# THE NORMS HYPOTHESIS AND THE DEMAND FOR MEDICAL CARE

PREPARED UNDER A GRANT FROM THE U.S. DEPARTMENT
OF HEALTH, EDUCATION, AND WELFARE

JOSEPH P. NEWHOUSE, M. SUSAN MARQUIS

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# PREFACE

This report was written as part of the Rand Health Insurance Study, funded under a grant from the U.S. Department of Health, Education, and Welfare. It addresses a question raised concerning previous work on the demand for medical care services, including work done under the Health Insurance Study grant: How variance in demand for medical care across individuals having different insurance plans has been used to estimate how the demand for care would vary as insurance plans change. An objection to this procedure holds that an important part of the demand for medical care is influenced by physician norms, and that if an across-the-board change in health insurance were to occur (as for example with a national health insurance plan), those norms would change. In that case the actual change in demand would be greater than would be predicted from studies that looked at variations in utilization across individuals with the norms held constant. This report examines the evidence available concerning the objection.

Reports issued under the Health Insurance Study grant that estimate the change in demand from data collected at the individual level include:

R-976-OEO, *The Effects of Coinsurance on Demand for Physician Services,* Charles E. Phelps and Joseph P. Newhouse, June 1972.

R-1197-NC/OEO, *Price and Income Elasticities for Medical Care Services,* Joseph P. Newhouse and Charles E. Phelps, June 1974.

R-2157-HEW, *Dental Care Demand: Point Estimates and Implications for National Health Insurance,* Willard G. Manning and Charles E. Phelps, March 1978.

An earlier draft of this report was published in the *Journal of Human Resources Supplement,* 1978.

# SUMMARY

This report tests what we term the norms hypothesis—that physicians treat patients in accordance with the average or modal insurance coverage in an area. If the hypothesis is true, a demand equation estimated from a cross-section of patients with different insurance will understate the effect of health insurance legislation that changes the average coverage. Our results, however, are inconsistent with the norms hypothesis.

Our review of the single study in the literature purporting to support the norms hypothesis shows that its specification is suspect; and when a theoretically more appropriate specification is used, the results give no support to the norms hypothesis.

Two variants of the norms hypothesis are distinguished. In the stronger variant, demand for care is solely a function of the community's average insurance coverage. In the weaker variant, demand is a function of both the individual's own coverage and the average coverage in the community.

We have tested the stronger variant with experimental data from the first site of the experimental portion of the Health Insurance Study. The data show the utilization of families in one medium-sized city who were assigned to different insurance plans. According to the stronger variant, there should be no difference among the plan means because all the individuals would be treated by physicians having the same norms. Yet statistically and practically significant differences appear among the plan means. We therefore reject the stronger variant of the norms hypothesis.

Experimental data from one site cannot be used to test the weaker version of the hypothesis, but data from additional sites are not yet available. We have therefore used data from two surveys done by the National Opinion Research Center for the Center for Health Administration Studies of the University of Chicago to test the weaker variant. Our method was to enter a variable measuring average insurance coverage in the individual's community, in addition to individual insurance coverage, in a demand equation. The coefficient of the average coverage variable was always insignificant at conventional levels, but its standard error was sufficiently large that the power of the test is low. In short, these results do not permit firm rejection of the weaker version of the hypothesis, but neither do they give it any support.

A nationwide study of utilization and insurance, the Medical Care Expenditure Survey, has recently been completed. When data from that survey and data from additional sites from the Health Insurance Study are available, it will be possible to test the norms hypothesis more completely.

# ACKNOWLEDGMENTS

The authors are grateful to Willard Manning and Carl Morris for assistance in computing results for the experimental data. Lee Lillard, Uwe Reinhardt, and the participants in the Conference on the Economics of Physician and Patient Behavior made helpful comments on an earlier draft.

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# TABLES

# I. INTRODUCTION

This report discusses a potential bias that may result from using individual (micro) data rather than aggregated data to estimate the insurance elasticity of demand for medical care services.[1] Some argue that the elasticity estimated from individual data underpredicts the demand response to a general change in health insurance benefits (as, for example, in a national health insurance plan). Those holding this view believe not only that the physician's role in determining demand is critical, but also that the physician finds it too expensive to ascertain the insurance coverage of the individual patient. As a result, the physician is assumed to make treatment decisions based on the average or modal insurance coverage in an area rather than on the particular patient's insurance. Put another way, the level of a community's insurance coverage determines physician norms. The norms, in turn, are assumed to determine treatment decisions. Estimates of demand that do not consider how national health insurance might alter physician norms will then underpredict the resulting change in demand. We term this argument the norms hypothesis. Our results give it little support.

Several references to the norms hypothesis are found in the economics literature. For example, Paul Ginsburg has commented on a paper that uses household-level data to estimate demand elasticities as follows: "... data on individuals are used instead of data on families or data aggregated by state.[2] Compared to aggregate data, [these] data have the advantage of avoiding general aggregation problems ... and provide superior opportunities for accurately specifying price. However, ... a change in the price faced by a single individual should have different effects from an across-the-board change for all individuals in an area as community norms of health care change. National health insurance may approximate the latter model (change in utilization through norms) more closely." (Ginsburg, 1976, p. 313.) A similar argument had been put forward in earlier work of Ginsburg (Ginsburg and Manheim, 1973, n. 17), and Martin Feldstein has also hinted at such a thesis (Feldstein, 1974, pp. 387-388).

The norms argument has arisen in part because studies using aggregated data (i.e., state averages) (Feldstein, 1971, 1977; Davis and Russell, 1972) appear to show higher estimated elasticities than studies using data from individual households (Scitovsky and Snyder, 1972; Phelps and Newhouse, 1972, 1974; Newhouse and Phelps, 1974, 1976). The norms argument would imply this discrepancy; however, our analyses do not support the hypothesis that the quantity of care received by an individual is affected by the community's average insurance coverage.[3] In our concluding remark, we suggest that misspecification in studies that use aggregate data is an alternative explanation for the discrepancy.

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[1]Individual data might come from household surveys, for example.

[2]It is not clear why estimates using data from families are less objectionable than estimates using data from individuals.

