIONAL CENTER
FOR HEALTH SERVICES RESEARCH AND DEVELOPMENT

CHARLES E. PHELPS
JOSEPH P. NEWHOUSE

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OCTOBER 1974

Rand
SANTA MONICA, CA. 90406

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PREFACE

One question that is not well understood in the financing of medical services is the effect of coinsurance on the consumer's demands for services. Some maintain coinsurance imposes barriers to care, while others argue that coinsurance is irrelevant to choice, since the physician makes decisions concerning resource allocation. This report treats the problem of coinsurance both theoretically and empirically, although the major thrust is empirical. Based on an expansion of the standard economic theory of consumer behavior, it shows how consumers would respond to changes in the coinsurance rate for various medical services. The report attempts to pull together the empirical evidence available and adds some from new sources. All the evidence indicates that coinsurance is not irrelevant to choice; irrespective of who makes the decisions, greater coinsurance implies less use of services. The amount of the reduction varies by type of service.

This revision updates a previous version of this report (R-964, April 1973). Several additional studies of the demand for dental services have been added, as well as another study of demand for prescription drugs and results from a study in which travel time for physician services was changed. The theoretical section has been generalized to include insurance policies with an upper limit, and a number of minor changes have been made in Sec. III, the review of the literature, to reflect new results.

This report should be of interest to those attempting to understand the impact of national health insurance on demand for medical care services and to those interested in demand for medical care in general. It is part of a considerably larger study that attempts, among other things, to ascertain the effects of alternative medical insurance plans on utilization of medical services. The research reported herein was supported by grants from the Office of Economic Opportunity and the National Center for Health Services Research and Development. A condensed version of this report was published in *The Review of Economics and Statistics*, Vol. 56, No. 3, August 1974.

SUMMARY

Using a formal model of utility maximization, this report develops a model of demand for medical services when reimbursement insurance is present, and when there are time costs involved in purchasing medical care. A major theoretical finding is that responsiveness to changes in insurance coverage diminishes as the cost of time becomes a relatively more important (and money price a relatively less important) proportion of total costs for medical services. It is also shown that, under plausible (and weak) conditions, the observed responsiveness of demand for medical care to either money price changes or insurance coverage changes approaches zero as insurance coverage becomes complete.

A number of data sources are used to estimate elasticities of demand for medical care services, and these results are compared with results of three other studies of the same problem in economics journals. Our sources show that the arc-elasticity for all medical services in the zero to 25 percent coinsurance range is on the order of 0.1. In percentage terms, we estimate that around 8 to 17 percent more services would be demanded at a zero coinsurance rate than at a 25 percent rate. This estimate is based on inferences from the premiums for various policies, as well as on a summation across demand changes for individual services. The zero to 25 percent range of coinsurance is relevant in assessing differences among many proposed universal health insurance plans, but these elasticities may not apply to other coinsurance ranges. In particular, these estimates may be inappropriate in assessing the effects of a shift from no insurance to complete or nearly complete coverage.

In the zero to 25 percent range of coverage, we estimate the elasticity of demand for hospital services to be 0.08; for medical office visits, 0.14; for ambulatory ancillary services, 0.07; and for home visits, 0.35. This pattern of a low elasticity for hospital services and ancillary services, a larger one for office visits, and the largest one for home visits is in accord with the underlying model developed here. The elasticity for prescription drugs appears to be 0.07 in this range and the elasticity for dental services between 0.07 and 0.16.

ACKNOWLEDGMENTS

The authors would like to thank Leonard Zucker and Clayton Hearst of the Stever Companies for providing some of the insurance premium data. They would particularly like to thank Bridger Mitchell, who not only found a critical mistake in a preliminary version of this report, but also spent considerable time convincing two rather stubborn authors of the error of their ways. Thanks are also due to Kenneth Arrow, David Chu, Martin Feldstein, and Michael Grossman for comments on a previous draft. Any remaining errors are the responsibility of the (demonstrably fallible) authors.

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> The real price of every thing, what every thing really costs to the man who wants to acquire it, is the toil and trouble of acquiring it.
>
> Adam Smith, *The Wealth of Nations*

I. INTRODUCTION

The effect of coinsurance payments on demand for medical services (and for other reimbursement insured goods and services) has been debated for many years.[1] Some assert that coinsurance helps control total expenditures by giving consumers a financial stake in how much medical care is purchased. Others assert that coinsurance is irrelevant to choice, because the physician makes the decisions about using medical services for his patients. Persons attempting to predict expenditures under various national health insurance systems are naturally interested in the degree to which coinsurance affects demand for services. The evidence we present in this report decisively rejects the assertion that coinsurance is irrelevant to choice; coinsurance clearly does affect the demand for services. Moreover, as we shall show, the impact of coinsurance varies across medical services in a systematic fashion depending on the time price of the service.

Appendix A provides the derivation for expressions relating responsiveness of demand to coinsurance rates, to market prices for medical care, and to "time costs" for medical care. The basic model assumes that consumers maximize a utility function, subject to a budget constraint. The utility function is in "other goods" (x) and "health" (H). For the moment we assume that medical care (h) is a homogeneous commodity that can be purchased in the market at a known and fixed price of p_h per unit, and other goods may be purchased at p_x per

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[1] Reimbursement insurance pays benefits in proportion to expenditures incurred, rather than a lump-sum payment to offset the loss. What is frequently described as indemnity insurance in the health field is actually of reimbursement form--expenditures are reimbursed up to a prespecified maximum.

unit. A production function g(h) changes medical care into health.[1] Purchases of medical care involve time inputs (t) caused by travel time, waiting time at the point of provision, and actual medical time (days spent in the hospital, time in a dentist's chair, and so on).[2] It is assumed that the opportunity cost for that time is given as w per unit of time.[3] The budget constraint is then $I = wT = p_x x + p_h h$, where T is the amount of productive time available to the person, defined by $T = T_0 - th$ (T_0 is fixed in this problem).[4] Purchases of medical services have two prices--the money price p_h per unit, and the time price wt per unit.[5]

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[1]Formally, we may consider the production function g(h,t) a function of both market medical care and own time. For expositional simplicity we write g(h) rather than g(h,t); that is, we assume h and t are used in fixed proportions. Relaxing this assumption would add another term to our equation reflecting the elasticity of substitution of medical goods for time in the production of H. Specifically, lowering the price of medical care facing the consumer would not only increase the demand for H by lowering its shadow price, but also would induce a substitution of h for t in production. Taking account of such a substitution would not alter any of the results derived here. For a more complete discussion, see Grossman (1972) and Newhouse and Phelps (1974a).

[2]For simplicity, we assume that there are no time inputs in the production of x. Relaxing this assumption would lead us to consider the relative time intensity in consumption of H and x, when analyzing changes in the wage rate (w) or medical service time (t); to do so would not lead to substantive changes in our conclusions about the effects of coinsurance. For a more complete treatment of the problem, see Acton (1973), Grossman (1972), and Newhouse and Phelps (1974a).

[3]For expository convenience, we do not allow here for varying prices of time depending on the individual's situation. For example, an individual who is sick enough to be hospitalized may have a considerably lower opportunity cost of time than a person who is visiting a physician for an inoculation. In the empirical section of this study, we do take account of such differences.

[4]An inter-temporal model based on the decrease in sick time that may be caused by purchase of medical services is developed by Grossman (1972) as the investment model. Although we ignore this aspect of medical services, to take account of it would not generate qualitatively different predictions with respect to any of the variables considered here.

[5]In the real world there is a third type of time cost, a simple delay time in seeing a physician or being admitted to a hospital. Such time may be spent doing other things and have no opportunity cost in a

The consumer's level of health may be considered random, which induces him to purchase insurance against the costs associated with randomness.[1] The final level of health consumed is $H = H_0 - \ell + g(h)$, which is dependent not only on the illness observed (ℓ is random) but the amount of medical care purchased in response to that illness (h is a function of ℓ) and the initial health stock level H_0. The consumer may purchase an insurance contract that specifies a coinsurance rate for medical expenditures--the consumer pays C percent and the insurer pays (1 - C) percent of incurred expenses during the period. In addition, there may be an upper limit h* on the amount the insurance policy will pay. The insurance contract is made before the actual value of ℓ is known, so the cost of the contract is

$$R = \int_0^{\ell^*} (1 + \theta)(1 - C)p_h h(\ell)\, f(\ell)\, d\ell + (1 + \theta)(1 - C)p_h h^* \int_{\ell^*}^{\infty} f(\ell)\, d\ell, \tag{1.1}$$

where $f(\ell)$ is the known distribution of losses and $h(\ell)$ is the loss-specific amount of medical care purchased; ℓ^* is the health stock loss associated with h*. The loading fee (θ) is any charge above expected benefits to cover administrative costs or risk assumption.

We are not concerned here with the selection of C, but how the consumer reacts to C when ℓ is observed. Assume that C has been previously chosen, or is imposed; in either event, C is fixed during the

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one-period model. To treat the problem formally would require a multi-period model with the consumer being unable to purchase medical services in the current time period. Such a model is beyond the scope of this study.

[1] Note that sick leave is insurance against the time price of medical care, and disability insurance protects against random drops in w due to illness. We focus here on conventional insurance against the money price of care.

period, and R (a function of C) is prepaid.[1] For a public program, R should be considered as the consumer's share of taxes necessary to finance the program.

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[1]For an analysis of how the amount of coinsurance is chosen, see Phelps (1973). The theory developed here is treated more extensively in that report. For another analysis of demand for insurance, including one where the loss distribution may be affected by the consumer, see Ehrlich and Becker (1972).

II. THEORETICAL RESULTS

The consumer's decision problem is to maximize a utility function $U = U[x, H_0 - \ell + g(h(\ell))]$ subject to the budget constraint $I = p_x x + Cp_h h + R$, and given a certain illness ℓ.

The first-order conditions from the budget-constrained utility maximization are

$$U_x + \lambda p_x = 0, \tag{2.1a}$$

$$U_H g'(h) + \lambda(Cp_h + wt) = 0, \tag{2.1b}$$

$$p_x x + Cp_h h + R - I = 0, \tag{2.1c}$$

where $(-\lambda)$ is the marginal utility of I.

The consumer views the total price (or shadow price) as $Cp_h + wt$, which may be greater than, equal to, or less than the market price p_h, depending on his insurance, the opportunity cost of time w, and the time per unit of service t. The net income available for consumption of x and h is I - R, with the additional income loss of wt per unit of h actually consumed.

Of interest here are relationships between demand for medical care and the various price components p_h and C and also w and t. It is shown in Appendix A that[1]

$$\eta_{hC} \approx \eta_{hp_h} \approx \frac{Cp_h}{Cp_h + wt}\left(\eta_h^T\right), \tag{2.2}$$

---

[1]The nearly exact equality between coinsurance elasticities and own-price elasticities as shown in Eq. (2.2) does not hold if the gross price p_h is a function of the insurance plan. Under this circumstance, $|\eta_{hp_h}| > |\eta_{hC}|$ because the former is the sum of the pure demand effect and the price of provider effect. We can envision several circumstances

where η_{hC} is the elasticity of demand for h with respect to C; η_{hp_h} is the elasticity of demand for h with respect to p_h; and η_h^T is the total-price elasticity. The approximations are due to income effects (from premium changes), which can be shown to be of small importance empirically.[1]

Similarly,

$$\eta_{hw} \approx \eta_{ht} \approx \frac{wt}{Cp_h + wt}\left(\eta_h^T\right), \tag{2.3}$$

where η_{hw} is the elasticity of h with respect to w, and η_{ht} is the elasticity of demand with respect to time per unit of h.[2] Except for the income effects from premium changes, the time-price and the money-price elasticities sum to the total-price elasticity.

These relationships show formally how one may estimate the own-price elasticity of demand for a good by observing response to different coinsurance rates. We shall make use of this result in the empirical work. Additionally, a relationship between response to coinsurance and response to time costs is given implicitly, which allows one to consider effects of changes in travel time to facilities and queues in estimating demand on a medical care system.

The theory developed thus far and in Appendix A assumes medical care is a single homogeneous commodity. Clearly there are various

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under which price might be affected by coinsurance level. At the individual consumer level, if higher price (for amenities or whatever) leads to shorter wait times (lower t), then the consumer will rationally substitute (insured) money price for (uninsured) time price. Thus a uniform coinsurance C would change the relative total-prices for various classes of provider. In a market setting, introduction of better insurance might induce a market shift toward more amenities, thus raising observed medical prices. Second, if the supply of physician services is not infinitely elastic in the long run, then decreasing coinsurance would increase demand, thus raising the market clearing price.

[1] See Appendix A for these derivations and a discussion of the small importance of the omitted income effect.

[2] In this model, the wage and time elasticities are identical if we consider a compensated wage effect, although this would not be true in an investment model.

medical services. To take account of this and derive further implications, we need the additional assumption that total-price elasticities are approximately equal across various medical services. Conditions under which this is true are outlined in Appendix B. To derive refutable hypotheses, we assume this to be the case. For the same reason, we also assume that η_h^T is approximately constant over losses and prices.

