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COPAYMENT AND UTILIZATION OF PUBLICLY INSURED
HOSPITAL SERVICES IN SASKATCHEWAN:
AN EMPIRICAL ANALYSIS

By
John MacGregor Horne

A Dissertation
Submitted to the Faculty of Graduate Studies
of Carleton University in Partial Fulfillment
of the Requirements for the Degree of

Doctor of Philosophy
in Economics

Department of Economics
Carleton University
Ottawa, Ontario
February, 1978

The undersigned recommend to the Faculty of
Graduate Studies acceptance of the thesis "Copayment
and Utilization of Publicly Insured Hospital Services
in Saskatchewan: An Empirical Analysis" submitted by
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ABSTRACT

This study examines the effects of copayment on families' access to and use of inpatient hospital services. Saskatchewan is an appropriate setting for the study, since it is the only Canadian province where these effects can be isolated from the more general price effects associated with the introduction of public universal hospital insurance. The Saskatchewan Hospital Services Plan (SHSP), the first program of its kind in North America, provided first dollar coverage of hospital bills to virtually all residents of the province for twenty-one years prior to the introduction of user charges in 1968.

In Chapter One, the key provisions of the insurance program prior to and during the copayment period are discussed, as are other essential features of the Saskatchewan setting. In addition, the performance of the program in the pre-copayment period 1947-67 is briefly reviewed to provide a perspective on the public debate which attended the government's decision to implement per diem utilization fees as a "cost containment" device.

In Chapter Two, previous attempts to quantify price effects in hospital use are critically reviewed. After identifying a variety of conceptual and statistical problems in these studies, it is tentatively concluded that admission and discharge decisions are only weakly sensitive to variations in the point-of-service price. It is also made clear throughout the review that further research, and, in particular, disaggregate analyses of quality micro-data runs no risk of being redundant.

The micro-analytic framework employed to analyze the

Saskatchewan "copayment experiment" is presented and discussed in Chapter Three. The elements of this framework are set out in terms of the available data, the method by which these data are sampled, the representativeness of the sample, the techniques of data assembly and organization, and the model by which the data are analyzed. Key features of the model are: (a) a set of dependent variables which measure utilization in terms of the probability, volume, frequency, and duration of hospitalization; (b) a disaggregation scheme under which each of these dependent variables is in turn narrowly defined by the diagnostic and treatment status of the patient, the practice setting and specialty of the attending physician, and the size, location, and service capability of the hospital; and (c) a vector of independent variables which, in addition to the user price of inpatient care, includes family income and other socio-demographic factors hypothetically relevant to hospital use.

Based on multiple linear regression results presented in Chapters Four and Five, it is concluded that direct charges on inpatients, equivalent to a coinsurance rate of five to seven per cent, do not significantly reduce hospital utilization under a public universal hospital insurance program. Evidence of a 6.2 per cent decline in broadly defined hospitalization during the copayment period is traced to reduced lengths of stay in several long-stay diagnostic categories, but these reductions are not unambiguous price effects. Indeed, examination of time trends in utilization, together with analysis of a non-copay control group, strongly suggests that the declines are unrelated to copayment.

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For any errors or inadequacies which remain, I alone am responsible.

Winnipeg, Manitoba
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John M. Horne

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CHAPTER I

INTRODUCTION

For almost a decade now there has been an official consensus in Canada that ways must be found to control the growth of expenditures financed under public universal hospital insurance programs.¹ Among the many cost-containment proposals that have been considered is one that would require consumer-patients to directly participate in the financing of hospital care through the payment of user charges. The economics of "copayment" are exceedingly simple. If the demand for inpatient care is not perfectly price inelastic, then raising the point-of-service price from zero -- its modal value under the universal schemes -- to some positive amount will reduce bed utilization and total expenditure. On the other hand, if demand is perfectly price inelastic, utilization and total expenditure will remain unaffected, but costs will be shifted from

¹The first publicizing of the consensus came in November, 1968 when the nation's health minister announced the establishment of a Federal-Provincial Committee and Task Force on the Costs of Health Services. Their pre-occupation with the costs of institutional care, and, in particular, the costs of care provided by public general and allied special hospitals, is clearly evident in the Task Force Reports on the Costs of Health Services in Canada (3 volumes) (Ottawa: The Queen's Printer for the Department of National Health and Welfare, November 1969). Their abiding concern is well-founded. Over the twenty year period 1953-73, the operating costs of these hospitals increased at an average annual rate of 13.6 per cent, outstripping population growth by a factor of seven and GNP growth by a factor of two. See Louis A. Lefebvre, Public General and Allied Special Hospitals in Canada -- Historical Summary of Inputs and Utilization of Facilities 1953 to 1973 (Ottawa: Statistics Canada, July 1976).

public to private budgets in direct proportion to the coinsurance rate.² In either event governments are relieved of some financial responsibility.