[3]We do not test a variant of the norms hypothesis—that the physician bases his decisions on the average coverage of his patients rather than on the average of the community. The studies using aggregate data use the average of the community as an explanatory variable, however.

Before proceeding to test the norms hypothesis, we consider one other criticism of studies using household-level data. Feldstein (1974) has argued that elasticity estimates based on behavior of households before and just after an insurance change are too low because they take account of only the short-term response. In his own work, Feldstein estimates a dynamic model that permits long-term adjustments, and concludes that the long-run demand elasticities are approximately double the short-run elasticities (Feldstein, 1971, 1974, 1977).

The principal data that compare utilization among the same households before and after a change in insurance have been collected by Scitovsky and Snyder and analyzed by them (1972) and by Phelps and Newhouse (1972). These data compared usage of a group of Stanford University employees in 1966 and 1968; in 1966 the coinsurance rate had been zero, whereas in 1968 a 25-percent coinsurance rate was imposed. Utilization fell 24 percent after imposition of the coinsurance rate.

If Feldstein's estimated lag for hospital days were applicable to physician visits, the 24-percent decline from 1966 to 1968 would be only about half or less of the ultimate response. Scitovsky and McCall collected additional data on 1972 utilization, both for persons who had been in the 1966-1968 samples and for all persons in the plan in 1972. Comparing the utilization of either group in 1972 with utilization in 1968, they found little change (Scitovsky and McCall, 1977).[4] Thus, the data are consistent with the hypothesis of a very short lag in adjusting to equilibrium. At the end of this report we argue that Feldstein's equation is misspecified, and one can therefore not infer from his results that adjustment to equilibrium proceeds slowly.

[4]Perhaps because of catch-up or transitory demand, Scitovsky and McCall expected that the effect of coinsurance might be less (rather than more) in 1972. The data are not consistent with substantial catch-up demand, either.

## II. EXISTING EMPIRICAL SUPPORT FOR THE NORMS ARGUMENT

A study by Bennett (1975) is the only empirical study in the literature that attempts to test the norms hypothesis. Bennett's evidence consists of average length of stay for maternity patients enrolled in the Federal Employees Health Benefits Program high option in 1970; all of these persons had the same insurance plan. The Federal employees' mean length of stay across the nine Census regions is positively associated with the proportion of the entire region's population insured by Blue Cross.

Bennett's explanation is that the amount of Blue Cross coverage determines physician norms. He asserts that Blue Cross has "traditionally covered the cost of hospitalization for all maternity cases," and that the proportion covered by Blue Cross "is positively related to length of stay because insurance decreases out-of-pocket cost to the *average* individual" (emphasis added). Bennett concludes: "The relationship between professional standards of physician practice and health insurance can explain why many estimates of the cost of national programs such as Medicare and Medicaid have been less than the actual cost. A national insurance program dramatically raises the level of insurance coverage held by an average or representative individual, and this influences the definition of medical need."

Closer examination of Bennett's evidence lends a quite different perspective. First, the norms argument is an unlikely culprit for the misestimation of Medicare and Medicaid costs. In the early 1960s, there were no good estimates of demand elasticities; one could hardly claim that the Medicare-Medicaid cost estimates were based on data from instances in which norms were unchanged and failed to adjust for a change in norms. Second, Medicare and Medicaid only raised the average percentage of the hospital bill reimbursed from 82 to 88 percent, and for physician services from 38 to 44 percent (Cooper, Worthington, and Piro, 1974, Table 7).[1] Such changes scarcely seem "dramatic" enough to change physician norms.

But the main point concerns Bennett's finding of a relationship between the extent of Blue Cross coverage and Federal employee length of stay. Bennett is simply incorrect in asserting that Blue Cross plans traditionally cover maternity services. No comprehensive data on Blue Cross coverage exist, but data collected by Reed in early 1968 belie Bennett's statement.[2] Reed published detailed provisions of the most widely held group contract for each of the 75 Blue Cross plans (Reed, 1970). No data are available on contracts other than the most widely held ones; we have therefore made the assumption that the maternity coverage of all other contracts is, on average, the same as that of the modal contract. Table 1 shows the resulting pattern of maternity coverage.

By these estimates, a little more than a third of the individuals insured by Blue

[1]These values overstate the change from Medicare and Medicaid. They assume that the entire change in benefits for services between 1966 and 1967 was attributable to Medicare and Medicaid and none to improvements in private insurance.

[2]Bennett's own references in support of his statement are from 1947 and 1955 publications. We have been unable to check those references, but we doubt that Bennett is correct even for 1947 and 1955.

Table 1

BLUE CROSS COVERAGE OF MATERNITY SERVICES

| Type of Coverage of Maternity Services | Number of Plans | Percentage of Total Blue Cross Enrollment |
|---|---|---|
| No coverage | 29 | 35 |
| Dollar limit on maternity coverage | 22 | 37 |
| $ 50 | *3* | *5* |
| $ 75 | *3* | *3* |
| $ 80 | *8* | *18* |
| $100 | *4* | *8* |
| $120 | *1* | *< 0.1* |
| $125 | *1* | *2* |
| $150 | *1* | *1* |
| $200 | *1* | *1* |
| 6-or-more-day limit on coverage | 22 | 25 |
| Not specified | 2 | 3 |
| Total | 75 | 100 |

SOURCE: Calculated from data in Louis S. Reed and Willine Carr, *The Benefit Structure of Private Health Insurance*, 1968 (DHEW/SSA/ORS/Research Report 32), U.S. Government Printing Office, Washington, D.C., 1970, Tables 2-1 and 2-5.

Cross had no maternity coverage at all, and a roughly equivalent number had a dollar limit on benefits, generally of $100 or less. The average cost per hospital day in 1968 was $56, and the cost of a normal delivery would therefore almost certainly exceed $100.[3] As a result, nearly 75 percent of individuals covered by Blue Cross had no coverage for the marginal day. In contrast, those who were covered for stays of 6 days or more were mostly covered for the marginal day, because the average length of stay (in the Federal employees sample) was approximately 4 days.

We have recalculated Bennett's estimated equation for average length of stay in the Federal Employees Plan, eliminating his variable of percentage with Blue Cross coverage and substituting the theoretically more appropriate variable, percentage with Blue Cross coverage for the marginal day.[4] We obtained an estimated equation, which we compare with Bennett's[5] in Table 2.