The following implications can then be drawn:

(1) Goods with relatively high time price components and nearly complete insurance coverage (C near zero) will show relatively small money-price and coinsurance elasticities, but high time-price elasticities. A student or other free clinic is one example of such a service. Hospital days might also be an example of such a service, although the opportunity cost of time, w, probably falls if one is seriously enough ill to be hospitalized.

(2) Goods with a low time price and poor insurance coverage may be highly sensitive to money price or insurance coverage changes. Physician home visits are an example of such a service.

(3) The effects of a given coinsurance change will differ substantially across services, depending on how large the time price component of total price is.

(4) One may infer time-price elasticities from money-price elasticities, if the appropriate *levels* of the two types of prices are known and if the total-price elasticity of the service in question is known (and vice versa).

(5) In models that specify constant elasticities, one must take care to specify whether it is the total-price elasticity or the money-price elasticity that is constant. If total-price elasticities are said to be constant, then money-price elasticities fall as C approaches zero, and time-price elasticities rise correspondingly.

(6) The responsiveness to coinsurance rates, money prices, and time prices may vary with the size of the illness observed

(that is, the elasticities may depend on ℓ), because the total-price elasticity may vary with ℓ, although this problem is assumed away in our empirical work.

In the subsequent sections we will subject the first five implications to empirical tests.

III. EMPIRICAL EVIDENCE: THREE STUDIES IN ECONOMICS JOURNALS

We begin by reviewing three recent studies by economists of the price elasticity of demand for medical services.[1] One studies the demand for hospital inpatient services, a second the demand for hospital inpatient and hospital outpatient services, and the third the demand for all medical services. The estimated elasticities in these studies are all much higher than the elasticities we report below, making it worthwhile to examine them in some detail.[2] Two other articles in the economics literature that are solely concerned with the length of hospital stay we shall defer until a later discussion of hospital services.

Feldstein uses state aggregate data for 1958 through 1967. He therefore has a time series of cross-sections, which permits him to estimate dynamic effects. Use of data at the state level abstracts from many features of insurance coverage. Thus Feldstein is forced to assume that an insurance policy can be represented by one parameter, the coinsurance rate. Although this is convenient, it may lead to a serious distortion, because many policies present a price schedule to the consumer rather than a constant price. For example, if there is a deductible and a coinsurance rate above the deductible, the consumer will pay 100 percent of the gross price until the deductible has been met, then the coinsurance rate times the gross price. This is not an atypical arrangement. To ignore this will lead to biases in the size of insurance-elasticity estimates.[3] In the

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[1] Feldstein (1971); Davis and Russell (1972); Rosett and Huang (1973).

[2] The reader should note that in all our empirical references, we denote price elasticities as *positive*. A more detailed discussion of all of these papers is contained in Newhouse and Phelps (1974b).

[3] Keeler, Newhouse, and Phelps (1974). More generally, there is a bias any time the marginal coinsurance rate falls with total expenditure. There is also a bias if a "limit" exists so that the consumer must pay 100 percent of the gross price (i.e., expenditures exceed a certain large amount). This latter problem is not likely to be of practical importance.

empirical work presented below, we have examined data in which the only variation was in the coinsurance rate so that there is no problem of bias. Operating at the state level, this is impossible to do with existing data.

Feldstein, however, needs still another assumption. There are no data to permit one to estimate a state coinsurance rate. Rather, Feldstein must assume that an individual in a given state who had insurance in any year had a policy that paid the average amount paid by insurance policies that year in the nation. Thus, in any given year, all variation across states in the insurance variable is attributable to differences in the proportion of the state's insured population. This procedure appears to lead to bias in the coefficients away from zero.[1] Feldstein admits that "the resulting values . . . are a very imperfect measure of hospital insurance."

Without allowing for dynamic adjustments, Feldstein estimates that price elasticity for hospital days is 0.67 when gross price elasticity is constrained to be equal to coinsurance elasticity (as we have argued above is appropriate).[2] When dynamic considerations are introduced, the elasticities fall to 0.55.[3] A substantial difference appears, however, when Feldstein relaxes the constraint that price and coinsurance elasticities be equal. Gross price elasticity remains large (0.71), but coinsurance elasticity falls to 0.17.[4] The unconstrained equations give coinsurance elasticities much closer to the size of the elasticities we infer from the evidence presented below. Feldstein prefers the 0.67 estimate, in part no doubt because of the discrepancy between gross price and coinsurance elasticities.[5]

Davis and Russell (1972) also use state data to estimate price elasticities, thus facing the same problems as Feldstein. However,

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[1] A more extensive discussion, together with some numerical results, is contained in Newhouse and Phelps (1974b).

[2] Feldstein's Eqs. (24) and (26).

[3] Feldstein's Eqs. (29) and (30).

[4] Feldstein's Eqs. (27) and (28).

[5] See Feldstein (1973).

their estimates are based on a single cross-section of states in 1969. Davis and Russell do not constrain insurance and price elasticities to be equal. Their insurance variable is simply the percentage insured in each state, which they enter in logarithmic form. This can be shown to be a misspecification, given their basic model.[1] Davis and Russell use three price variables for inpatient care: revenue per patient day, inpatient revenue per admission, and basic charge for a two-bed room. They use gross hospital outpatient revenue per hospital outpatient visit as the price of hospital outpatient services. Since insurance is not adequately controlled for, one cannot know a priori what biases enter into estimation of the coefficient of price, but evidence in Newhouse and Phelps (1974b) suggests that the estimated effect of price in the Davis-Russell study is roughly twice as large as the true value. Davis and Russell obtain admission elasticities ranging from 0.19 to 0.46 and length of stay elasticities ranging from -0.35 to +0.30.[2] A summary measure of their estimated price elasticities for hospital days might be 0.5, and for hospital outpatient visits, 1.0. Although these estimates are consistent with Feldstein (as Davis and Russell note), they are much larger than the elasticities we present below. As in Feldstein's case, this could be because of differences in the range of price variation between their data and ours.

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[1]The proportion of insured can vary from zero to one. In the multiplicative form implied by Davis and Russell's log-linear specification, this means demand varies from zero (when none are insured) to whatever demand is when all are insured. Such a multiplicative combination is clearly not a linear combination of the insured and uninsured demands. If we adopt Davis and Russell's assumption that demand is given by $(Cp)^a$, where C is the coinsurance rate and p is gross price, and if we let w be the proportion of consumers with coinsurance rate C (leaving 1 - w percent of consumers facing the full gross price), then the aggregate demand equation would be $d = w(Cp)^a + (1 - w)p^a = p^a(1 - w(1 - C^a))$ or $\log(d) = a \cdot \log(p) + \log[1 - w(1 - C^a)]$ rather than the form used by Davis and Russell: $\log(d) = a \cdot \log(p) + b \cdot \log(w)$.

[2]Their introductory remark that inpatient elasticities are 0.32 and 0.46 apparently refers only to the admissions elasticities. The reader should recall that we report price elasticities as positive; hence, the -0.35 was reported by them as +0.35, and so on.

A study by Rosett and Huang (1973) also assumes that insurance can be represented by a single parameter--the coinsurance rate. Unlike the two previous studies, their study uses household level data. Unfortunately, the data (the 1960 *Survey of Consumer Expenditure*) only describe the amount of *out-of-pocket* medical expenses and *out-of-pocket* premiums; any employer insurance premium payments are omitted. The seriousness of this omission may be gauged by the statistic that in 1963, employer premiums represented over 30 percent of total health insurance premiums (and thus a greater proportion of the effective coverage, since group insurance provided through employers has a much lower loading fee than family-paid insurance, much of which is individually purchased).[1] Given Rosett and Huang's methodology, those with any employer-paid premiums will have a coverage value (k) assigned lower than the true coverage parameter, thus biasing their insurance elasticity away from zero. They must also assume that all insurance policies have the same loading fee, which induces systematic error into their estimates of coverage, and biases their elasticity estimate further from zero.[2] Rosett and Huang are placed in the unenviable position of attempting to estimate a price elasticity without being able to directly measure either the dependent variable (expenditure) or the key explanatory variable (insurance coverage).

By far the most serious problem facing Rosett and Huang, however, is that their data aggregate various types of medical services that are insured at different levels. Because expenses that are large are well insured on average (hospital care), while expenses that are small are poorly insured on average (physician care, drug purchases, etc.), the regression of total expenditure on insurance leads to an overestimate of the price elasticity of care. In Newhouse and Phelps (1974b) we have shown that this aggregation across services induces

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[1] Phelps (1973). The 30 percent figure is computed from unpublished data from the 1963 National Health Survey conducted by the Center for Health Administration Studies, University of Chicago. Families with verified premiums were used for this calculation.

[2] For a systematic treatment of the problems in the Rosett and Huang paper, see Newhouse and Phelps (1974b).

an upward bias of an order of magnitude (using a data base similar to that used by Rosett and Huang).

These studies face some common problems. All have attempted to model behavior of consumers on the assumption that there is only one parameter in their insurance policies--the coinsurance rate. None of the studies have provided any evidence that the empirical results are not sensitive to this assumption.

Second, actual policies frequently cover outpatient and inpatient services differently. Feldstein avoids the problem of differential coverage of services by not specifying an outpatient price term. Under reasonable assumptions this leads to a further upward bias, which is probably small empirically.[1] Davis and Russell include separate price terms but do not have an adequate measure of the insurance coverage of either. Rosett and Huang study total expenditures, thus masking covered and uncovered service differences.

Third, Feldstein and Davis and Russell work with aggregate data, rather than data on individual consumers. The difficulties of working with such data are well known, but could be reiterated here in the context of these studies. Specifically, for such estimates to remain stable, either the coefficients for individuals making up the aggregate must be identical (or randomly distributed through time from the same distribution), or the distribution of the explanatory variable must remain unchanged. Because a primary reason for generating evidence on elasticities is to estimate the effects of national legislation, it must be assumed that the distribution of insurance among individuals is likely to change. Hence the case for a stable relationship must rest on the assumption that all individuals have identical response function over all coverage levels (or that their responses are randomly distributed over time), an assumption that cannot be strongly supported, and which implies that η_h^T rises to infinity as C falls to zero (see implication 5, page 7).

Additionally, Davis and Russell and Rosett and Huang treat

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[1] Newhouse and Phelps (1974b).

insurance coverage as exogenous. Since those at greater risk appear to systematically purchase more insurance,[1] their estimates are further biased away from zero.

We conclude from analyzing these studies in economics journals that there is little firm information on demand elasticities for medical care. If anything, there appears to be something of a consensus that demand elasticities are large. The consequences of these large elasticities should be pointed out. If, for example, there is an arc-elasticity of 0.67 for a coinsurance ranging from zero to 0.25, then the difference in demand from insurance plans offering those two coinsurance rates is 134 percent, which seems extraordinarily large and certainly at odds with data we present below. (In fairness to the authors of these studies, it should be pointed out that they do not attempt to extrapolate their elasticity estimates to this range.)

It is important for policy purposes to establish the magnitude of these elasticities. In what follows we present some new evidence and review evidence that is not in the economic literature to obtain an estimate of the average arc-elasticity in the range of zero to 25 percent coinsurance.[2]

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[1] Phelps (1973).

[2] It is also policy relevant to consider deductibles, but they cannot simply be treated as 100 percent coinsurance. See Keeler, Newhouse, and Phelps (1974).

IV. OTHER EMPIRICAL EVIDENCE

TOTAL MEDICAL EXPENDITURES

A variety of data sources may be used to estimate the coinsurance (or money-price) elasticity of demand for medical services. We have gathered a number of different sources of data (some published, some original) that allow us to infer the elasticity of demand for h with respect to C, either directly or through insurance premiums.

Our first source of data comes from four insurance companies. These data show the cost of insuring medical expenditures in a group of persons with characteristics equivalent to the national average. The insurance covered all expenditures except mental and dental treatment, up to a lifetime maximum of $25,000 (with a $1,000 annual reinstatement provision). We asked for quotes from actual experience, over as broad a range of deductibles and coinsurance as was possible. The companies quoted premiums for policies with coinsurance rates of 0.10, 0.15, 0.20, and 0.25, and deductibles ranging from $25 to $1,000.[1] Although the dollar amount of the premiums differed among the companies, the ratio of premiums for various coinsurance rates was independent of the deductible. The resulting index of premiums I for the four companies (20 percent coinsurance equals 1.00) is shown in Table 1.

We have fitted a simple linear equation to these data, predicting the value of index I as a function of C, where C ranges over the values 0.10 to 0.25. This equation is

$$I = 1.34 - \underset{(t = 42.28)}{1.68\ C}, \qquad R^2 = 0.993. \tag{4.1}$$

[1] In a 1970 nationwide survey of insurance companies issuing *comprehensive major medical* insurance, 78.5 percent of the insurance in force had a coinsurance rate of 0.20; 4.5 percent had a coinsurance rate of 0.25. See Health Insurance Association of America (1970). Other hospital plans pay "full semi-private" rates for hospitalization, and still others impose "internal limits" of some dollar-per-day maximum.