Other considerations do however complicate the proposal. First, there can be no guarantee, as its protagonists seem to imply, that only "unnecessary" hospitalizations are price-sensitive. Equally, if not more probable, is a deterrent effect on those prospective patients whose ability to pay for care is relatively low and whose ability to benefit from care is relatively high. Obviously this would be at variance with the equity objective which ultimately differentiates public from private hospital insurance. Second, if there is no deterrent effect, copayment is little more than a tax on inpatients. This too would be difficult to justify since taxation of illness contradicts both the spirit and the logic of the public programs. As Evans observes "One could justify such a tax only by appeal to a 'benefits received' principle of public finance, but if that principle applied to health care then the government should never have intervened in this market."³ These potential "perversities" perhaps explain why most provincial governments remain wary of resorting to the price mechanism to secure program cost reductions. It remains a policy option, however, and as such commands serious empirical assessment. Hitherto this has been precluded by the absence of hard data.

² Assuming that a uniform copayment charge applies to each and every unit of care, the coinsurance rate is given by the ratio of the net (point-of-service) price to the gross (supply) price per unit of care.

³ R.G. Evans, Price Formation in the Market for Physician Services in Canada 1957-69 (Ottawa: Information Canada, 1973) p. 118.

The purposes of this study are: (a) to review and critically appraise previous empirical studies on "price effects" in hospital use; (b) to collect and assemble utilization data generated before and after the introduction of copayment in Saskatchewan, the only Canadian province where the effects of per diem utilization fees can be isolated from the more general price effects associated with the implementation of public universal hospital insurance;⁴ (c) to develop a micro-analytic framework for assessing price, income, and other relevant effects on families' access to and use of hospital services; and (d) to quantify and interpret these effects in the Saskatchewan data.

Saskatchewan: The Setting of the Problem

Implemented in April 1968 by the Liberal government of Premier Ross Thatcher, the copayment policy remained in effect until August 1971 when it was removed by the newly elected NDP government of Premier Alan Blakeney. During this period the Saskatchewan Hospital Services Plan (SHSP) authorized hospitals in the province to charge adult and child inpatients a utilization fee of \$2.50 per day for the first thirty days and \$1.50 per day thereafter, to a maximum of ninety consecutive days.⁵ Concurrently, the Medical Care Insurance Commission, the agency

⁴In British Columbia and Alberta, direct charges on inpatients have been government policy since the introduction of their universal programs in 1949 and 1958 respectively. As the text subsequently notes, there was no copayment provision in the Saskatchewan program for the first twenty-one years (1947-67) of its operation.

⁵In 1968, the \$2.50 utilization fee represented a coinsurance rate of 6.5 per cent, i.e., 6.5 per cent of the \$38.62 average total per diem cost of care in the province's public general and allied special hospitals. By 1971, the implicit coinsurance rate had fallen to 5.3 per

established in 1962 to administer the companion universal medical care insurance plan, authorized physicians in the province to charge their patients \$1.50 for each office visit and \$2.00 for each home, emergency and outpatient visit.⁶ These fees for inpatient care and physicians' services were collectively subject to an annual ceiling of \$180 per single adult or family, and were payable as incurred by all but a small percentage of the insured population.⁷

Point-of-service charges on inpatients were unprecedented in the twenty-one year history of SHSP. Indeed, since its inception on January 1, 1947, the Plan had entitled the insured inpatient to receive a wide range of hospital services at no direct cost. Services "freely" provided over the period 1947-67 included public ward or minimal accommodation, general nursing care, use of operating and case rooms, surgical dressings and casts, X-ray, laboratory and other diagnostic procedures, radio therapy, anaesthetic drugs and equipment,

cent due to an increase in average per diem operating costs to \$47.15. Ibid., p.53.

⁶The effects of these visit fees on families' utilization of physicians' services have been extensively analyzed in R.G. Beck, "An Analysis of the Demand for Physicians' Services in Saskatchewan," (unpublished doctoral dissertation, Department of Economics, University of Alberta, 1971). The relevant results from this study will be noted in subsequent chapters.