In the revised specification, the variable representing the percentage of the region with Blue Cross coverage for the marginal day has a statistically insignifi-

[3]This is the adjusted cost figure from the American Hospital Association, *Guide Issue.*

[4]For the purpose of calculation, we assumed that all of those enrolled in plans wherein the largest group had no coverage or a dollar limit also had no coverage for the marginal day. If the largest group had a day limit, we assumed that all enrollees had full coverage for the marginal day. We then calculated the percentage of the civilian population having Blue Cross coverage that represented coverage for the marginal day. Our calculations ignore commercial insurance coverage, but so do Bennett's.

[5]We could not replicate Bennett's results by using his data, but our estimated coefficients and t-statistics were sufficiently close to his for his specification that the inferences he drew would not have changed had our estimates been used.

Table 2

ORIGINAL AND REVISED EQUATIONS FOR MATERNITY LENGTH OF STAY[a]

| | | | |
|---|---|---|---|
| Bennett[b] | | | |
| 2.92 + 0.03 | percentage with Blue Cross | + 0.23 beds/1000 | − 0.88 physicians/1000 |
| | (6.86) | (1.95) | (3.46) |
| $R^2$ = 0.94; d.f. = 5 | | | |
| Revised[c] | | | |
| 1.15 + 0.01 | percentage with Blue Cross for marginal day | + 0.56 beds/1000 | + 0.24 physicians/1000 |
| | (0.90) | (1.74) | (0.39) |
| $R^2$ = 0.44; d.f. = 5 | | | |

[a]Absolute value of t-statistics is shown in parentheses.

[b]Max D. Bennett, "Influence of Health Insurance on Patterns of Care: Maternity Hospitalization," *Inquiry*, Vol. 12, No. 1, March 1975, pp. 59-66.

[c]Joseph P. Newhouse and M. Susan Marquis, this study.

cant coefficient. Because this variable is theoretically more appropriate, we conclude that Bennett's effort to establish the norms hypothesis fails, and that the norms hypothesis remains without convincing empirical support in the literature.[6]

[6]It might be argued that these results indicate that coverage of nonmaternity services determines a norm that spills over to maternity services. This argument is not persuasive because (1) there is little variation across regions in hospital insurance coverage (as opposed to Blue Cross coverage), and there is no logical reason that a norm should ignore commercial insurance; (2) maternity length-of-stay decisions are frequently made by obstetricians, who are likely to know the properties of maternity coverage; (3) maternity coverage is widely known to differ from coverage for other services, and imputing the coverage of other services to maternity services is implausible.

# III. A TEST OF THE STRONG NORMS HYPOTHESIS

Two versions of the norms hypothesis can be distinguished. The stronger version holds that the appropriate insurance variable in an individual demand equation is the area's insurance coverage and that the individual's insurance plan is irrelevant. A weaker version of the hypothesis is that inclusion of only the individual's insurance plan underestimates the insurance elasticity of demand in response to a change in everyone's insurance, although the individual's insurance plan can explain some variation in demand. In this section, we will present data that are inconsistent with the stronger version of the norms hypothesis. The data collected by Scitovsky and McCall and cited above are not consistent with the stronger version. But even more powerful data can be brought to bear.

These data come from the experimental portion of Rand's Health Insurance Study. In this experiment, families were enrolled in one of several different insurance plans that varied the coinsurance rate and set an upper limit on out-of-pocket expenditures by the families—either 5, 10, or 15 percent of income, or $1000, whichever was less. Table 3 describes the plans; the design is described in detail in Newhouse (1974).

The results presented here come from experience in the first year in Dayton, Ohio, the first site in which families were enrolled. Because the families are all located in one moderate-sized Standard Metropolitan Statistical Area (1970 population 850,000), we assume that any variable measuring the area's insurance coverage would show no variation across the families. In such a case, according to the stronger version of the norms hypothesis, one should observe no differences in demand among different plans.

Before presenting our results, we describe the sample briefly. The population eligible to participate in the experiment is civilian, noninstitutionalized, and under age 62. Veterans with service-connected disabilities and the disabled population covered by Medicare are also excluded. Those in the lowest tertile of income distribution were mildly oversampled and those in the middle tertile were mildly undersampled; no corrections for this sampling pattern have been made in the results presented here. The upper tertile of the income distribution was truncated at a family income of $25,000 (1973 income), thus excluding families in the upper 5 percent of the income distribution. Those eligible to participate were selected at random within the Dayton metropolitan area, although geographical stratification was employed to ensure representation from all parts of the Dayton area.

The participating family is paid a lump sum equal to its worst case to prevent adverse selection at enrollment; thus, a family cannot be worse off financially by participating. In such circumstances few families refused offers to participate. Seven percent of those who received offers to participate declined, and an additional 3 percent left the experiment during the first year in Dayton. The results presented below do not include data on those 3 percent, or on children who were born or adopted during the year, or on families whose heads became separated or

Table 3

DISTRIBUTION OF PERSONS ACROSS INSURANCE PLANS IN DAYTON

| Coinsurance Rate[a] | Upper Limit on Annual Out-of-Pocket Family Expenditure (% of income)[b] | Number of Persons | |
|---|---|---|---|
| 0 | 0 | 276 | |
| 25 | 5 | 95 | 254 |
| 25 | 10 | 82 | |
| 25 | 15 | 77 | |
| 50 | 5 | 61 | 174 |
| 50 | 10 | 54 | |
| 50 | 15 | 59 | |
| 100 | 5 | 102 | 268 |
| 100 | 10 | 77 | |
| 100 | 15 | 89 | |
| 100[c] | $150 per person[d] | 95 | |
| Total | | 1067 | |

[a]The coinsurance rate is the fraction of the bill paid by the family.

[b]The upper limit is $1000 if 5, 10, or 15 percent of income would exceed $1000.

[c]In this plan the coinsurance rate applied only to outpatient services; inpatient services were free. In all other plans, the coinsurance rate applies equally to inpatient and outpatient services.