Table 1

INDEX OF PREMIUMS

Coinsurance Rate	Insurance Company			
	A	B	C	D
25	0.94	0.91	0.92	0.91
20	1.00	1.00	1.00	1.00
15	1.09	1.08	1.08	n.a.
10	1.18	1.17	1.17	n.a.

As is obvious, the fit is nearly perfect. Using this equation, $\hat{I}$ for 25 percent coinsurance rate is 0.92 and for 10 percent is 1.17.

In Appendix A we have derived the appropriate formula for converting these predicted premium ratios into the insurance companies' perceived change in demand for services, given the change in coinsurance. Approximately what we do is compute the change in premium if utilization did not change and attribute the residual to increased utilization; the interested reader should consult Appendix A for details. The implied insurance elasticity of demand is 0.07 between 25 and 10 percent. The implied elasticity falls through the range of coinsurance; it is 0.12 between 25 and 20 percent, 0.08 between 20 and 15 percent, and 0.04 between 15 and 10 percent. Extrapolating, it is 0.01 between 10 and 0 percent. These calculations assume that there is no self-selection by "sickly" people toward the lower coinsurance rate plans in these data; if that is true, then our estimates are further from zero than the "true" effect of coinsurance. The decrease in elasticities between 25 percent and zero coinsurance is consistent with implications 1 and 5 (page 7) that η_{hC} will fall with C.

The concept of arc-elasticity can be deceptive when the ranges over which the elasticities are computed differ. We have therefore used a standardized interval to compute arc-elasticities. For a standard interval we have chosen a range we regard as policy relevant--25 percent to zero coinsurance. When our insurance data are extrapolated to zero, the value of the premium index is 1.34, and the estimated arc-elasticity over the 25 percent to zero range is 0.043. This represents

an 8 percent increase in expenditure as the coinsurance rate is dropped from 25 percent to zero.[1]

In order to test the validity of this calculation, we have estimated arc-elasticity in the 25 percent to zero range of coinsurance for various types of medical expenditures. We can then aggregate these figures (using budget shares as weights) and compare the result with the 0.043 value derived from the premiums. Data on the elasticities of various services will also permit us to test our implications derived above on the variation of observed elasticities by type of service.

PHYSICIAN VISITS

A variety of published articles contain sufficient information to calculate arc-elasticities of demand for various medical care services. These articles are generally outside the economics literature. In 1967, a 25 percent coinsurance rate was initiated for all members of a prepaid medical care plan (who were employees of Stanford University). Before 1967 the coinsurance rate was zero. The coinsurance was introduced to reduce budget deficits for the plan. Individual observations on 2,567 members of the plan in the first full year before the coinsurance (1966) and the first year after the coinsurance (1968) were obtained by Scitovsky and Snyder (1972). Multiple regression analysis of those data shows the partial effects of the coinsurance rate, from which an arc-elasticity of 0.138 for all office visits was calculated.[2] For ancillary services (X-ray, lab tests, and so on), the implied arc-elasticity was 0.07.

Further data from Scitvosky and Snyder show that home visits

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[1]There are numerous other data from the insurance literature, but most are of little value in assessing effects of coinsurance. A number of actuarial studies, for example, show differences in per capita expenditures from a whole spectrum of geographic locations, where the nominal price of medical care and amount of insurance coverage differ. Without an investigation of the partial effects of coinsurance (that is, holding the real price of medical care constant), these data are of limited value; they may show only the geographic distribution of insurance policies rather than effects of coinsurance on demand.

[2]Phelps and Newhouse (1972).

decreased much more in the face of the coinsurance; the implied arc-elasticity from their data for home visits was 0.35.[1] The higher elasticity of demand for home visits is strong evidence supporting our theory of time-price and money-price responsiveness. The time-costs for home visits are essentially zero, because the doctor rather than the patient does the traveling. Typically, a person receiving a home visit is sick in bed, and the opportunity cost of even waiting for the doctor may be said to be quite low. Using Eq. (2.2), and assuming wt equals zero, 0.35 can be given as an approximation of the total-price elasticity of demand for home visits.

Data showing the same pattern of elasticities for home and office visits come from Canada. In 1946 the Swift Current District in the Province of Saskatchewan instituted a prepaid medical plan providing full coverage for all enrollees, partly financed by personal premium payments.[2] Eligibility was determined by municipal residence. Doctors were paid on a fee-for-service basis, according to a fee schedule. Because of high utilization, a copayment provision was instituted in 1953 according to the schedule in Table 2. We estimate that the charges

Table 2

COPAYMENT RATES, SASKATCHEWAN HEALTH PLAN

Office call	$1
Minor surgery in office	$1
First maternity care visit	$1
Home call--day	$2
Home call--night	$3

SOURCE: Straight (1962), p. 75.

[1] Scitovsky and Snyder (1972), Table 10.

[2] Although it seems unlikely, there may have been self-selection because of the family premium payments. Enrollment was near 90 percent of the eligible population.

for office and home visits represent 42 percent coinsurance rates in each case.[1]

The copayments for home visits went into effect on January 1, and the payments for office visits on August 1. In Table 3 we show the number of visits per subscriber between 1952 and 1954 for home visits and office visits. The effects of copayment on utilization are shown by comparing office visit use between 1952 and 1954 and home visit use between 1952 and 1953.[2]

Table 3

CHANGES IN DEMAND WHEN COPAYMENTS WERE INTRODUCED IN SASKATCHEWAN IN 1953

Year	Office Visits Per 1000	Home Visits Per 1000	Hospital Calls
1952	2187	398	1239
1953	2136	158	1259
1954	1812	175	1416

SOURCE: Straight (1962), p. 75.

[1] Gross prices were computed from total costs paid by the plan divided by total visits, divided by 0.68, because there was a 68 percent payment ratio in 1952 from the plan to participating doctors. This yields values $2.40 and $4.72, as the market price for office and home visits respectively; 0.42 = 1/2.40 = 2/4.72. The 42 percent rate for home visits assumes that all such visits were daytime visits and so is a lower bound. An upper bound is 63 percent, if all visits were night visits; 0.63 = 3/4.72.

[2] For office visits, the year when copayments were introduced (1953) cannot be used to show the effect of copayment on use. Because of seasonal fluctuations in demand for physician visits, one cannot accurately "normalize" the portion of 1953 when money prices were charged for office visits. There may be other effects operative as well: Income levels may have changed; an infectious disease epidemic may have been present in one of the years; there may have been self-selection out of the plan when the copayment charges were introduced. These factors are assumed to have had no net effect. (Enrollment decreased from 48,000 in 1953 to 47,000 in 1954, the only such decrease in an eight-year span. From 1952 to 1959, enrollment increased from 46,000 to 52,000.)

These data cannot properly be used to derive an elasticity of demand. Unlike the Palo Alto case, where the Stanford employees constituted a small fraction (16 percent) of the medical group's practice, in Saskatchewan the insurance coverage of an entire region was changed. Thus, supply effects may have played a role in the outcome. For example, there might have been an even higher number of visits in 1952 had a larger supply existed. In addition, since demand fell as copayments rose, queues and delays in the system could also fall, reducing the net price to the consumer. The reduction in utilization is thus a lower bound on the amount demand fell. We cite these data because they show the same pattern for home and office visits as the Palo Alto data. The reduction in office visits (1952 to 1954) was 17 percent, and the reduction in home visits (1952 and 1953) was 60 percent.[1] This difference supports the theory of demand developed above.

Some Evidence on the Time-Price Elasticity of Physician Visits

Our theoretical formulation suggests that time should act as a rationing device, just as money does, in the demand for medical care. Data supporting this hypothesis come from a natural experiment analogous to the Palo Alto study we report above. The data show monthly visits to a university health service clinic before and after a change in the physical location of the clinic (Simon and Smith, 1973). The mean time for students to travel to the clinic before the move was 5-10 minutes, and the mean time after the clinic was moved was 20 minutes. A multiple regression equation was estimated with utilization as a function of "time" variable (measured in months) and a dummy

[1] Both hospital and home visits rose slightly in 1954 (home visits rose by 10 percent and hospital visits by 12 percent), indicating that some outside force (perhaps an infectious disease epidemic) increased demand for all medical services. If, say, because of a flu epidemic, there was a transitory 10 percent increase in demand for all medical services in 1954, then the permanent reduction in office visits attributable to the copayment change might be on the order of 550 visits per 1000 population, rather than the 375 visits derived from Table 3. The increase in home visits and hospitalization may also indicate a substitution away from office visits toward alternative medical treatments in the face of increased relative prices between office visits and the alternative services.

set to 1.0 for months after which the clinic had moved.[1] The estimated equation was

$$(4.2)\quad y = \underset{(t = 0.52)}{-5.9} \cdot (\text{time}) - \underset{(t = 1.95)}{489} \cdot \begin{pmatrix}\text{clinic location}\\ \text{dummy}\end{pmatrix} + 1789,$$

where $\bar{y} = 1491$, $R^2 = 0.37$, Durbin-Watson = 1.99, and $F(2, 34) = 9.799$. From this equation, we know that $\Delta y/\bar{y} = -489/1491 = -0.328$. A time-price elasticity may be calculated by assuming some value for the change in travel time. If initial travel time was 5 minutes, then $\Delta p = 15$ minutes, $\bar{p} = 12.5$, and $\eta = 0.27$. Alternatively, if initial travel time was 10 minutes, then $\Delta p = 10$, $\bar{p} = 15$, and $\eta = 0.49$. These two calculations bound the time-price elasticity given Simon and Smith's estimates of initial travel time.[2]

One should only use this estimate with caution in other settings, since the population covered was young (and presumably healthy) and the clinic would typically treat only less serious illnesses. We present the data primarily to demonstrate an independent effect of a change in the time-price of a service when the money price is held constant (at zero).

Additional support for our theoretical formulation can be found in Acton's (1973) results. Acton estimates travel-time-price elasticities in free care clinics to be between 0.6 and 1.0 using survey data, while for private physicians, he estimates a travel-time-price elasticity of 0.3 or lower. The lower value for private physicians,

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[1] Visit levels estimated from Simon and Smith (1973), Fig. 1.

[2] Several factors may prevent one from interpreting this estimate as a true demand elasticity. First, since the travel time changed for almost all students and since total use in the system fell, the actual time *at* the clinic may have fallen, making the change in total time for a visit less than the change in travel time. If so, the true time-price elasticity would be larger than the figures reported above.

Second, other features of the clinic system changed when the clinic moved. Seven new treatment rooms, additional staff, and X-ray, physical therapy, and laboratory services were added, and the new facilities may have been more pleasant than the old ones.

where patients also faced a money price is consistent with implications 1 and 2 (page 7).

Indirect Estimates of the Value of Time for Physician Visits

Two of the above studies allow estimation of the value of time for an office visit (implication 4), given certain assumptions.[1] In the student clinic study, we have assumed that average actual treatment time is 15 minutes (we have no data), implying that average total time is 30 minutes. From Eq. (2.3), this would imply that η_h^T is twice η_{ht}, or about 0.79 (using the midpoint of the two estimates of η_{ht}). From the Palo Alto data, where there was no time price for home visits, we can use the coinsurance elasticity of 0.35 as an estimate of η_h^T. Using these values of 0.35 and 0.79 for η_h^T in Eq. (2.2), with an estimate of $\eta_{hC} = 0.14$ and an average money price in Palo Alto of \$3.46 (Phelps and Newhouse, 1972), the implicit value of time estimated from these two values of η_h^T ranges from \$5.19 to \$16.06.[2] Although the estimated range is large, these values are within reasonable expectation, thus lending support to our theory. The point of policy interest is that for conceivable values of η_h^T, the time cost is estimated to be a significant and substantial proportion of the total cost of ambulatory medical care and will rise in relative importance as C goes to zero. Thus we can expect time to act as a significant rationing device in the absence of money prices (see also Acton, 1973).

HOSPITAL SERVICES

A study by Heaney and Riedel (1970) provides evidence on the elasticity of demand for non-maternity hospital services. They studied changes in utilization of group-enrolled persons in Connecticut in 1966-1968, when the Blue Cross plan in that area offered a change from indemnity insurance (paying \$15 per day for hospital care) to full semi-private coverage. Some of the groups studied were small (under 25 persons enrolled), and others were quite large (over 400 enrollees).

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[1] See Appendix B.

[2] $5.19 = (0.35 - 0.14)\ (3.46)/0.14$; $16.06 = (0.79 - 0.14)\ (3.46)/0.14$.