⁷Those exempted included: recipients of social assistance under the Saskatchewan Assistance Plan; War Veterans' Allowance Recipients; Indians on whose behalf the Federal Government paid the fees; patients transferred to general hospitals from Saskatchewan mental hospitals, tuberculosis sanatoria, and schools for the mentally retarded; patients admitted involuntarily to psychiatric wards of general hospitals; and newborns. Saskatchewan Hospital Services Plan, Annual Report 1968 (Regina: Queen's Printer 1968), p.75.

physiotherapy, most drugs in general use, and blood plasma.⁸ The costs of these services in Saskatchewan hospitals were insured without limit and without regard to age or pre-existing conditions providing they were judged medically necessary by the patient's attending physician. Per diem utilization fees represented a significant revision in the Plan's schedule of inpatient benefits, and the first departure from its long-standing policy of providing "complete" insurance against the costs of inpatient care.

Other basic features of the program remained unaffected by the copayment policy. Thus, participation in the program continued to be compulsory for all persons who had resided in the province for a specified period, with the exception of those who were covered under federal or other provincial programs, e.g., members of the Armed Forces and R.C.M.P., inmates of federal and provincial penitentiaries, and patients in provincial mental hospitals and tuberculosis sanatoria. From January 1, 1947 to June 30, 1958 the residence requirement was six months; it was reduced to three months on July 1, 1958 when, under the terms of the Hospital Insurance and Diagnostic Services Act, the province began to receive federal financial assistance for the program.⁹ Eligibility for coverage of hospital bills continued to be contingent upon prior payment of a hospitalization tax, not to be confused with the

⁸Among the inpatient services not insured by SHSP were the extra costs of private and semi-private ward accommodation (unless ordered by the attending physician), and the services of doctors, private duty nurses and other personnel not employed by the hospital. SHSP, Annual Reports 1947-67 (Regina: Queen's Printer).

⁹Canada, Statutes of Canada, 1957, 5-6, c.28. For a compact review of this legislation see Malcolm Taylor, "The National Hospital Insurance Act," University of Toronto Law Journal, Vol. XII, No.2, 1958, pp.300-303.

utilization fees. Unlike private insurance premiums, the tax was not based on actuarial principles. It involved no distinction between "good" and "bad" risks, since a basic purpose of the program has always been to spread these risks in a province-wide pool and thereby equalize the terms and conditions of coverage for all residents. Rather the tax was levied annually on a family basis, and varied only in relation to family size and the age limit of persons included as dependents for family tax assessments.¹⁰ Moreover, the tax was never intended to fully finance the costs of the program. Indeed, in 1947 proceeds from the tax defrayed only 50 per cent of total expenditure, the other half being financed from the province's general funds. Proceeds peaked at 55 per cent of total expenditure in 1949, and thereafter declined fairly steadily to 35 per cent of 1961 expenditure, and 12 per cent of 1971 expenditure.¹¹ Reliance on general funds to finance the program became complete on January 1, 1974 when the Blakeney government eliminated the hospitalization tax, leaving "permanent residence" the only condition of

¹⁰Various tax rates have prevailed since the Plan's inception. For example, in 1947-48 the annual tax was \$5 per person with a maximum of \$30 for a family consisting of father, mother, and dependent children under the age of twenty-one; in 1957-58 it was \$20 per adult and \$5 per dependent child under the age of eighteen to a maximum of \$5 per family; and from 1965 to 1971 it was \$24 per adult and \$48 per family with no tax for dependent children under the age of eighteen. Indians, and recipients of social assistance, War Veterans' Allowance, and the Old Age Assistance Pension comprised the majority of beneficiaries on whose behalf these taxes were paid by federal, provincial, or municipal governments. SHSP, op. cit.

¹¹ Ibid.

eligibility for coverage.¹²

A synopsis of the Plan's performance during the first two decades of its operation will be conducive to a keener perspective on the copayment policy and the public debate which preceded its introduction. It will also afford an opportunity to place the Saskatchewan setting -- its population and health care delivery system -- in sharper focus.