[d]This plan represents a $150 per-person-per-year deductible, with a $450 limit on family expenditure. Other plans apply the limit on out-of-pocket expenditure across all family members.

divorced during the year.[1] The distribution of persons in the analysis sample across insurance plans is shown in Table 3.

The experiment acts as the participant's insurance company, and the results discussed below come from data reported on insurance claim forms filed with the experiment. The participant assigns the benefits of his existing policy to the experiment, and uses only the experimental policy. Insurance is thus exogenous.

As Table 4 shows, the plans differ significantly in the annual expenditure rate for all nondental services[2] and for ambulatory physician services, irrespective of whether a simple analysis of variance model or an analysis of covariance model is used.[3] Total nondental expenditure includes expenditure for physicians, hospitals,

[1]The insurance plan of such families was modified when a "head split" occurred.

[2]Dental expenditures are not included because, in the first year at this site, adult dental services were covered only for those with a zero coinsurance rate. (For children, dental services were covered as in any other service, and for other sites and other years in Dayton, dental services for adults were covered as in any other service.)

[3]The regression equations (plan means and response surfaces) are not presented because they are not central for present purposes, and their release might alter responses yet to be observed.

Table 4

F-TESTS ON PLAN-RELATED VARIABLES

| Dependent Variable | Model: Analysis of Variance | Model: Analysis of Covariance[a] |
| --- | --- | --- |
| Log nondental medical care expenditure | $F_{10,1056} = 2.79$ | $F_{10,1030} = 3.85$ |
| Log ambulatory physician expenditure | $F_{10,1056} = 4.06$ | $F_{10,1030} = 5.33$ |
| $F_{.01(10,1000)} = 2.34$ | | |

[a]The 26 control variables in the regression in addition to 10 plan dummies include: AFDC Eligible (0,1); Black (0,1); Self-Reported Health Status at Time of Baseline: Excellent, Good, Fair, or Poor (3 dummy variables); Self-Reported Amount of Pain at Time of Baseline: Great, Some, Little, None (3 dummy variables); Self-Reported Amount of Worry Related to Health at Time of Baseline: Great, Some, Little, None (3 dummy variables); Separated or Divorced at Time of Baseline (0,1); Single but Not Separated or Divorced at Time of Baseline (0,1); Dummy = 1 if 13 or more years of education and adult; Dummy = 1 if no M.D. visits in previous year; Log of M.D. visits in previous year (if positive); Log Age × Dummy = 1 if less than 18 years; $(\text{Log Age})^2$ × Dummy = 1 if less than 18 years; Log Age × Dummy = 1 if 18 years or older; Log Age × Dummy = 1 if 18 years or older and female; Dummy = 1 if less than 18 years; Dummy = 1 if less than 18 years and female; Dummy = 1 if 18 years or older and female; Log Family Size; Log Average Income in 1972 and 1973 (in 1972 dollars); $(\text{Log Average 1972-1973 Income})^2$.

drugs and supplies, optometrists, psychologists, and certain other providers. The dependent variable in these regressions is the logarithm of annual expenditure plus $5, because the transformation removes the skewness in the residuals and makes the error term approximately normal. Generalized least squares estimators are used to fit the model because of intercorrelations of the disturbance among family members.[4]

Families were assigned to plans to maximize balance (i.e., to ensure that the distribution of families on each plan would resemble the distribution on every other plan). Given that demographic characteristics were balanced across plans, the control variates in the analysis of the covariance model serve principally to improve efficiency; there is little imbalance across the plans for which they might control. Not surprisingly, the F-values increase in the analysis of covariance model, because the covariates do increase efficiency.

[4]The model is

$$Y_{if} = X_{if}\beta + u_{if},$$

where i denotes the individual and f, the family. The disturbance $u_{if}$ is assumed to have the following structure:

$$E(u_{if}, u_{i'f'}) = \begin{cases} \sigma^2, & i = i', f = f', \\ \rho\sigma^2, & i \neq i', f = f', \\ 0, & \text{otherwise.} \end{cases}$$

the parameter $\rho$, the intrafamily correlation, is estimated by the maximum-likelihood method.

Statistically significant differences are not necessarily practically significant. Although we do not report response surfaces in this report for reasons explained in footnote 3 on page 7, the differences among the plans are practically significant. To give some idea of how much the plans differ one from another, we have computed the coefficient of variation among the (untransformed) plan means.[5] The coefficients of variation for total nondental and ambulatory physician expenditure plan means are both 0.38, indicating substantial difference in demand among the plans. If the transitory (catch-up) demand that was correlated with a plan existed in the first year, these coefficients of variation are overstated; the existence of such demand, however, is inconsistent with the stronger version of the norms hypothesis.

In sum, our results are certainly inconsistent with the stronger version of the norms hypothesis; according to that hypothesis there should be little, if any, difference among the plans in observed utilization, and yet the differences among the plans are both practically and statistically significant. Clearly, when ordering services, physicians are taking account of the individual patient's insurance coverage.

It might be argued that the differences we observe result from services under the control of the consumer, and those under the control of physicians are still governed by the norms hypothesis. Such an argument would imply that the consumer plays a more powerful role than those who espouse the norms hypothesis might concede; nonetheless, the consumer role cannot be rejected with these data. To test that hypothesis, we must use other data.

[5]The coefficient of variation among the estimated means in the analysis of covariance model would be practically identical because of the balance across plans.

# IV. THE WEAKER VERSION OF THE NORMS HYPOTHESIS

The weaker version of the norms hypothesis states only that the responsiveness to insurance coverage estimated by the experimental data within one site would underpredict the responsiveness of demand to changes in insurance induced by a national health insurance plan. The data reported above are inherently unable to test the weaker version of the norms hypothesis, precisely because they come entirely from one site, and data from other sites are not yet available.

In the absence of experimental data, we have tested the weaker version of the norms hypothesis with nonexperimental data, namely the 1963 and 1970 surveys from the Center for Health Administration Studies at the University of Chicago (CHAS). Our method is to estimate a demand equation similar to those found in other work by J. P. Newhouse (see Newhouse and Phelps, 1974, 1976), but include in each equation a measure of the average insurance coverage of all individuals in the primary sampling unit (PSU). (The PSU in these surveys was a single metropolitan area or a rural county; there were 72 such units.)[1] If the weaker version of the norms hypothesis is correct, the average coverage in the area should have a negative and significant sign, in addition to a negative sign for the individual coverage. Our results do not support these predictions.