There may have been some self-selection in those groups choosing the more complete coverage; no evidence is available from the study to determine that. The change in demand does not appear to be large relative to total demand, so the results may be interpreted as demand elasticities.[1]

The study compares hospital admission rates and length of stay of the groups opting for better coverage six months before and six months after the more complete insurance was introduced. These groups switched insurance coverage at various points during a two-year period, so there may be seasonal effects dominating any one group's behavior. But a comparison of the total activity of these groups to their previous activity should give some indication of the elasticity of demand for hospital services.[2]

It is difficult to impute a coinsurance level to the indemnity insurance plans, because no information was given on the total expenditure for hospital services, or (more appropriately) on the average per diem rates charged by the hospitals in the area. Using average patient revenue per day for Connecticut hospitals in 1967 of $65.89,[3] and noting that room and board charges are on average 53.5 percent of total hospital charges,[4] one can infer that the average room and board charge facing these persons was $35.20. The $15 per day insurance benefit represents a 57 percent coinsurance rate for room and board charges ($20.20 out-of-pocket per day). Because ancillary services were fully covered in both the indemnity and full-service contracts,

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[1] From data in Heaney and Riedel, it appears that the coverage of some 600,000 persons was changed, or slightly over 20 percent of Connecticut's population. The occupancy rate in short-term general hospitals in Connecticut in 1966 was 80 percent.

[2] Heaney and Riedel also provide data on groups whose coverage did not change; they characterize such groups as controls. We do not use these data because the so-called controls are self-selected and because such data do not assist in adjusting for seasonality.

[3] *Hospitals* (1968), p. 442.

[4] Health Insurance Association of America (1968), Table 3.

the average out-of-pocket expense of $20.20 represents a coinsurance rate of 31 percent of the average total cost of a hospital stay. A full semi-private plan, of course, represents a zero coinsurance rate.

It is clear that the 31 percent coinsurance is more closely correct for hospital admission decisions. For length of stay, the 57 percent room and board figure is more likely to be correct; because ancillary expenses systematically decline with length of stay, charges for marginal days in the hospital are essentially room and board charges.[1] Table 4 shows hospital utilization at the different coinsurance levels and gives the corresponding arc-elasticities of demand.

Table 4

UTILIZATION DIFFERENCES, CONNECTICUT HOSPITAL STUDY

Item	Actual Utilization		Arc-Elasticity	
	Indemnity Plan	Full Coverage Plan	Actual Range[a]	0-25% Coinsurance[b]
Hospital admissions	4.56	5.09	0.05	0.04 (t = 1.70)
Average stay per admission	7.26	8.17	0.06	0.03 (t = 1.48)
Patient days			(c)	0.07

[a] For admissions, actual coinsurance was 31 to 0.0 percent. For average stay, actual coinsurance was 57 to 0.0 percent.

[b] Computed by fitting a linear demand curve to two points observed, and computing elasticity from fitted demand at C = 0.25 and C = 0.0. t-statistics are taken from Heaney and Riedel (1970), Tables 3 and 6.

[c] Elasticities cannot be added since they are computed over different coinsurance intervals.

[1] Note also that, with a low indemnity payment, one would anticipate that Blue Cross subscribers would systematically choose hospitals with lower, rather than average, prices. This would lower their apparent coinsurance rate below 0.57, but we have no data to estimate the magnitude of this effect. The implication for our results is

As discussed above, we wish to obtain standardized results by having all elasticities refer to the 25 percent to zero coinsurance range. To do this, we must interpolate from these data. For hospital admissions, using the 31 percent coinsurance figure for the indemnity plan, the arc-elasticity in our standardized range is 0.04. For length of stay (using C = 0.57 for the indemnity plan), the arc-elasticity in the standardized range is 0.03. The total patient-days elasticity in the standardized range is the sum of these, or 0.07.

This estimate is a lower bound on the expenditure elasticity with respect to coinsurance, since no adjustment is included for changes in resources used per patient day. Newhouse and Phelps (1974a) present data indicating that increases in resources demanded per day may increase the expenditure elasticity to 0.09 or 0.10.

A second study of hospital utilization (Williams, 1966) showed differences in demand under two Blue Cross plans in 1964--one with full payment for hospital days, and the other with a $4 per day copayment. Table 5 shows the basic data. The $4 per day payment represents an estimated coinsurance of 12 percent.[1]

The arc-elasticity of patient days per 1,000 persons per year is 0.07 in the 0.0 to 0.12 coinsurance range. Of more interest is the elasticity of resource use per subscriber, since this is closest to the concept of an expenditure elasticity. The resource-use elasticity is 0.04 over the coinsurance range of 0.0 to 0.12 coinsurance. When we extrapolate this to the 0.0 to 0.25 coinsurance range, the figure is 0.08, quite close to the 0.09 to 0.10 figure obtained from the Connecticut hospital study.

Note that for patients on the $4 per day plan, there is an incentive to use more resources per patient day (if there is a tradeoff between intensity of treatment and length of stay) because they have full insurance at the margin for any additional resources used per

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negligible; for example, the 0.03 value for length-of-stay elasticity shown in Table 4 could perhaps be 0.04. See Newhouse and Taylor (1970) for a discussion of insurance effects on choice of hospital price.

[1]The average benefit was $29.44, so average payment was $29.44 + $4.00 = $33.44; 0.12 = $4/$33.44.

Table 5

HOSPITAL USE UNDER DIFFERENT COPAYMENT PLANS

$ Per Day Copayment	Effective Coinsurance Rate	Patient Days Per 1,000	Adjusted Benefits Per Case	Benefits Per Patient Day	Average Stay[a]	Admissions Per 1,000	Medical Resources Per Subscriber
0	0.0	1310	$261.00	31.11	8.39	156.14	40.75
4	0.12[b]	1129	$245.80	29.44	8.35	135.21	37.75[c]

SOURCE: Williams (1966). The data are from plan D, adjusted for age and sex. Coinsurance rates for plans A, B, and C are not available, and only unadjusted data are available for plan E.

[a]Computed as benefits per case/benefits per patient day.

[b]0.12 is $4 divided by the average benefit of $29.44 plus the $4 copayment.

[c]The medical resources per subscriber for the $4 per day copayment plan included $245.80 per case plus 8.35 × $4, the average payment by subscribers, or $279.20 per admission. Applying the admissions rate of 135.21 per 1,000 gives resources per subscriber of $37.75, as shown.

day. Those on the copayment plan indeed used $33.45 of resources per day, compared to $31.11 per day for those on the full payment plan. This is why the resource-use elasticity is less than the patient-day elasticity (0.04 vs. 0.07) in the actual range of copayment observed.

In another study, Rosenthal (1970) estimated the elasticity of length of stay with respect to several forms of price for hospital services, using data from 15,685 hospital admissions in New England in 1962. Of greatest interest here are his estimates of the elasticity that use the average coinsurance rate (cash payment/total bill) as the price variable.[1] The estimated coinsurance elasticities were all near zero; the largest elasticity in any of Rosenthal's 28 disease categories was 0.08. These estimates are likely to be inconsistent toward zero because the average charge per day falls with length of stay, as we pointed out in our discussion of length of stay coinsurance rates on page 24.[2] But even allowing for substantial inconsistency, Rosenthal's estimates are relatively small, as suggested by implication 1 (page 7).

In another study, Joseph (1972) purports to estimate elasticities for hospital length of stay from patient discharge data.[3] A feature of Joseph's study is that he holds diagnosis constant. Unfortunately, his only measure of insurance is a dummy variable measuring whether the individual had insurance or not; hence he measures differences in length of stay between insured and uninsured populations. He is able to control for age, sex, and type of hospital room, but these are imperfect controls for the numerous demographic and socioeconomic differences between insured and uninsured populations, which could be expected to affect length of stay and so bias his estimates. Perhaps

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[1] Rosenthal also presents estimates using average daily charges as a price variable. We do not discuss these estimates because price defined in this fashion does not hold constant hospital quality.

[2] Rosenthal's price variable is the average over the entire length of stay, rather than for the marginal day. Because of the concentration of ancillary service charges in initial days of a hospital stay, the average price will exceed the marginal price, and the difference is likely to narrow as length of stay increases. Thus there is a measurement error in Rosenthal's price variable that falls as marginal price falls with length of stay. This will lead to an inconsistency toward zero in the elasticity estimate (see Newhouse and Phelps, 1974b).

[3] Joseph (1972), pp. 152-161.

for this reason, in 7 of 22 diagnoses, the mean length of stay among the uninsured is greater than among the insured.

Joseph claims to measure elasticities, but actually his "elasticity" is the difference in length of stay between two groups divided by the mean length of stay in the uninsured population, i.e., $\Delta y/y_0$, where y_0 represents demand of those without insurance, y_1 is demand of the insured, and $\Delta y = y_1 - y_0$. Under an assumption which is known to be invalid, this figure is analogous to a point elasticity at the point of no insurance (C = 1), but nowhere else.[1] We thus conclude that Joseph's estimates cannot be considered as elasticities, and one cannot compute an elasticity from the data Joseph presents. There is no measure of price (or coinsurance) change in Joseph's data.

PRESCRIPTION DRUGS

A study using data from Windsor, Ontario, showed the effects of reimbursement insurance on demand for prescription drugs. In 1958, Prescription Services Incorporated (PSI) (sponsored by Ontario pharmacists) was founded, offering all prescriptions at zero coinsurance rate above a payment of 35 cents per prescription. Since the average price per prescription in that community was $3.78, this corresponds to an effective average coinsurance rate of about 0.09 (varying with the price of the prescription). Per capita utilization of those with the PSI plan was compared with utilization by a sample of people who had prescriptions filled in the community during the same year (June 1, 1962, to May 31, 1963). The comparison is shown in Table 6.

The implied arc-elasticity of the *expenditure* on prescriptions is 0.40 over the range 9 to 100 percent. Extrapolating linearly to zero coinsurance, the arc-elasticity of expenditures in the zero to 25 percent

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[1] If one makes the assumption that all insurance in force provides C = 0, then Joseph's estimate is analogous to a point elasticity at C = 1. However, many insurance policies in force provide less than full coverage. The argument may be expressed algebraically: If all insurance in force has C = 0, then $\Delta C = -1$ for persons with insurance, and a figure analogous to a point elasticity at C = 1 is given by $(\Delta y/y_0)/(\Delta C/C_0) = (\Delta y/y_0)/(-1/1) = -\Delta y/y_0$, which is what Joseph estimates. However, Joseph cannot know that $\Delta C = -1$ because he does not know the effective coinsurance of those with insurance.

range is 0.07.[1] There may have been considerable self-selection into this drug insurance program, in which case these figures are biased away from zero.

Table 6

PRESCRIPTION RATES WITH AND WITHOUT DRUG INSURANCE FOR ONE YEAR IN WINDSOR, ONTARIO

Item	Without Insurance[a]	With Insurance	With Insurance (age adjusted)
Prescriptions per person	2.19 (1.81-2.57)	4.08	4.20
Expenditures per person	$8.29 ($6.89-$9.69)	$16.48	$16.64
Mean price per prescription	$3.78 ($3.70-$3.86)	$4.03	$3.96

SOURCE: Greenlick and Darsky (1968), p. 2125.

[a]Data based on sample of community; 95 percent confidence limits in parentheses.

There also appears to have been a higher average price per prescription in the insured plan than for the community as a whole. Because the only payment is a 35-cent deductible, the marginal cost above that rate is zero; as a result, patients have no incentive to shop for lower priced prescriptions, and physicians have no incentive to use generic (rather than brand name) prescriptions. The average price per prescription with this type of prepayment is $4.03, which is over three standard error units from the mean price per prescription in the community at large. Under normal hypothesis testing, we

[1]To compute this, we pass a straight line through expenditures 16.64 (or 16.48) at 9 percent coinsurance and 8.29 at 100 percent coinsurance. We find the values on this line corresponding to zero and 25 percent coinsurance and compute an elasticity.

would reject the proposition that the average prices per prescription in the two plans were equal, with probability less than 0.01 of making a type 1 error. This provides some evidence on the effects of fixed-payment insurance on price selection. For a discussion of an application of this phenomenon to hospitalization insurance, see Newhouse and Taylor (1970).[1]

Other data exist that allow estimation of a money-price elasticity of demand for drugs from British National Health Services (NHS) experience. The data show the total number of prescriptions filled annually for 13 years under the NHS, and the price charged per prescription. The nominal price was varied from zero shillings to 2.5 shillings throughout the period involved. We converted out-of-pocket payments to a coinsurance rate (by dividing payments by total value

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[1]Some actuarial data on hospital prices are available that also support Newhouse and Taylor's hypothesis, although there are other potential explanations. The data give the average price of hospital selected by persons with different categories of per diem reimbursement insurance. Representative figures (for male employees) are shown below.

	Per Diem Payable Under Policy ($)				
	Under 8.51	8.51-10.50	10.51-12.50	12.51-19.50	Over 19.50
Average daily benefit payable under terms of policy (mean of range)	7.73	9.85	11.77	14.89	20.04
Average hospital per diem incurred	11.75	12.98	14.91	16.24	18.92

SOURCE: Gingery (1967), Table 1.