The population insured by SHSP changed in several notable dimensions over the period 1947-67. Reckoned in terms of the number of beneficiaries, it grew from 780,445 to 959,310, an overall increase of nearly twenty-three per cent.¹³ Three-quarters of this growth occurred between 1957 and 1967 when its numbers increased by almost 132,000.¹⁴ More significantly, there was substantial change in the age structure of

¹²On this date the government also eliminated the medical care premium of \$12 per single adult and \$24 per family, annual payment of which had been a condition of eligibility for coverage of doctors' bills under the Medicare program.

¹³SHSP, Annual Report 1967, p.59. These figures are based on slightly different methods of enumeration; the former is a count of all persons with any SHSP coverage between January 1 and August 31, 1947, while the latter is a count of all persons with SHSP coverage on June 30, 1967. Moreover, Treaty Indians and recipients of War Veterans' Allowance are excluded from the 1947 figure, but are included in the 1967 figure, since their insurance coverage was not SHSP's responsibility until July 1, 1958. On that date, they added 23,097 persons to the Plan's insured population and raised the Plan's coverage to 97.5 per cent of the (inter-censal estimate of the) provincial population. The Plan's coverage ranged from 92.1 per cent to 94.4 per cent of the total population between 1947 and 1957, and since 1966 has never fallen below 99.0 per cent. SHSP, Annual Reports 1958, 1972.

¹⁴Ibid. Over this eleven-year period inter-censal estimates of the provincial population show an increase of only 79,000, from 879,000 to 958,000. The difference is largely due to administrative changes which became effective July 1, 1958, viz., the above footnoted addition of Treaty Indians and War Veterans' Allowance recipients to SHSP's insured population, and the reduction of the waiting period for

the insured population. Paralleling post-war experience in the rest of Canada, the most pronounced shift occurred in the age group 65 years and over. Between 1947 and 1967, beneficiaries in this age group increased their number at a rate (66.2 per cent) three times faster than the overall average and, in consequence, increased their representation in the population from 7.0 per cent to 9.5 per cent. Concurrent with this aging of the population there was: a marked decrease in the proportion of beneficiaries in the age group 25-64 years; a marked increase in the proportion of beneficiaries in the age group 5-14 years; and little or no change in the proportion of beneficiaries in the age group 15-24 years and the group 4 years and under. These developments are detailed in the following table:

TABLE 1.1
AGE DISTRIBUTION OF SHSP BENEFICIARIES

Age Group	1947		1967*		Change % 1947-67
	No.	%	No.	%	
0 - 4	84,177	10.8	103,500	10.8	23.0
5 - 14	143,349	18.4	216,662	22.6	51.1
15 - 24	139,668	17.9	158,088	16.5	13.2
25 - 64	358,684	46.0	390,362	40.7	8.8
65+	54,567	7.0	90,698	9.5	66.6
All Ages	80,445	100.0	959,310	100.0	22.9

*Includes Indians and War Veterans' Allowance Recipients.
Source: SHSP, Annual Report 1967, p. 59.

It may be noted that the increase in the combined proportions of young persons (under 15) and old persons (65 and over) served to raise the "dependency ratio" from 36.2 per cent in 1947 to 42.9 per cent in 1967. Out-migration of relatively mobile working-age persons (15-64) in

newcomers to the province from six to three months.

partial response to labour-saving technological change within the province's agricultural sector was doubtless one of several factors underlying this trend.¹⁵

Steady reductions in the demand for agricultural labour find further reflection in the quite striking changes which occurred in the geographic distribution of the population. In particular, there was an acceleration of the rural to urban shift which began in the 1930's. Census estimates based on a very liberal definition of urban areas (i.e., unincorporated and incorporated cities, towns, and villages of 1000 and over) indicate that the urban population increased from 21.3 per cent of the 1941 population to 30.4 per cent of the 1951 population, an average annual increase of 0.82 per cent.¹⁶ However, by 1966 it had increased to 49.0 per cent on the strength of year-to-year increases averaging 1.03 per cent.¹⁷ These shifts notwithstanding, the rural-urban population ratio has always been well above the national average and,

¹⁵For argument and evidence that job opportunities in the mining, manufacturing and service industries did not expand at a rate sufficient to absorb the labour released by the agricultural sector, see J.C. Stabler, "Regional Development Theory and the Growth of the Canadian Prairie Region, 1870-1961," (unpublished doctoral dissertation, University of Utah, 1969).

¹⁶Statistics Canada, Perspective Canada (Ottawa: Information Canada, July 1974).