Using the 1963 data, we have estimated the effect of the individual's coverage and the average coverage in the community in predicting the probability of a hospital admission and demand for physician office visits and hospital days conditional on some use of medical care. We have not included an equation relating the decision to use any physician services to the community norm. Although the physician may play a critical role in determining subsequent visits or in decisions about hospitalization, the initial decision to use outpatient services is usually made by the individual consumer rather than the physician. Hence, use of any physician services is unlikely to be affected by physician norms.

The measures of the individual's coverage and the average community coverage require some explanation before turning to the results. The presence of deductibles creates difficulty in defining a price variable except in the hospital length-of-stay equation. Because almost every hospital admission leads to a total expense greater than the deductible, the deductible should not influence the individual's decision about length of stay once he is admitted. By contrast, the amount of the deductible does affect the out-of-pocket price for an admission. Fortunately, deductibles for hospital admissions were relatively infrequent; only 10 percent of individuals had an insurance policy with a deductible for hospital care in 1963. Hence, for purposes of computing a value for the *community* coinsurance rate, we have ignored the deductible.[2] The community coinsurance rate variable in the

[1]The sample for the 1970 study was selected to overrepresent specific subgroups. One subsample was drawn from rural areas in PSUs selected especially for the 1970 study. Because only persons living in the rural areas of the PSUs were sampled, these PSUs have been excluded from our analysis.

[2]It could also be argued that deductibles should not be taken into account to test the norms hypothesis because of their infrequency; i.e., that the modal coverage does not include them.

hospital admission and length-of-stay equation is the average coinsurance rate for hospital room and board services (above a deductible) for all persons sampled in the geographic unit (primary sampling unit).[3] Communities do differ substantially in the level of insurance coverage for hospitalization; the coefficient of variation among community coinsurance rate means is .40 in the 1963 data.

A deductible will nonetheless influence individual admission decisions. We have therefore excluded individuals with deductibles when estimating the admissions equation; we have, however, kept these persons in the length-of-stay equation because the deductible should be irrelevant, as explained above.

A different approach is necessary to measure the appropriate insurance variable in the physician visits equation because, in 1963, deductibles were commonplace in insurance policies covering outpatient physician care; 64 percent of persons with insurance coverage for physician visits had a deductible to satisfy. Additionally, the deductible is unlikely to be met after the first, or even first several, doctor visits. In fact, in 1963 only 10 percent of those persons with deductibles who used physician services satisfied the deductible. Hence, many persons, despite insurance coverage, were required to pay the full charge for their outpatient care.

Our procedure to measure the *community* insurance coverage for outpatient physician care was first to compute a marginal coinsurance rate for each individual as of the end of 1963; if an individual had not met the deductible, his coinsurance rate was set equal to 1.0.[4] The community insurance coverage is then measured as an average of the marginal coinsurance rates; i.e., we assume that the community norm takes into account the expected proportion of persons satisfying the deductible. Two measures of the community insurance variable were tested in the physician visit equation. The first is simply the community average for all persons in the geographic area. The second is the community average for persons in the primary sampling unit who used services, because the persons the physician treats may establish the norm. The results obtained by using the two measures were qualitatively the same.[5] The former is used in the results below. The resulting variation across communities is considerably less than in the case of hospitals; the coefficient of variation is only 0.10. Although accounted for in the calculation of the community coinsurance rate, those with deductibles have been excluded from the sample because their inclusion would cause the estimated insurance elasticity to be biased away from zero (see below).

All of our calculations treat the insurance variables as exogenous. If anything, this would favor the norms hypothesis (because of adverse selection), but the results are nonetheless inconsistent with it.

The estimated hospital admissions equation is summarized in Table 5. The dependent variable is binary (one if an admission occurred, zero otherwise); we

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[3]For computational convenience, the coinsurance rate was set at 0.2 in this calculation for persons with a policy that contained a deductible; this is the modal coinsurance rate for such policies. The community coinsurance rate was calculated by using all persons with verified insurance; there were on average 70 individuals per community.

[4]This assumes that the physician and individual act as though they know with certainty that the deductible will not be exceeded (Keeler, Newhouse, and Phelps, 1977).

[5]Sample sizes were too small to allow us to test the alternative measure of community insurance coverage in the hospital equations, and the community insurance coverage is measured over all persons in the primary sampling unit, irrespective of whether the individual was admitted to the hospital.

employed a linear discriminant function as an estimator.[6] In the hospital admissions equation, the individual's insurance variable is negative and significant. The community insurance variable, however, is not statistically significant at conventional levels, which does not accord with the weaker version of the norms hypothesis.

Table 5

COEFFICIENTS ON OWN AND COMMUNITY COINSURANCE RATES FOR HOSPITAL ADMISSIONS, 1963 DATA[a]

| Explanatory Variable | Hospital Admission: Coefficient (t-statistic) | Hospital Admission: Elasticity (computed at the mean) |
|---|---|---|
| Own coinsurance rate | -.011 (5.11) | -.56 |
| Community coinsurance rate | -.006 (1.27) | -.30 |

[a]The number of observations is 3358. Excluded from the original sample are persons whose insurance policies are not verified; persons with more than three insurance policies; persons with reported wage income on imputed value of time exceeding $500 per week or less than $0; persons receiving free care or on welfare; persons with positive hospital or office deductibles; persons exceeding the limits of the insurance policy; persons who paid more than $50 per visit for office visits. The mean probability of admission is .059. The full equation is given in the Appendix.

The demand equations for length of stay conditional on an admission and the number of outpatient visits conditional on some use of physician services (based on 1963 data) are shown in Table 6. Although we do not find a significant effect of the average coinsurance rate in these equations, the coefficients on the individual's price are also not significant (and have the wrong sign). These findings do not support the norms hypothesis; however, the 1963 data have little power to measure moderately large effects.[7] Suppose average community coverage changed from no insurance to full insurance. If hospital length of stay and physician visits changed by 40 percent (about 3 days and two visits), the probability of a Type II error is 73 percent and 56 percent, respectively.[8]

[6]This procedure yeilds unbiased estimates of the parameters of the logistic function if the explanatory variables have a multivariate normal distribution; the significance tests are likely to be correct in any case. (See Halpern et al., 1971.)