These figures show a strong positive correlation between the average daily benefit provided and the per diem of the hospital chosen by the patients, as suggested by Newhouse and Taylor. However, the correlation may be due to regional geographic price differentials, as well as selection of high-coverage insurance by those who planned on using high-cost hospital services. To analyze the effects fully, one would need a simultaneous equation model predicting both the hospital price chosen and the amount of insurance chosen. The preliminary results of estimating such a model are in Newhouse and Phelps (1974a) and support the Newhouse-Taylor effect.

of prescriptions for the year); the resulting coinsurance rate varied from 0.0 to 0.23. The fitted equation from these data (where y = real expenditure on prescription drugs) was

$$(4.3)\quad y = 63.36 - \underset{(t = 9.16)}{53.16} \cdot \begin{pmatrix}\text{coinsurance}\\ \text{rate}\end{pmatrix} + \underset{(t = 43.35)}{4.49} \begin{pmatrix}\text{annual time}\\ \text{trend}\end{pmatrix},^{1}$$

where $\bar{y} = 92.41$, $\bar{C} = 0.129$, $R^2 = 0.995$, Durbin-Watson = 1.60, and N = 15.

Using Eq. (4.3) at C = 0.0 and at C = 0.25, the estimated arc-elasticity at the average of the time trend (7.5) is 0.07, identical to the estimate in this range from the Windsor drug data.

A study of Medicaid recipients in Mississippi shows similar effects of coverage on demand for drugs (Smith and Garner, 1974). The study drew a random sample of Medicaid recipients from one town and monitored drug purchases for three months before Medicaid drug coverage was instituted. The data were obtained from a complete search of all pharmacies in the town, so underreporting should not be a problem. Data on drug expenses by the same persons were obtained directly from Medicaid files for the period after drugs were covered by Medicaid; hence, the data sources for the two periods differ. Because Medicaid files will not show any drug purchases paid for directly by recipients, the data-gathering procedure will likely, if anything, understate the change in demand due to Medicaid coverage. The data are shown in Table 7.

The arc-elasticity of resource use (expenditure) is 0.38, where the coinsurance range is from C = 1 to C = 0. Demand with full coverage is 2.24 times demand with no insurance. If one imputes a straight-line demand curve between these two points, the estimated arc-elasticity in the range of C = 0.25 to C = 0 is 0.07, identical to the results from previously discussed studies.

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[1] Expenditures were deflated by the general British price index (Central Statistical Office, 1965, 1971). Data on other variables are from Myers (1972).

Table 7

COMPARATIVE DRUG USE BEFORE AND AFTER MEDICAID COVERAGE OF DRUG EXPENSE

Item	3 Months Before Medicaid[a]	3 Months After Medicaid[b]
No. prescriptions per person	5.43	9.48
Expense per prescription	$3.58	$4.49
Expense per person	$18.96	$42.54
Sample size	(241)	(241)

SOURCE: Mickey C. Smith and Dewey D. Garner, "Effects of a Medicaid Program on Prescription Drug Availability and Acquisition," *Medical Care*, Vol. XII, No. 7, July 1974, pp. 571-581.

[a]Data obtained from records of all pharmacies in town of study.

[b]Data obtained from State of Mississippi Medicaid files.

DENTAL SERVICES

We have four sources of data on effects of dental insurance on demand. First, we have obtained insurance premium data (from one carrier) on coverage of limited amounts of dental services at 20 percent coinsurance and at zero coinsurance. Benefits include examination, prophylaxis, and X-rays not more frequently than every six months. These data were analyzed in a manner similar to the data on total premiums (see Appendix A for methodology), and the estimated arc-elasticity of demand in the zero to 20 percent coinsurance range was 0.13. Extrapolated to the zero to 25 percent coinsurance range, the figure is 0.16. Lest this be thought of as a "small" response, note that the increase in services is 30 percent as coinsurance falls from 20 percent to zero.

A second study is drawn from the medical care literature. A Teamster Comprehensive Care Plan in New York (TCCP) provided free comprehensive dental care to participants. The number of visits per per-

son rose from 2.5 for the year before insurance to 4.9 per person per year for the year after introduction (Morehead, Donaldson, and Zanes, 1971). This represents a 96 percent increase in visits when persons were shifted to full insurance from no insurance.

Self-selection probably accounted for at most a small amount of this difference, because persons who joined the dental plan had lower dental expenses in the previous year ($78 median) than those who did not join the plan ($105 median). The percentage of persons seeing a dentist was also lower among those who joined TCCP--43 percent had seen a dentist in the previous year, with average visits of 2.4 per person, compared with 52 percent seeing a dentist in the group who did not join TCCP, with 3.1 visits per year average. Thus, those with high dental utilization were not overrepresented among those who joined the plan.[1]

The change in demand from 2.5 to 4.9 visits per person implies an arc-elasticity over the C = 1 to C = 0 range of 0.32. If we fit a linear demand curve to those two points and compute an arc-elasticity in the zero to 25 percent coinsurance range, the estimated demand at C = 0.25 is 4.3 visits, and the arc-elasticity is 0.07.

A third estimate is available by comparison of the dental utilization of two groups. While this is not a direct comparison of the same group at two different coinsurance levels, the groups may be treated in a way that sheds some light on the effects of dental insurance on demand. We compare a group of insured persons in the New York Group Health Dental Insurance (GHDI) plan for 1958-1964 with demand in the U.S. population as a whole for 1964 (where dental insurance was practically nonexistent). The data allow direct comparison of use of four major dental services (examination, prophylaxis, filling, and extraction) by age of person. Since age is an important determinant of demand, the between-groups comparison must be standardized by age-mix of the populations to be meaningful.

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[1]It might be argued that those who joined the plan had a large number of dental problems because of previous underutilization. However, the discrepancy in prior utilization does not appear to be large enough to sustain this argument.

The GHDI data are drawn from Avnet (1967), reflecting dental use over a period from 1958 to 1964. The plan provided full coverage of virtually all dental services. The four services considered constitute 84 percent of all services rendered in the plan, and two-thirds of the services as weighted by the American Dental Association's Relative Value Schedule (RVS). Table 8 shows absolute levels of demand for the U.S. population in 1964 for the four services, by age group (National Center for Health Statistics, 1965), as well as demand for two insured groups in GHDI. The "Basic" group from GHDI is primarily persons automatically enrolled (group enrollment) through employment agreements. The "Voluntary" group consists of families given the option to enroll, who pay all or part of their premium. (The population from which the voluntary sample was drawn is slightly represented--3.5 percent--in the basic group.) Comparison of the basic group's demand with average demand in the U.S. population should show effects of insurance, with little or no adverse selection present, since enrollment was automatic. The voluntary group's demand shows additional effects of adverse selection on per-person program costs.

Table 9 shows service and age-specific demand for the two GHDI groups, each relative to the comparable U.S. population level of demand. Overall demand in the basic group is 80 percent higher in GHDI, weighting age categories by the U.S. population age distribution. The arc-elasticity is 0.29, computed over the coinsurance range of zero to 1.0. If a linear demand curve is fitted between these two points, the computed arc-elasticity between C = 0.25 and C = 0 is 0.06, quite close to the estimate from the Teamster's dental plan.

In the voluntary group demand is 180 percent higher than in the U.S. sample, showing additional effects of adverse selection on per-person program costs. A possible assessment of the effect of self-selection can be made if the voluntary group in GHDI is compared with the U.S. population. The arc-elasticity in the zero to 1.0 coinsurance range is 0.47, and the interpolation in the zero to 0.25 coinsurance gives an elasticity of 0.09.

The experience in the GHDI plan is largely replicated in a separate study of dental use under an insurance plan for employees of the

Table 8

DENTAL SERVICES PER PERSON/YEAR

Age	Fillings	Extractions	Cleaning	Examination	Sum[a]	Age Group as Percent of U.S. Population
All U.S. Population, 1964						
Under 5	0.12	0.01	0.05	0.16	0.34	0.115
5-14	0.78	0.22	0.33	0.51	1.73	0.205
15-24	0.90	0.31	0.22	0.39	1.82	0.145
25-44	0.73	0.31	0.30	0.36	1.70	0.244
45-64	0.50	0.28	0.25	0.28	1.31	0.202
65+	0.16	0.14	0.09	0.11	0.50	0.092
All ages	--	--	--	--	1.38[b]	1.00
GHDI--Basic Group						
Under 6	1.03	0.09	0.28	0.28	1.63	--
6-14	2.26	0.20	0.53	0.51	3.50	--
15-24	2.23	0.24	0.43	0.42	3.32	--
25-44	1.40	0.26	0.42	0.40	2.48	--
45-54	0.92	0.33	0.34	0.33	1.92	--
55+	0.49	0.21	0.24	0.22	1.16	--
All ages	--	--	--	--	2.49[b]	--
GHDI--Voluntary Group						
Under 6	1.21	0.17	0.34	0.36	2.08	--
6-14	3.17	0.21	0.77	0.74	4.89	--
15-24	3.35	0.23	0.64	0.61	4.83	--
25-44	2.27	0.31	0.62	0.60	3.80	--
45-54	1.80	0.39	0.66	0.60	3.45	--
55+	1.38	0.61	0.62	0.55	3.16	--
All ages	--	--	--	--	3.86[b]	--

SOURCE: For All U.S. Population data: National Center for Health Statistics, "Volume of Dental Visits, United States--July 1963-June 1964," NCHS Series 10, No. 23, *Vital and Health Statistics*, U.S. Department of Health, Education, and Welfare, Washington, D.C., October 1965, Tables 22, 26. For GHDI data: Helen Hershfield Avnet and Mata Kouvari Kisias, *Insured Dental Care*, Group Health Dental Insurance, Inc., New York, 1967, Tables 5.7 and 6.36.

[a]Visits summed directly since dental RVS values are nearly equal for all four services.

[b]Calculated by weighting age-specific visit sums by the corresponding proportion of the U.S. population in each age class.

Table 9

RATIO OF GHDI DEMAND TO U.S. DEMAND

Age	Fillings	Extractions	Cleaning	Examination	All Visits[a]
GHDI--Basic Group					
0-5[b]	8.58	9.00	5.6	1.75	4.79
6-14[b]	2.89	0.91	2.41	1.00	2.02
15-24	2.47	0.77	1.95	1.08	1.82
25-44	1.92	0.84	1.40	1.11	1.46
45-54[c]	1.08	1.18	1.36	1.18	1.47
55+[c]	1.63	1.50	2.67	2.00	2.32
All ages	--	--	--	--	1.80[d]
GHDI--Voluntary Group					
0-5[b]	10.08	17.0	6.80	2.25	6.12
6-14[b]	4.06	0.95	3.5	1.45	2.83
15-24	3.70	0.75	2.91	1.56	2.65
25-44	3.10	1.00	2.07	1.67	2.24
45-54[c]	3.60	1.39	2.64	2.14	2.63
55+[c]	8.63	4.36	6.89	5.00	6.32
All ages	--	--	--	--	2.80[d]

SOURCE: Derived from data in Table 8.

[a]Visits summed directly, since dental RVS values are nearly equal for all four services.

[b]Comparison age groups are 0-4 and 5-14 in U.S. data.

[c]Comparison age groups are 45-64 and 65+ in U.S. data.

[d]Calculated by dividing total visits for GHDI groups by U.S. total. See Table 8 for total visits by group: 1.80 = 2.49/1.38; 2.80 = 3.86/1.38.

Dentists' Supply Company (DSC) located in York, Pennsylvania (Grubb, 1964). That plan paid for 80 percent of all services after a $10 deductible; the deductible was waived for routine oral examinations; and receipt of an examination fulfilled the deductible for other services. (This clause was included to obtain maximum benefit from preventive services.) Table 10 shows levels of utilization for the four basic services (fillings, extractions, cleaning, and examination), and Table 11 shows ratios of demand to U.S. population levels for the same services and age categories. Holding constant the age distribution of

Table 10

SERVICES/PERSON-YEAR, DENTAL SUPPLY COMPANY INSURANCE PLAN--1959-1962

Age	Fillings	Extractions	Cleaning	Examinations	Sum
0-4	0.522	0.022	0.186	0.151	0.880
5-14	2.490	0.317	0.761	0.499	4.067
15-24	2.904	0.379	0.894	0.919	5.096
25-44	1.500	0.264	0.593	0.586	2.943
45-54	0.925	0.272	0.498	0.479	2.174
55-64	0.513	0.324	0.327	0.344	1.508
65+	0.726	0.313	0.394	0.544	1.977
All ages	--	--	--	--	2.946

SOURCE: Herbert J. Grubb, *A Statistical Analysis of the Dentists' Supply Company Dental Health Plan*, Continental Casualty Co., Chicago, January 1964 (Table 35).

Table 11

RATIO OF DENTAL SUPPLY COMPANY DEMAND TO U.S. DEMAND

Age	Fillings	Extractions	Cleaning	Examinations	Sum
0-4	4.35	2.20	3.72	0.94	2.59
5-14	3.19	1.44	3.46	0.98	2.35
15-24	3.23	1.22	4.06	2.36	2.80
25-44	2.05	0.85	1.98	1.63	1.73
45-64	1.44	1.06	1.65	1.47	1.41
65+	4.54	2.24	4.38	4.94	3.95
All ages	--	--	--	--	2.13[a]

SOURCE: Table 10 for DSC data; Table 8 for U.S. data.