¹⁷Dominion Bureau of Statistics, 1966 Census of Canada, Population Age Groups I (1-10) (Ottawa: Queen's Printer, March 1968). Alternate estimates for the years 1951 and 1966 based on a more stringent definition of urban areas (population 10,000 or more) yield urban per centages of 24.1 (cf. 30.4) and 36.7 (cf. 49.0) respectively. SHSP, Annual Reports 1951 and 1966 (Regina: Queen's Printer).

indeed, well above the ratios in the other prairie provinces.¹⁸ In geo-demographic terms, therefore, the province has always been quite atypical of the rest of Canada.

This atypicality is also reflected in certain salient characteristics of the Saskatchewan hospital sector. For example, an abundant supply of general hospital beds has long been one of the distinguishing features of the province's health care delivery system. At the inception of the hospital insurance program, there were 163 institutions with an aggregate rated capacity of 4,229 beds.¹⁹ This represented 5.1 beds per thousand population, the highest in Canada at the time. Between 1947 and 1958, there was a small decline in the number of institutions but this was more than offset by increases in their average size from twenty-six to thirty-nine rated beds. Accordingly, the stock increased to 6,090 beds and the bed-population

¹⁸The following data are illustrative of these differentials;
Rural-Urban Population Ratio

	1951	1961	1971
Sask.	2.29:1	1.33:1	.89:1
Alberta	1.10:1	.58:1	.36:1
Manitoba	.79:1	.56:1	.44:1
Canada	.60:1	.43:1	.31:1

Sask./Alta.	2.1	2.3	2.5
Sask./Man.	2.9	2.4	2.5
Sask./Canada	3.8	3.1	2.9

Source: Statistics Canada, op.cit. See also Appendix A, Table A.3.

¹⁹Saskatchewan, Department of Public Health, Report of the Hospital Survey Committee (Regina: Queen's Printer, 1962), p.10.

ratio to 6.9.²⁰ It was during this period that the bulk of post-war hospital construction and renovation occurred, fostered to a considerable extent by the financial assistance provided by the federal government under the National Health Grants Programme introduced in 1948. Between 1958 and 1967, net additions to the bed stock were comparatively small, but the resultant ratio of 7.0 beds per thousand population was still the highest in Canada.²¹

The geographic distribution of general hospitals has been very much influenced by the settlement patterns of the population. Indeed, notwithstanding the historic prevalence of sparsely populated rural areas, many with densities as low as one person per square mile, it has been the case that nearly all individuals have resided within twenty miles of a general hospital. Thus, in 1948, with the exception of nine institutions located in urban areas, all general hospitals of less than 100 beds were located in small towns and villages and, indeed, 95 per cent of these hospitals had less than twenty-five beds. By 1967 the picture had changed very little, with 13 of the 144 institutions situated in urban areas, and with 92 rural institutions in the under twenty-five bed category.

The performance of the hospital insurance program during the period under consideration may be assessed in the light of the relevant expenditure, cost, and utilization data. During the first year of operation, total program expenditure was \$7.56 million, equivalent to \$9.68 per capita. Year to year increases of 12.5 per cent between 1947

²⁰ Ibid.

²¹ SHSP, Annual Report 1967.

and 1957 resulted in total expenditures of \$24.6 million by the end of its first decade of operation. Per capita spending in 1957 was \$29.66, reflecting an annual rate of growth only slightly lower than that of total expenditures. Over the second decade of its operation, total expenditures increased to \$63.5 million, an overall increase of \$39.0 million, and an average annual rate of growth of 10.0 per cent. Per capita expenditures increased to \$66.14, an average annual rate of growth of 8.4 per cent. These figures and their breakdown into administrative costs and payments to hospitals for insured care are presented in Table 1.2. It will be noted that administrative expense consumed a declining share of total expenditures, from 7.9 per cent in 1947 to 2.0 per cent in 1967.

TABLE 1.2

SASKATCHEWAN HOSPITAL SERVICES PLAN EXPENDITURE,
SELECTED YEARS

	Total Expense		Hospitalization Expense		Administration Expense	
	Amount	PerCapita	Amount	PerCapita	Amount	PerCapita
1947	\$ 7,560.80	\$ 9.68	\$ 6,963.30	\$ 8.92	\$ 597.50	\$0.76
1957	24,553.60	29.66	23,757.00	28.70	796.60	0.96
1967	63,499.90	66.14	62,152.40	64.79	1,297.50	1.35
%Change						
1947-57	12.5%	11.9%	13.1%	12.4%	2.9%	2.4%
1957-67	10.0	8.4	10.1	8.5	5.0	3.5
1947-67	11.2	10.1	11.6	10.4	4.0	2.9

Source: SHSP, Annual Reports, various issues.