[7]The lack of power is not due to a high correlation between the two insurance variables. The correlation between the community's average hospital coinsurance and the individual's own coinsurance is .28 for those admitted to the hospital; in the sample using physician services, the community outpatient coinsurance correlates .46 with the individual's own coverage.

[8]This calculation sets probability of Type I error at 5 percent (one-tail) and uses a normal approximation to the noncentral t-statistic. We have held constant the individual insurance variable for computational ease, implying that everyone's insurance except those in our sample changed.

Table 6

COEFFICIENTS ON OWN AND COMMUNITY COINSURANCE RATES FOR HOSPITAL LENGTH OF STAY AND PHYSICIAN OFFICE VISITS, 1963 DATA[a]

| | Dependent Variable | | | |
|---|---|---|---|---|
| | Log Hospital Length of Stay | | Log Physician Office Visits | |
| Explanatory Variable | Coefficient (t-statistic) | Elasticity (computed at the mean) | Coefficient (t-statistic) | Elasticity (computed at the mean) |
| Own coinsurance rate | .0003 (.22) | .009 | .0003 (.32) | .03 |
| Community coinsurance rate | -.0046 (1.41) | -.21 | -.0015 (.67) | -.14 |

[a]For hospital length of stay, n = 295; for physician office visits, n = 1682. Excluded from the same are persons with more than three insurance policies; persons with reported wage income or imputed value of time exceeding $500 per week or less than $0; persons receiving free care or on welfare; persons exceeding the limits of the insurance policy; persons who exceeded $50 per visit for office visits. The hospital length-of-stay equation excludes persons with no charge for a (nonobstetrical) hospital admission. The physician office visit equation excludes persons with no charge for physician outpatient care and with deductibles in their insurance policy. The full equations are given in the Appendix.

In light of the lack of power in the 1963 data, we have also tested the norms hypothesis using the 1970 survey data. Unlike the 1963 survey, the data collected in 1970 included relatively little detailed information on terms of the insurance policy. As a result, we do not know the coinsurance rate for persons who did not use services. We have therefore limited our estimates to demand conditional on some use of medical care, and the community coinsurance rates are defined as the average over persons who used the services.[9]

For those who used medical care, we must infer the coinsurance rate from survey reports of out-of-pocket and total expenditures. The formula used was

$$\text{coinsurance} = (\text{out-of-pocket expense} - \text{deductible}) / (\text{total expense} - \text{deductible}).$$

This computation of coinsurance rates biases insurance elasticity estimates away from zero if expenditures are aggregated across medical services having different

[9]As in the 1963 data, a coinsurance rate of 1.0 if the individual did not satisfy the policy deductible was used to determine the community norm in the demand-for-physician-visits equation. For the calculation of community coinsurance variables from the 1970 data, it was necessary to weight the data to correct for different probabilities of selection among sample observations. From the 72 PSUs in the NORC master sample, three subsamples were drawn for the 1970 study: a sample selected from segments designated as having a high proportion of low-income urban families (U sample); a sample selected from the remaining segments in the master sample (A sample); and a sample of families with low income or a member aged 66 or older obtained by screening households in all NORC segments (S sample). (See Andersen et al., 1972.) Our procedure for measuring the community coinsurance variable was to calculate the PSU mean by using observations from the U and A samples weighted by the differential probabilities of selection. The regressions, however, use observations (unweighted) from the three subsamples. Across communities, the coefficient of variation in the mean hospital coinsurance rate was 1.25 and .05 for physician office visit coverage.

kinds of insurance coverage and if the well-covered services are more expensive (Newhouse, Phelps, and Marquis, 1979). The bias is probably small in the hospital length-of-stay equation because most inpatient services are covered equally. The bias may be somewhat more serious in the office-visit equation. Outpatient preventive care is typically not covered by insurance, even if other outpatient services are, and the proportion of expenditure on preventive services should decrease as total expenditures rise.

Any bias from aggregation is reinforced by inclusion of persons with a deductible in the sample.[10] The coinsurance rate equals one if the deductible has not been satisfied and is less than one (typically 0.2) once the deductible is met. Thus, those consuming larger quantities of medical care will have a lower coinsurance rate, and the insurance variable is negatively correlated with the disturbance term. Such a correlation of price and the error term will bias the price coefficient away from zero. Phelps (1975), however, shows that the bias is probably small; although the estimated elasticity at the mean is large, the mean is near no insurance. When examined near full insurance, the estimated response of demand is close to that of other estimates not subject to this problem. If the bias in the individual coinsurance rate is small, the bias in the coefficient of the community insurance variable should also be small.

The results from the 1970 survey data are presented in Table 7. The individual's coverage shows a significant effect on both the length of hospital stay and the number of physician visits. By contrast, the average coverage in the community is

Table 7

COEFFICIENTS ON OWN AND COMMUNITY COINSURANCE RATES FOR HOSPITAL LENGTH OF STAY AND PHYSICIAN OFFICE VISITS, 1970 DATA[a]

| | Dependent Variable | | | |
|---|---|---|---|---|
| | Log Hospital Length of Stay | | Log Physician Office Visits | |
| Explanatory Variable | Coefficient (t-statistic) | Elasticity (computed at the mean) | Coefficient (t-statistic) | Elasticity (computed at the mean) |
| Own coinsurance rate | -.009 (3.35) | -.05 | -.010 (8.44) | -1.00 |
| Community coinsurance rate | -.005 (.96) | -.03 | -.002 (.41) | - .19 |

[a]The complete model is shown in the Appendix. For hospital days, n = 747; for physician office visits, n = 2812. Excluded from the sample are persons who received free medical care; persons with incomes over $75,000; persons with greater than 100 disability days.

[10]Phelps (1975) included individuals with deductibles because if they were excluded, there was little variation in price. To save time and expense, we have simply replicated his equation.

insignificant in both equations. The power is still not large, however. The probability of Type II error is 84 percent and 83 percent in the length-of-stay and physician-office-visit equations, respectively, using the test applied to the 1963 data.[11]

In sum, the results from the two cross-section surveys do not support the norms hypothesis. Unfortunately, the power of these tests using these data is not high. Therefore, we cannot firmly reject the norms hypothesis with these data, but neither can we support it.