[a]Calculated by dividing total DSC use (2.946) by U.S. total use (1.38): 2.13 = 2.946/1.38.

the population, demand was 2.13 times higher under the 20 percent coinsurance plan than for the U.S. population, the largest increases (as with GHDI) being for children and persons over 65. The data represent a three-year average utilization between 1959 and 1962. Note that the level of use is higher with only 80 percent coverage (20 percent coinsurance) and a deductible than was demand under the full coverage GHDI plan (compare Table 8 to Table 10). These differences may be due to a variety of factors (e.g., time costs, water fluoridation levels, eating habits, or the fact that DSC employees, being engaged in the manufacture of dental equipment and dentures, were more "dental conscious" than the average citizen).

Using the values from DSC plan and the U.S. data, one can compute an arc-elasticity of 0.54 over the range of C = 1 to C = 0.2.[1] Letting demand at C = 1 equal 1, fitting a linear demand curve to these two points, and extrapolating, the estimated levels of demand are 2.06 at C = 0.25 and 2.41 at C = 0.[2] Fitting an arc-elasticity between these two points gives an estimate of $\eta_{hC} = 0.08$ in our standardized range, quite close to the Teamsters' and GHDI figures obtained above.

Estimates based on the insurance premium data show a considerably larger elasticity than those from the three groups. This discrepancy may be due in part to our assumption of linear demand curves, because the data from the three groups represent much larger shifts in the coinsurance rate than do the premium data, or it may be due to adverse selection in the insurance premium data. To test sensitivity to the linearity assumption, we assumed a semi-logarithmic demand curve of the form $y = \exp(a + bC)$, or $\ln(y) = a + bC$. For this functional form, the elasticity falls with C. Using this form, the elasticity estimates in the 25 percent to zero coinsurance ranges for these data are 0.08 for the Teamsters' data (rather than 0.07); 0.08 for the GHDI data (rather than 0.06); and 0.13 for the DSC data (rather than 0.08).

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[1] We assume that all persons exceeded the $10 deductible or had it waived because of an examination. The average expense per person was $35.36 per year.

[2] Thus estimated demand is 141 percent higher at C = 0 than at C = 1.

Thus the assumption of a linear demand curve may account for part, but not all, of the discrepancy between our data sources. Applying the semi-logarithmic form to the voluntary enrollment group of the GHDI data gives an elasticity estimate of 0.17 in this range (rather than 0.08), which is virtually identical to the insurance premium data estimate, suggesting that it is a combination of the assumed functional form and adverse selection that creates the discrepancy. Adverse selection is likely to be a much greater problem for dental services than all other services because so few (5 percent of the population) have insurance for dental services (Health Insurance Institute, undated). Thus groups that apply for dental coverage may well tend to be those with higher than average expenditures. (This should not be a problem with the medical insurance data used above because group medical coverage is nearly universal.)

To summarize, the arc-elasticity of demand for basic dental services appears to be around 0.1 in the coinsurance range of 0.25 to zero. Relative demand for basic dental services is some 80 to 140 percent higher at full coverage than at no coverage.

AGGREGATION ACROSS SERVICES

One can derive an average elasticity for all services by aggregating across services using expenditure weights. As noted above, it is important to convert elasticities to similar ranges, and the range we have chosen is zero to 25 percent coinsurance because this range is proposed in almost all national health insurance bills and because this range corresponds roughly to the range found in much of our data.

Our purpose in this aggregation is to derive an independent estimate of the price elasticity for medical care services which may be compared with the estimate obtained from Eq. (4.1) using insurance premium data. Because the insurance premium data did not include dental services, we omit them in this aggregation.

The relative (normalized) budget shares of hospital services, physician visits (other than house calls), ambulatory ancillary services, physician house calls, and prescription drugs are 0.68, 0.14,

0.06, 0.01, and 0.11, respectively.[1] Using these budget shares and elasticity estimates of 0.07 for hospital services, 0.14 for physician services, 0.07 for ambulatory ancillary services, 0.35 for physician house calls, and 0.07 for prescription drugs, we obtain an aggregate arc-elasticity in the 25 percent to zero coinsurance range of 0.08, compared with the figure of 0.04 obtained from the insurance company data cited earlier. Given the approximations involved, this is a reasonably close agreement. In relative terms, an elasticity of 0.04 implies an 8 percent increase in demand as coinsurance falls from 25 percent to zero; an elasticity of 0.08 implies a 17 percent increase in demand. The discrepancy between our two estimates should be viewed against the elasticities of 0.5 or greater found in the economics literature cited above.

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[1]The estimates are from Newhouse, Phelps, and Schwartz (1974), using data from Cooper and Worthington (1972). House calls are estimated to be slightly over 1 percent of all physician expenditures and prescription drugs to be two-thirds of all expenditures on drugs, based on data from the 1963 Center for Health Administration Studies Survey. These budget shares would presumably change somewhat if uniform coverage were introduced, but that should not much affect our calculation. Our calculation assumes that the elasticity for physician services in a hospital is identical to that for other inpatient services. Our theoretical construct suggests that inpatient physician services are likely to be less elastic than outpatient services; however, the aggregate elasticity estimate is not sensitive to alternative plausible estimates of η_{hC} for inpatient physician services.

V. SEVERAL CAVEATS

As previously noted, the data we analyze are primarily micro-studies that unambiguously show *demand* effects. However, until long-run supply adjustments have been made, demand shift estimates will overstate utilization estimates because of inelastic supply. Expenditures, of course, may change more or less than utilization, depending on supply and demand elasticities. Although this point may seem obvious, confusion concerning it exists (U.S. Congress, 1971).

Our estimates of elasticity have obvious weaknesses stemming primarily from the paucity of data. We have data giving the cost of an insurance policy covering all medical services for a quite limited range of coinsurance rates. For individual services, we have data (typically) at only two levels of coinsurance; such data cannot show how the elasticity changes over coinsurance rates or price levels.

Our results use the assumption of linear demand curves (to extrapolate or interpolate results into the zero to 25 percent coinsurance range) while much of the economics literature uses constant elasticity demand curves. We do not know how sensitive the results in the literature are to this assumption; because our results generally come from observations of two points, we are not well equipped to test for differences in functional form.[1] However, we believe our results are accurate for the ranges given; because most of our data come from approximately the zero to 25 percent coinsurance range, the chance for large error from interpolation or extrapolation seems minimal. And in most cases where we have two or more estimates for the same service over different ranges of coinsurance (hospital and prescription drugs)

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[1] We have performed the following test of the linearity assumption. If one assumes that the true demand curve is of the form $\log(y) = a + b \cdot \text{Coinsurance}$, then one can compute an elasticity using relative demand between two points that can differ from the arc-elasticity using the linear demand curve assumption. We have made such calculations, and the results are not changed except for the estimates of dental demand discussed in the text, and for two estimates of drug expense (the semi-logarithmic form gave an elasticity in the zero to 25 percent coinsurance range of 0.10 rather than the 0.07 estimate from the linear demand curve assumption). The methods diverge most when ΔC is large.

there appears to be sufficient consistency between our estimates to support the assumption of a linear demand curve. (Dental services are something of an exception.) Our data also do not allow us to address the question of how the response to insurance changes as the size of loss changes (that is, how η_{hC} interacts with ℓ). One is tempted to suggest that the elasticity goes to zero as ℓ becomes large, since the quantity demanded (which appears in the denominator of the elasticity expression) always increases as ℓ increases. Parallel demand curves (each shifted outward for larger losses) have this feature of decreasing elasticity with larger losses.[1] Most cost estimates for catastrophic insurance have implicitly assumed that the coinsurance elasticity approaches zero for large losses. If the opposite is true, such proposals could be considerably more costly than present estimates indicate.

We have little evidence to suggest what the interaction is, if any, between income and the insurance elasticity--η_{hC}. Any interaction will obviously affect the distribution of benefits under a national health insurance program. Our theory suggests that those with lower time costs will be more sensitive to coinsurance changes than others, but the evidence on this issue is scanty.[2] We note in passing that if those with low time values *are* more sensitive to coinsurance, then introduction of full coverage (C = 0) will increase the share of medical services consumed by that group. If queues also increase because of supply restrictions, the increased time price will have its strongest effects on those who value their time most, reducing their share of resources consumed.

Policy formation based on inferences from such scanty data is obviously hazardous. A social experiment in health insurance has now begun to alleviate these data deficiencies.[3]

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[1] Kenneth Arrow suggests that, because of the increasing budget share element of the income effect involved in $\partial h/\partial p_h$, larger losses may in fact lead to larger price elasticities.

[2] See Phelps and Newhouse (1972) and Newhouse, Phelps, and Schwartz (1974) for a discussion.

[3] Newhouse (1974).

VI. CONCLUSIONS

The data sources presented here support our theory of demand for health care services. Those services with a relatively high time price exhibit relatively low coinsurance (or price) elasticities (implications 1 and 3, page 7). Services with a relatively high money-price to time-price ratio show considerably higher own-price elasticities (implications 2 and 3). Services with coinsurance rates near zero show low price elasticities (implications 1 and 5). Table 12 summarizes our results, and Table 13 restates the results of the three studies in economic journals.

In our introductory remarks, we noted that some persons feel coinsurance is irrelevant to decisions about consumption of medical services, because physicians make all the relevant expenditure choices. We feel that the results presented here are strong evidence against that hypothesis. Consistently, across a number of studies based on diverse data, coinsurance has been found to exert an impact on utilization. However, our estimates of the responsiveness of demand to coinsurance are not as large as a number of estimates found in the economics literature. Part of this discrepancy may be due to differences in the range of effective coinsurance between our study and the others. But as we pointed out earlier, the difference between our estimates and those in the economics literature could also be due to an upward inconsistency in those studies (Newhouse and Phelps, 1974b). Another proposed explanation of the discrepancy is that behavior is based on the average coverage in an area, so that our methods would underestimate the change in demand which would result if everyone's coverage were changed (Ginsburg and Manheim, 1973). While this argument cannot be dismissed a priori, there is no unambiguous evidence supporting it and some evidence to the contrary (Newhouse and Phelps, 1974b).

Our data may overstate the effects of coinsurance on demand, since there may be some self-selection involved in many of our data sources. Also, in a number of our data sources, transitory factors

Table 12

ARC-ELASTICITIES FROM VARIOUS DATA SOURCES

Data Source	Services Covered	Arc-Elasticity Over Actual Range	Actual Range of C	Arc-Elasticity[a] Over C=0 to C=0.25	Relative Demands at C=0 to C=0.25
Fitted insurance premiums from 4 companies	All hospital, physician, prescription drug	0.07	0.10 to 0.25	0.04	1.08
Palo Alto pre-payment plan	Office visits Home visits Ancillary services	0.14 0.35 0.07	0 to 0.25	0.14 0.35 0.07	1.33 2.08 1.15
Blue Cross Indemnity vs. full coverage	Hospital admissions Average stay Patient days	0.04 0.03 --	0 to 0.31[b] 0 to 0.57[b]	0.04 0.03 0.07	1.08 1.06 1.15
Blue Cross intra-plan comparison	Patient days Resource use	0.07 0.04	0 to 0.13[b] ($4 per day)	0.14 0.08	1.33 1.17
Rosenthal hospital study	Hospital average stay	0 to 0.08[c]	??	0 to 0.08[c]	1 to 1.17
Windsor, Ontario drug plan	Prescription drug expense	0.40	0.09 to 1.0[b]	0.07	1.15
British NHS time series data	Prescription drug expense	0.07[d]	0 to 0.23	0.07[d]	1.15
Mississippi Medicaid	Prescription drug expense	0.38	0 to 1.0	0.07	1.16
Dental Insurance premiums	Dental examination, prophylaxis, X-ray	0.13	0 to 0.2	0.16	1.38
Teamsters' Dental Plan	Comprehensive dental care	0.32	0 to 1.0	0.07	1.14
GHDI dental plan vs. U.S. population	Examination, cleaning, filling, extraction	0.29	0 to 1.0	0.07	1.15
Dentists' Supply Company	Examination, cleaning, filling, extraction	0.54	0.2 to 1.0	0.08	1.17

[a] Linear extrapolations or interpolations, where necessary.

[b] See text for derivation.

[c] Constant, not arc-elasticity.

[d] Computed from linear regression values of demand at C = 0 and C = 0.25.

Table 13

ELASTICITY ESTIMATES REPORTED IN ECONOMICS JOURNALS

Study	Services Covered	Approximate Point Elasticities	Range of Coinsurance
Feldstein state aggregates (time series of cross-sections, 1958-67)	Hospital days	0.67	Constant elasticity
Davis-Russell state aggregates (1969)	Hospital days Outpatient visits	0.5 1.0	Constant elasticity
Rosett-Huang 1960 Survey of Consumer Expenditure	All physician and hospital expenses	0.35 to 1.5	0.20 to 0.80

may distort our inferences. A complete analysis of the problem would incorporate a simultaneous equations model, estimating the amount of medical care purchased, holding constant such things as income, time cost, the severity of the illness, quality of provider of medical services (the marginal productivity g'(h)), and the price of substitute methods of recouping health lost through illness.[1] Until such estimates are completed, we offer these estimates of the *average* elasticities at coinsurance levels which we feel to be relevant to policy decisions in universal health insurance debates.