[11]Again in the 1970 data, multicollinearity does not explain the lack of power. The correlation between the community hospital coverage and the individual coinsurance variable is .24; the community average outpatient coinsurance rate correlates .17 with the own coinsurance rate.

# V. ARE THE STUDIES USING AGGREGATE DATA INCONSISTENT WITH THOSE USING HOUSEHOLD DATA?

One of the reasons why the norms hypothesis arose was the apparent inconsistency between insurance elasticities found when using agggregate data (for example, data across states) and those found when using household-level data. Earlier work of Feldstein (1971) and Davis and Russell (1972) estimated that insurance elasticities for hospital days exceeded 0.5,[1] whereas the earlier work of Newhouse and Phelps (1974, 1976) estimated elasticities less than half that value. The norms hypothesis would, of course, imply such a difference, but there are other reasons for the discrepancy.

There are several possible misspecifications of models estimated with aggregate data (Newhouse, Phelps, and Marquis, 1979). The potential for aggregation bias is well known. Moreover, as pointed out above, a deductible that is exceeded during the hospital stay is relevant to a decision about admissions, but is of little relevance to length-of-stay decisions. Because studies based on aggregate data use an average coinsurance rate to define the price variable, this distinction in estimated insurance elasticities is lost; the same price variable is used in both admissions and length-of-stay equations. Further, use of the average coinsurance rate when a deductible is present and all users exceed it would cause the estimated elasticity to be biased away from zero.[2] Typically constant elasticity demand functions have been fitted, but there is evidence that elasticities fall with coinsurance (Phelps and Newhouse, 1974). Thus, the functional form may be misspecified. If the true functional form is not constant elasticity, care must be taken in determining that elasticities come from comparable ranges.

The magnitude of possible problems in estimates based on aggregate data may be seen in Feldstein (1977); he estimates a demand function for hospital days by using a time series of state cross-sections from 1959 to 1973. He has altered his earlier (1971) specification by adding a variable representing the quality of hospital services. (He also added data from the years 1968 to 1973.) Feldstein estimates the price elasticity of hospital days by using a constant elasticity demand function. In the earlier study with a similar specification, Feldstein found a price elasticity of hospital days of 0.55;[3] after adding the quality variable and 5 more years of data, the estimated price elasticity drops to 0.13.[4] The 0.13 value appears roughly equivalent to the estimates from household-level data, and so is inconsistent with the norms hypothesis. However, when Feldstein breaks his data into two subperiods (1959-1965 and 1966-1973), he estimates price elasticities of 0.58 and 0.29

[1]We give price elasticities in absolute value.

[2]Those who stay longer have a lower average coinsurance rate even though they could have the same marginal price.

[3]Feldstein (1971), Eqs. 29 and 30.

[4]Feldstein (1977), Table 1, Eqs. 1 and 2.

for the two subperiods, respectively.[5] Both are clearly above his estimated elasticity based on data from the entire period; such instability emphasizes potential specification problems. In sum, misspecification is at least as plausible an explanation of the difference in estimates between individual and aggregate level data as is the norms hypothesis.

[5]Ibid., Table 1, Eqs. 3, 4, 5, and 6.

# VI. SUMMARY AND CONCLUDING REMARK

We can find little support for the norms hypothesis. The one study in the literature that purports to support the hypothesis is open to serious question. The stronger version of the hypothesis—that physicians treat all patients according to a norm that is a function of the area's insurance coverage—is quite clearly rejected by Rand's Health Insurance Study data. The weaker version of the hypothesis—that individual level data understate insurance elasticities when everyone's insurance is changed because these data do not consider changed physician norms—is not clearly rejected by the data, but neither is it supported. The discrepancy between estimates based on aggregate data and those based on household level data can be plausibly explained as misspecification.

The National Center for Health Services Research and the National Center for Health Statistics are currently sponsoring the Medical Care Expenditure Survey, a nationwide probability sample survey. This survey plans to collect very complete information on medical care expenditure and insurance. When it becomes available in a few years, the data from this survey, together with Health Insurance Study data from other sites, should provide the best opportunity yet for testing the norms hypothesis.

# APPENDIX

The estimated demand equations for hospital admissions, length of hospital stay, and physician office visits based on the 1963 CHAS data are given in Tables A.1 and A.2. The dependent variable in the hospital admission equation is binary and takes the value 1 if the individual had a hospitalization during the year. The dependent variables in the equations predicting demand for services conditional on some use are the log of hospital length of stay and the log of the number of physician visits during the year. Hospital length of stay is the number of nonobstetrical hospital days weighted by the average price across the sample for the type of accommodation (a one, two, three, or more bed medical or surgical accommodation). Differences in the average price of the accommodations are assumed to reflect productivity differences; the weighting scheme is therefore assumed to

Table A.1

HOSPITAL ADMISSION EQUATION, 1963 DATA
(n = 3358)

| Explanatory Variable | Coefficient[a] | |
|---|---|---|
| Hospital coinsurance rate | -.011 | (5.11) |
| M.D. office coinsurance rate | -.004 | (1.13) |
| Community coinsurance rate | -.006 | (1.27) |
| Wage income/week (0 if no wage income) | .0009 | (.431) |
| Estimated value of time (0 if wage income > 0) | .005 | (1.52) |
| Nonwage income | .00003 | (.62) |
| Education of head 9-11 years | .028 | (.13) |
| Education of head 12 years | .074 | (.34) |
| Education of head 13-15 years | -.509 | (1.70) |
| Education of head 16 or more years | -.460 | (1.48) |
| Age | -.006 | (1.31) |
| Family size | -.004 | (.09) |
| Sex (= 1 if female) | -.053 | (.32) |
| Race ( = 1 if nonwhite) | -.486 | (2.11) |
| Health status good | .614 | (3.60) |
| Health status fair | 1.586 | (6.05) |
| Health status poor | 2.707 | (6.47) |
| M.D.s per 100,000 population ratio | -.002 | (.22) |
| $(\text{M.D.s per 100,000 population ratio})^2$ | -.00001 | (.52) |
| Beds per 1000 population ratio | -.050 | (.41) |
| $(\text{Beds per 1000 population ratio})^2$ | .013 | (1.07) |
| Rural dummy | -.297 | (1.54) |
| Constant | -1.82 | (2.47) |
| F | 6.68 | |
| d.f. | 22,3335 | |

[a]The t-statistics are shown in parentheses.

Table A.2

UTILIZATION EQUATIONS, 1963 DATA

| Explanatory Variable | Dependent Variable = Log Hospital Length of Stay[a] (n = 295) | | Dependent Variable = Log Physician Office Visits[a] (n = 1682) | |
|---|---|---|---|---|
| Hospital coinsurance rate | .0003 | (.22) | -.002 | (3.57) |
| M.D. office coinsurance rate | -.0007 | (.49) | .0003 | (.32) |
| Community coinsurance rate | -.0046 | (1.41) | -.0015 | (.67) |
| Wage income/week (0 if no wage income) | .0005 | (.35) | .0003 | (.48) |
| Estimated value of time (0 if wage income > 0) | .002 | (.74) | -.0005 | (.59) |
| Nonwage income (0 if > $3000) | -.00005 | (.69) | -.00009 | (3.00) |
| Nonwage income if > $3000 | .00002 | (1.60) | .000007 | (.34) |
| Dummy = 1 if nonwage income > $3000 | .080 | (.34) | .049 | (.31) |
| Education of head 9-11 years | -.040 | (.27) | .062 | (1.02) |
| Education of head 12 years | -.106 | (.77) | .050 | (.80) |
| Education of head 13-15 years | -.0003 | (.00) | .181 | (2.28) |
| Education of head 16 or more years | -.113 | (.54) | .086 | (1.03) |
| Age 7-17 years | -.004 | (.02) | -.243 | (3.41) |
| Age 18-24 years | .317 | (1.20) | -.117 | (1.25) |
| Age 25-34 years | .241 | (.91) | -.113 | (1.07) |
| Age 35-54 years | .774 | (2.93) | -.143 | (1.47) |
| Age 55-64 years | .893 | (2.73) | -.099 | (.86) |
| Age 65 or more years | .765 | (2.76) | -.009 | (.09) |
| Family size | -.034 | (.97) | -.037 | (2.72) |
| Sex (= 1 if female) | .117 | (1.03) | .063 | (1.34) |
| Race (= 1 if nonwhite) | .462 | (1.98) | -.077 | (.98) |
| Disability days | .003 | (3.29) | .004 | (6.61) |
| Health status good | -.069 | (.55) | .282 | (5.66) |
| Health status fair | .033 | (.20) | .644 | (9.19) |
| Health status poor | .189 | (.90) | 1.042 | (9.23) |
| M.D.s per 100,000 population ratio | .001 | (.25) | .001 | (.51) |
| (M.D.s per 100,000 population ratio)$^2$ | -.000006 | (.32) | -.000002 | (.24) |
| Beds per 1000 population ratio | -.075 | (1.04) | .022 | (.65) |
| (Beds per 1000 population ratio)$^2$ | .007 | (1.10) | -.002 | (.73) |
| Married | -.015 | (.09) | .055 | (.86) |
| Constant | 1.406 | | 1.092 | |
| $R^2$ | .34 | | .19 | |
| Adjusted $R^2$ | .27 | | .17 | |
| F | 4.56 | | 12.63 | |
| d.f. | 30,264 | | 30,1651 | |

[a]The t-statistics are shown in parentheses.

convert the quantities into efficiency units. Similarly, the dependent variable in the physician-office-visit equation is the number of visits weighted by the average price across the sample of the type of primary care provider (general practitioner, specialist, clinic).

The demand equations estimated from the 1970 CHAS data are shown in Table A.3; the dependent variables are the log of the number of hospital days and the log of the number of physician office visits. The specifications for the equations follow Phelps (1975), and a detailed discussion of the variables can be found there.

## Table A.3

### Utilization Equations, 1970 Data

| Explanatory Variable | Dependent Variable = Log Hospital Length of Stay[a] (n = 747) | | Dependent Variable = Log Physician Office Visits[a] (n = 2812) | |
|---|---|---|---|---|
| Hospital coinsurance rate | -.009 | (3.35) | (b) | |
| M.D. office coinsurance rate | (b) | | -010 | (8.44) |
| Community coinsurance rate | -.005 | (.96) | -.002 | (.41) |
| Wage income | -.018 | (.54) | --- | |
| Income | -.00001 | (1.76) | .000002 | (.74) |
| Education | -.006 | (.38) | -.003 | (.23) |
| Age | -.008 | (4.86) | .006 | (5.51) |
| Sex (= 1 if female) | -.116 | (1.56) | .038 | (1.15) |
| Race (= 1 if nonwhite) | .164 | (2.37) | -.097 | (2.52) |
| Disability days | .013 | (9.12) | .012 | (10.19) |
| Health status good | .040 | (.43) | .131 | (3.44) |
| Health status fair | .236 | (2.28) | .449 | (8.83) |
| Health status poor | .401 | (3.03) | .603 | (6.53) |
| Pain very often | .145 | (1.28) | .404 | (5.37) |
| Pain fairly often | .093 | (.89) | .351 | (5.60) |
| Pain occasionally | -.026 | (.30) | .156 | (4.30) |
| Welfare (1 = yes) | -.110 | (1.01) | -.043 | (.30) |
| Deceased | .196 | (1.01) | .401 | (1.69) |
| Dummy = 1 if under age 14 | --- | | .294 | (3.03) |
| Appointment delay | --- | | .005 | (.55) |
| Travel time × wage | --- | | .003 | (.88) |
| Wait time × wage | --- | | .001 | (1.16) |
| Constant | 1.431 | (11.04) | 1.469 | (.49) |
| $R^2$ | .29 | | .24 | |
| Adjusted $R^2$ | .28 | | .23 | |
| F | 17.79 | | 43.26 | |
| d.f. | 17,729 | | 20,2791 | |

[a] The t-statistics are shown in parentheses.

[b] Cross-price coinsurance rates were excluded because the cross-price coinsurance rate could not be calculated if the individual did not use the other service.

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