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[1] See Newhouse and Phelps (1974a), Newhouse and Phelps (forthcoming).

Appendix A

COMPARATIVE STATICS OF THE CONSUMPTION OF MEDICAL CARE-- A DECISION UNDER CERTAINTY

The model specifies that consumers derive utility from health services (H) and other goods and services (x), with H a random variable defined by $H = H_0 - \ell + g(h)$, where H_0 is the entering level of health, and g(h) is a given production function relating quantities of medical care (h) to the amount of health (H). We specify that the consumer acts as if he maximized utility, subject to a budget constraint, where market prices (p_x and p_h), income (I), the loss from health (ℓ), the opportunity wage (w), the entering level of health (H_0), the production function (g(h)), travel time to and waiting time for care (t), and the insurance policy parameter (C) are taken as *given* for this decision. The consumer is said to maximize utility, subject to a budget constraint:

$$\text{(A.1)} \qquad \underset{x,h,\lambda}{\text{Max}}\ Z = U(x,H) + \lambda(-I + p_x x + Cp_h h + R).$$

Income is said to be productive time (T) times the wage rate (w), and time spent receiving market medical care is said to be taken entirely from productive time.[1] If $I = w \cdot T$, and medical care uses t per unit of h, then $\partial I/\partial h = w \cdot \partial T/\partial h = -w \cdot t$. This relationship is used to compute the first-order conditions for maximizing utility with respect to x, h, and λ:

$$\text{(A.2a)} \qquad U_x + \lambda p_x = 0,$$

$$\text{(A.2b)} \qquad U_H g'(h) + \lambda(Cp_h + w \cdot t) = 0,$$

$$\text{(A.2c)} \qquad -I + p_x x + Cp_h h + R = 0.$$

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[1] Consumption of medical services is assumed not to alter the total amount of productive time available. To assume otherwise would not change the substance of this paper. For an analysis of this case, see Grossman (1972).

CHANGES IN INCOME

Any change in income (I) would alter this system in the following fashion:

$$U_{xx}\frac{\partial x}{\partial I} + U_{xH}g'(h)\frac{\partial h}{\partial I} + p_x\frac{\partial \lambda}{\partial I} = 0,$$

$$U_{xH}g'(h)\frac{\partial x}{\partial I} + U_{hh}\frac{\partial h}{\partial I} + NPH\frac{\partial \lambda}{\partial I} = 0, \tag{A.3}$$

$$p_x\frac{\partial x}{\partial I} + NPH\frac{\partial h}{\partial I} = -\frac{\partial R}{\partial I} + 1,$$

where the "net" price of health care (NPH) is defined by $NPH = Cp_h + w \cdot t$, and $U_{hh} = U_{HH}g'(h)^2 + U_H g''(h)$.[1] Cramer's Rule gives

$$\frac{\partial h}{\partial I} = \frac{\begin{vmatrix} U_{xx} & 0 & p_x \\ U_{xH}g'(h) & 0 & NPH \\ p_x & 1-\frac{\partial R}{\partial I} & 0 \end{vmatrix}}{|M|} = \left(1-\frac{\partial R}{\partial I}\right)\left(\frac{p_x U_{xH}g'(h) - NPH\, U_{xx}}{|M|}\right), \tag{A.4}$$

where $$|M| = -p_x^2 U_{hh} - (NPH)^2 U_{xx} + 2p_x NPH\, U_{xH}g'(h),$$

and $$\frac{\partial R}{\partial I} = (1-C)(1+\theta)\int_0^{\ell^*} p_h\frac{\partial h}{\partial I} f(\ell)d\ell;$$

that is, the premium reflects, on average, any changes in demand for h due solely to income changes.[2] If utility is indeed maximized, then $|M|$ is positive, and U_{xx}, U_{hh} are both negative. We make no a priori assignment of the sign of U_{xH}, although it may be "reasonable" to assume that U_{xH} is positive (that the marginal utility of x is higher if the level of health is higher).

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[1] If we had specified the effects of h on income in nonlinear fashion, second derivatives of $\partial I/\partial h$ would enter the total expression for U_{hh}.

[2] The premium R is defined on p. 3.

CHANGES IN "TOTAL PRICE" OF MEDICAL CARE

We have defined $NPH = Cp_h + w \cdot t$ as the "total price" facing the consumer. *Any* change in "total price" alters the solution to (A.2) in the following manner:

$$U_{xx}\frac{\partial x}{\partial NPH} + U_{xH}g'(h)\frac{\partial h}{\partial NPH} + p_x\frac{\partial \lambda}{\partial NPH} = 0, \tag{A.5a}$$

$$U_{xH}g'(h)\frac{\partial x}{\partial NPH} + U_{hh}\frac{\partial h}{\partial NPH} + NPH\frac{\partial \lambda}{\partial NPH} = -\lambda, \tag{A.5b}$$

$$p_x\frac{\partial x}{\partial NPH} + NPH\frac{\partial h}{\partial NPH} = -h - \frac{\partial R}{\partial NPH}, \tag{A.5c}$$

where $$\frac{\partial R}{\partial NPH} = \frac{\partial R}{\partial C}\cdot dC + \frac{\partial R}{\partial p_h}\cdot dp_h + \frac{\partial R}{\partial w}\cdot dw + \frac{\partial R}{\partial t}\,dt$$

reflects any changes in R as any (or all) components of NPH change.

The Cramer's Rule solution is[1]

$$\frac{\partial h}{\partial NPH} = \frac{\begin{vmatrix} U_{xx} & 0 & p_x \\ U_{xH}g'(h) & -\lambda & NPH \\ p_x & \left(-h - \frac{\partial R}{\partial NPH}\right) & 0 \end{vmatrix}}{|M|} = \left(-h - \frac{\partial R}{\partial NPH}\right)\frac{\partial h}{\partial I} + \frac{\lambda p_x^2}{|M|}. \tag{A.6}$$

Since this derivative is somewhat artificial (we will typically observe changes in market price, wage rates, coinsurance rates, or travel time rather than changes in the "total price" per se), we defer discussion of the "price effects" in this model until the comparative statics with respect to p_h itself are analyzed. Derivatives such as $\partial h/\partial NPH$ will be referred to as $\partial h/\partial TP$ when $\partial R/\partial NPH = 0$ is assumed.

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[1] At this point, and hereafter, $\partial R/\partial I = 0$ is assumed in all expressions involving $\partial x/\partial I$, $\partial h/\partial I$, and $\partial \lambda/\partial I$.

CHANGES IN THE MARKET PRICE OF MEDICAL CARE

If the market price of medical care changes, then the system of first-order conditions set forth in (A.2) changes in the following manner:

$$U_{xx}\frac{\partial x}{\partial p_h} + U_{xH}g'(h)\frac{\partial h}{\partial p_h} + p_x\frac{\partial \lambda}{\partial p_h} = 0, \tag{A.7a}$$

$$U_{xH}g'(h)\frac{\partial x}{\partial p_h} + U_{hh}\frac{\partial h}{\partial p_h} + NPH\frac{\partial \lambda}{\partial p_h} = C\lambda, \tag{A.7b}$$

$$p_x\frac{\partial x}{\partial p_h} + NPH\frac{\partial h}{\partial p_h} = -Ch - \frac{\partial R}{\partial p_h}. \tag{A.7c}$$

Using Cramer's Rule gives:

$$\frac{\partial h}{\partial p_h} = \frac{\begin{vmatrix} U_{xx} & 0 & p_x \\ U_{xH}g'(h) & -C\lambda & NPH \\ p_x & \left(-Ch - \frac{\partial R}{\partial p_h}\right) & 0 \end{vmatrix}}{|M|} = \left(-Ch - \frac{\partial R}{\partial p_h}\right)\frac{\partial h}{\partial I} + \frac{C\lambda p_x^2}{|M|}. \tag{A.8}$$

Ignoring effects of changes in R (which produce income effects on demand for h), we find by comparing (A.6) and (A.8) that

$$\frac{\partial h}{\partial p_h} = C \cdot \frac{\partial h}{\partial NPH} \quad (+ \text{ income effects as R changes}), \tag{A.9a}$$

which, in elasticity form, is

$$\eta_{hp_h} = \frac{\partial h}{\partial p_h}\cdot\frac{p_h}{h} = \frac{Cp_h}{h}\frac{\partial h}{\partial NPH} = \frac{Cp_h}{NPH}\frac{\partial h}{\partial NPH}\cdot\frac{NPH}{h} = \frac{Cp_h}{NPH}\eta_h^T, \tag{A.9b}$$

where η_h^T is the elasticity with respect to *total* price. Ignoring income effects due to changes in R, we see that the money-price elasticity is approximately the total-price elasticity weighted by $Cp_h/(Cp_h + w \cdot t)$.

CHANGES IN THE TIME PRICE OF MEDICAL CARE

Two changes are possible that would alter the time price of medical care--the opportunity wage w may change, or the time per unit of service t may change. These changes are seen to be approximately equal in elasticity form, as will be shown below. First, consider how the first-order conditions are altered if w is changed:[1]

$$U_{xx}\frac{\partial x}{\partial w} + U_{xH}g'(h)\frac{\partial h}{\partial w} + p_x\frac{\partial \lambda}{\partial w} = 0, \tag{A.10a}$$

$$U_{xH}g'(h)\frac{\partial x}{\partial w} + U_{hh}\frac{\partial h}{\partial w} + NPH\frac{\partial \lambda}{\partial w} = -t\lambda, \tag{A.10b}$$

$$p_x\frac{\partial x}{\partial w} + NPH\frac{\partial h}{\partial w} = -th - \frac{\partial R}{\partial w}; \tag{A.10c}$$

then

$$\frac{\partial h}{\partial w} = \frac{\begin{vmatrix} U_{xx} & 0 & p_x \\ U_{xH}g'(h) & -t\lambda & NPH \\ p_x & \left(-th - \frac{\partial R}{\partial w}\right) & 0 \end{vmatrix}}{|M|} = \left(-th - \frac{\partial R}{\partial w}\right)\frac{\partial h}{\partial I} + \frac{\lambda t p_x^2}{|M|}. \tag{A.11}$$

Ignoring the effects of wages (w) on R, this expression is just the total-price slopes $\partial h/\partial TP$ multiplied by t (the time per unit of service), which may be written in elasticity form as:

$$\eta_{hw} = \frac{\partial h}{\partial w}\cdot\frac{w}{h} \approx t\frac{\partial h}{\partial TP}\cdot\frac{w}{h} = \frac{w\cdot t}{Cp_h + w\cdot t}\cdot\frac{\partial h}{\partial TP}\cdot\frac{Cp_h + w\cdot t}{h} \tag{A.12}$$

$$= \frac{w\cdot t}{Cp_h + w\cdot t}\eta_h^T.$$

[1] We are deliberately ignoring changes in productive time (T) in response to changes in the wage rate w. In other words, these derivatives represent partial derivatives holding income (I) constant. The work-leisure decision is assumed to be made outside of this framework. Recall, in (A.10c), that $I = T_0 w = w\cdot(T_0 - th)$, so $\partial I/\partial w = -t\cdot h$, if T_0 does not change with w. We consider a compensated change, so $T_0\cdot dW$ is exactly offset by income changes, and no income effect is given.

The "approximately equals" sign is used because the effects of $\partial R/\partial w$ are deleted.

Corresponding changes in the "service time" t produce similar effects on demand:

$$U_{xx}\frac{\partial x}{\partial t} + U_{xH}g'(h)\frac{\partial h}{\partial t} + p_x\frac{\partial \lambda}{\partial t} = 0, \tag{A.13a}$$

$$U_{xH}g'(h)\frac{\partial x}{\partial t} + U_{hh}\frac{\partial h}{\partial t} + NPH\frac{\partial \lambda}{\partial t} = -w \cdot \lambda, \tag{A.13b}$$

$$p_x\frac{\partial x}{\partial t} + NPH\frac{\partial h}{\partial t} = -w \cdot h - \frac{\partial R}{\partial t}. \tag{A.13c}$$

The appropriate derivative is solved for as

$$\frac{\partial h}{\partial t} = \frac{\begin{vmatrix} U_{xx} & 0 & p_x \\ U_{xH}g'(h) & -w\lambda & NPH \\ p_x & \left(-wh - \frac{\partial R}{\partial t}\right) & 0 \end{vmatrix}}{|M|} = \left(-w \cdot h - \frac{\partial R}{\partial t}\right)\frac{\partial h}{\partial I} + \frac{w \cdot \lambda \cdot p_x^2}{|M|}. \tag{A.14}$$

Similar to changes in w, these derivatives (ignoring $\partial R/\partial t$) are seen to be simply w times the corresponding derivatives with respect to total price (TP), and the elasticity for h is

$$\eta_{ht} = \frac{\partial h}{\partial t} \cdot \frac{t}{h} \approx \frac{w \cdot t}{Cp_h + w \cdot t}\frac{\partial h}{\partial TP} \cdot \frac{TP}{h} = \frac{w \cdot t}{Cp_h + w \cdot t}\eta_h^T \approx \eta_{hw}, \tag{A.15}$$

where, again, the approximation is indicated because effects of $\partial R/\partial t$ have been deleted.

CHANGES IN THE HEALTH LOSS "ℓ"

The level of health (H) is considered random in this model, with the loss from health ℓ being drawn from a random distribution f(ℓ). The loss from health may be taken as exogenous in this model, and we can investigate how purchases of x and h change if ℓ changes (if a different drawing from f(ℓ) had occurred). Note that the premium R

does *not* change for each drawing of ℓ. That premium was written to reflect all possible losses initially and is *fixed* at the time ℓ is actually observed. No deceit or imperfect knowledge is implied by this phenomenon. We have simply structured one insurance policy to average over many events.

Noting that $H = H_0 - \ell + g(h)$, so $\partial H/\partial \ell = g'(h)\partial h/\partial \ell - 1$, the first-order system given in (A.2) changes with the loss ℓ in the following fashion:

$$U_{xx}\frac{\partial x}{\partial \ell} + U_{xH}g'(h)\frac{\partial h}{\partial \ell} + p_x\frac{\partial \lambda}{\partial \ell} = U_{xH}, \tag{A.16a}$$

$$U_{xH}g'(h)\frac{\partial x}{\partial \ell} + U_{hh}\frac{\partial h}{\partial \ell} + NPH\frac{\partial \lambda}{\partial \ell} = U_{HH}g'(h) = \frac{U_{hh} - U_H g'(h)}{g'(h)}, \tag{A.16b}$$

$$-p_x\frac{\partial x}{\partial \ell} + NPH\frac{\partial h}{\partial \ell} = 0. \tag{A.16c}$$

Whenever $g''(h)$ is zero, these derivatives may be reduced to

$$\frac{\partial x}{\partial \ell} = \frac{-NPH}{g'(h)}\frac{\partial x}{\partial I} < 0, \tag{A.17a}$$

$$\frac{\partial h}{\partial \ell} = \frac{+p_x}{g'(h)}\frac{\partial x}{\partial I} = \frac{1}{g'(h)} - \frac{NPH}{g'(h)}\frac{\partial h}{\partial I} > 0, \tag{A.17b}$$

$$\frac{\partial \lambda}{\partial \ell} = \frac{-NPH}{g'(h)}\frac{\partial \lambda}{\partial I} < 0, \tag{A.17c}$$

which clearly demonstrate that a loss from health can be readily translated into a normal income effect through use of the shadow price of health--$NPH/g'(h)$.[1]

CHANGES IN THE INSURANCE PARAMETER C

Just as we may investigate how demand would change if money prices, income, or time prices change, it is also possible to investigate how demand patterns shift if a different coverage ratio C has been chosen

[1]The derivation will be supplied by the authors upon request. Equation (A.17b) (right-hand side of equality sign) makes use of the Engel aggregation.

for the insurance plan. Note that the insurance decision is considered endogenous but that the ratio C has been chosen before the loss ℓ is observed, and may thus be considered fixed. These *partial* derivatives tell how demand would change if a different level of C had been chosen, *for whatever reason.*

The system of first-order conditions (A.2) changes as C changes, according to

$$U_{xx}\frac{\partial x}{\partial C} + U_{xH}g'(h)\frac{\partial h}{\partial C} + p_x\frac{\partial \lambda}{\partial C} = 0, \tag{A.18a}$$

$$U_{xH}g'(h)\frac{\partial x}{\partial C} + U_{hh}\frac{\partial h}{\partial C} + NPH\frac{\partial \lambda}{\partial C} = -\lambda p_h, \tag{A.18b}$$

$$-p_x\frac{\partial x}{\partial C} + NPH\frac{\partial h}{\partial C} = -R_c - p_h h, \tag{A.18c}$$

which gives the derivative

$$\frac{\partial h}{\partial C} = \frac{\begin{vmatrix} U_{xx} & 0 & p_x \\ U_{xH}g'(h) & -\lambda p_h & NPH \\ p_x & (-R_c - p_h h) & 0 \end{vmatrix}}{|M|} = (-R_c - p_h h)\frac{\partial h}{\partial I} + \frac{\lambda p_h p_x^2}{|M|}. \tag{A.19}$$

The derivative is seen to be simply p_h times the "total price derivative" (see A.6) plus income effects due to changes in R as C is changed. Momentarily ignoring those income effects due to R_c, the derivative for h may be converted into elasticity form as

$$\eta_{hC} = \frac{\partial h}{\partial C}\cdot\frac{C}{h} \approx \frac{Cp_h}{Cp_h + w\cdot t}\left(\frac{Cp_h + w\cdot t}{h}\frac{\partial h}{\partial TP}\right) = \frac{Cp_h}{Cp_h + w\cdot t}\eta_h^T. \tag{A.20}$$

The approximation in this expression is caused by omission of income effects due to R_c. The full expression is written

$$\eta_{hC} = \frac{Cp_h}{Cp_h + w\cdot t}\eta_h^T + \frac{C(-R_c)}{I}\frac{\partial h}{\partial I}\cdot\frac{I}{h} \tag{A.21a}$$

$$= \frac{Cp_h}{Cp_h + w \cdot t}\,\eta_h^T + \frac{C(-R_c)}{I}\,\eta_{hI}.$$

The second term on the right-hand side of (A.21a) is typically very small. Using observed premium data, for C = 0.20 to C = 0.25, ΔR = \$48.84.[1] Hence, $\Delta R/\Delta C$ = 48.84/-0.05 = -976.80; for a "typical" income of \$10,000,

$$C\,\frac{-R_c}{I} = 0.20\left(\frac{976.80}{10{,}000}\right) = 0.02.$$

If η_{hI} is (arbitrarily) set at η_{hI} = 0.10,[2] then from (A.21a):

(A.21b) $$\eta_{hC} = \frac{Cp_h}{Cp_h + w \cdot t}\,\eta_h^T + 0.002.$$

The empirically observed bias from using (A.20) rather than (A.21a) for η_{hC} is quite small.

Calculation of η_{hC} from Premium Data

The premium function is

(A.22) $$R = [1 + \theta]\left[\int_0^{\ell^*} (1 - C)p_h h\, f(\ell)\, d\ell + \int_{\ell^*}^{\infty} (1 - C)p_h h^*\, f(\ell)\, d\ell\right],$$

so

[1] \$48.84 is the change with a \$25 deductible. In the context of a public plan, R_c refers to the change in taxes with respect to the co-insurance rate. The data were used to derive Table 1.

[2] Newhouse and Phelps (1974a). See also Andersen and Benham (1970) for additional estimates.

$$\text{(A.23a)} \quad R_c = [1 + \theta]\left[\int_0^{\ell^*} \left(-p_h h + (1 - C)p_h \frac{\partial h}{\partial C} \right) f(\ell)\, d\ell \right.$$

$$\left. - p_h h^* \int_{\ell^*}^{\infty} f(\ell)\, d\ell \right].$$

Let $\int_{\ell^*}^{\infty} f(\ell)\, d\ell = Q(\ell^*)$. Then (A.23a) becomes

$$\text{(A.23b)} \quad R_c = (1 + \theta)\left[\int_0^{\ell^*} - p_h h \left(1 - \frac{(1 - C)}{C} \eta_{hC}\right) f(\ell)\, d\ell \right.$$

$$\left. - p_h h^* Q(\ell^*) \right].$$

Assuming η_{hC} is constant over all $\ell \leq \ell^*$ gives

$$\text{(A.24a)} \quad \frac{R_c}{R} = \frac{(1 + \theta)\left\{\left[\frac{(1 - C)}{C} \eta_{hC} - 1\right] \int_0^{\ell^*} p_h h\, f(\ell)\, d\ell - p_h h^* Q(\ell^*)\right\}}{(1 + \theta)\left\{\int_0^{\ell^*} (1 - C)p_h h\, f(\ell)\, d\ell + (1 - C)p_h h^* Q(\ell^*)\right\}}.$$

Now define $\int_0^{\ell^*} p_h h\, f(\ell)\, d\ell = p_h \bar{h}$, the average expense under h^*, and

$$\text{(A.24b)} \quad \frac{R_c}{R} = \frac{\left[\frac{(1 - C)}{C} \eta_{hC} - 1\right] p_h \bar{h} - p_h h^* Q(\ell^*)}{(1 - C)p_h \bar{h} + (1 - C)p_h h^* Q(\ell^*)}$$

$$= \frac{\frac{(1 - C)}{C} \eta_{hC} - (1 + \delta)}{(1 - C)(1 + \delta)}$$

where $\delta = Q(\ell^*)p_h h^*/p_h \bar{h}$. Thus we can solve for η_{hC} using values for $R_c = dR/dC$ and R:

$$\text{(A.25a)} \quad \eta_{hC} = (1 + \delta)\left[\frac{R_c}{R} (1 - C) + 1\right]\left[\frac{C}{1 - C}\right],$$

$$(A.25b) \qquad = (1 + \delta)\left[\eta_{RC} + \frac{C}{1 - C}\right]$$

where η_{RC} is the elasticity of R with respect to C.

Empirically, δ is small and may be dropped. Since $\delta = (p_h h^*/p_h \bar{h})Q(\ell^*)$, we can estimate δ from insurance data. For $p_h h^*$ = \$25,000 (as specified in the premium data we use), $p_h \bar{h} \approx$ \$500, and $Q(\ell^*) < 0.0005$.[1] Thus $\delta < (25000/500)(0.005) = 0.025$, so $1 \leq (1 + \delta) < 1.025 \approx 1$. In arc form (dropping δ) (A.25a) becomes

$$(A.25c) \qquad \bar{\eta}_{hC} = \frac{\Delta h}{\Delta C} \cdot \frac{\bar{C}}{\bar{h}} = \left[\frac{1 + (1 - \bar{C})\dfrac{\Delta R}{\bar{R}} \cdot \dfrac{1}{\Delta C}}{1 - \bar{C}}\right]\bar{C}.$$

From the regression of premiums on coinsurance rate (see text), we have computed a premium index of 1.34 for C = 0, and 0.92 for C = 0.25. Using these values,

$$\begin{aligned} R_1 &= 0.92, & C_1 &= 0.25, \\ R_2 &= 1.34, & C_2 &= 0.0, \\ \bar{R} &= 1.13, & \bar{C} &= 0.125, \\ \Delta R &= 0.42, & \Delta C &= -0.25, \end{aligned}$$

so

$$(A.25d) \qquad \bar{\eta}_{hC} = \left[\frac{1 + 0.875\left(\frac{0.42}{1.13}\right)\left(\frac{1}{-0.25}\right)}{0.875}\right](0.125)$$

$$= -0.043.$$

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[1] Data on $Q(\ell^*)$ from unpublished insurance files. Calculations are courtesy of Bridger M. Mitchell and are available upon request.

Appendix B

A NOTE ON A MULTI-GOOD PRODUCTION OF HEALTH

We have assumed in our model that there is one uniform "medical service" input into production of health, so that $g(h, t) = H$. But we observe many inputs into the production of health--office visits, house calls, hospital stays, drugs, laboratory examinations, and the like. If we write an n-factor production function as $H = g(h_1, t_1, \ldots, h_n, t_n)$, then there is a systematic relationship between observed price elasticities of demand for any two services (say, office visits and home visits). That relationship is the same as any derived demand for factors of production. Let η be the underlying price elasticity of demand for health (H), and σ_{ij} be the Allen partial elasticity of substitution in the production process between inputs h_i and h_j. Assume factor supply is perfectly elastic. Assume h_r and t_r are used in fixed proportions, and the production function is linear homogeneous in (h_r, t_r) composites for $r = 1, \ldots, n$. Then the total-price elasticity for h_i is[1]

$$\eta^{T}_{h_i p_i} = \omega_i(\sigma_{ii} - \eta) \tag{B.1}$$

where ω_i is the share of all medical expenditures attributable to goods of the ith composite (h_i, t_i).

In assessing the implicit time costs involved in office visits, we have assumed that the underlying total-price elasticity of demand for h_i (office visits) and h_j (home visits) is identical, or at least nearly so. One way for that to occur is for ω_i to be equal to ω_j *and* for σ_{ii} to equal σ_{jj}. The budget shares for home and office visits

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[1] Allen (1938), p. 508. We have assumed that the supply of (h, t) composites is perfectly elastic. The supply of time to the production of medical services is clearly not perfectly elastic over a sufficiently large range. However, for many medical services the use of time may be a sufficiently small part of the total amount of time that its shadow price may not change appreciably as consumption of medical services changes.

are obviously unequal. However, the relationship holds approximately if σ_{ii} is large relative to η and $\omega_i\sigma_{ii} \approx \omega_j\sigma_{jj}$. Whether this holds is unknown. To the extent that the equality of the underlying total-price elasticities is not met, our estimates of the value of time for an office visit are in error.

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