The increase in payments to hospitals providing insured services reflects increases in both the volume and the unit costs of care. Measured in terms of patient-days per thousand, the volume of care increased markedly between 1947 and 1957, from 1,565 to 2,120, but from 1957 to 1967 it declined to 1,992. Annualized over the two decades, utilization change was a modest 1.2 per cent. More dramatic were the increases in the unit costs of care. Average payments per insured-patient-day rose from \$4.97 in 1947 to \$13.49 in 1957, and to \$28.69 in 1967. Over the two decades, unit costs rose at 9.2 per cent per annum though, significantly, the most rapid rate of increase occurred during the plan's first decade of operation. These figures clearly indicate that the major contributing factor to program cost escalation was the increase in average payment per patient-day rather than increases in utilization. It is arguable therefore that conditions on the supply side of the hospital market rather than demand shifts triggered by the insurance program were the most probable cause of the escalation in expenditures that occurred in the time leading up to the introduction of copayment.

Against this background of utilization and cost performance the Liberal government initiated a discussion of copayment policies. The argument advanced by the then Minister of Health, Honorable Gordon Grant, was very simple. Speaking to the draft legislation, he stated that "the amendment (to the Hospitalization Act) is directly related to the problem of rapidly rising hospital costs" and cited the fact that

"in the last two years the costs have risen by about \$5 million per year or close to ten per cent."²² Underlying his concern was a belief that the provision of services without direct charge to users necessarily results in "abuse of health privileges" and, in particular "the abuse of hospital services".²³ As the Minister stated, "This is a measure which is intended to control responsible [sic] demand."²⁴ The Minister's position rested upon two highly questionable assumptions. First, he presumed that direct charges would eliminate or at least reduce "unnecessary" use of facilities. Second, he presumed that the imposition of copayment would constitute neither a financial barrier to care, nor a financial burden on those who were admitted to hospital. There was no factual basis for these assumptions.

The opposition's response in the debate also rested on empirical issues. Mr. W. Smishek, the then N.D.P. health critic, argued that "deterrent charges tend to concentrate on the low income groups"²⁵ on the assumption that they are over-represented in the inpatient population. He also argued that the magnitude of the proposed utilization fee was non-trivial: "From the old age pensioner who now receives a cheque of \$75 and who happens to go to hospital for a period

²²Saskatchewan, Debates and Proceedings; Official Report, Session 1968, N.S. Vol. VI, part 2. First Session of the sixteenth legislature, p.1047.

²³ Ibid., p.1048.

²⁴ Ibid., p.1048.

²⁵ Ibid., p.1059.

of 30 days this Government is going to take his old age security cheque away."²⁶ Finally, he contended that abuse of the "free" care system was not attributable solely or even mainly to decisions by patients. Patients are admitted to and discharged from hospital by their physicians. As he stated, "The physician as well as the patient has a contributory responsibility to ensure that his services are wisely used."²⁷

The disputants in this debate were more at odds in opinion than in fact. Indeed, as is often the wont of the politician, they speculated on the facts where the facts were not available to speak for themselves. It is perhaps appropriate, therefore, to regard the copayment policy as an experiment in public policy -- an experiment designed, though not ideally, to generate many of the very facts upon which its successful advocacy or rebuttal depended. Systematic analysis of the data generated during this policy period is thus essential if future provincial, national, and international debates on user charges are to be less speculative than those of the past.

Outline of the Study

The plan of this study corresponds closely to the purposes outlined at the beginning of this chapter. Chapter 2 presents a critical review of previous empirical work on price effects in the hospital sector. It is argued that the accumulated evidence is

²⁶ Ibid., p.1053.

²⁷ Ibid., p.1058.

inconclusive. Chapter 3 sets out the elements of the micro-analytic framework within which the Saskatchewan experience will be examined. This includes a discussion of the available data, the sampling procedures, and the conceptual model and its component dependent and independent variables. A regression model is employed to generate estimates of the effects of copayment and selected socio-demographic control variables on families' use of inpatient facilities. The results of these statistical experiments are presented in Chapter 4 and 5. Chapter 4 contains an analysis of hospitalization probabilities before and after the introduction of copayment. These probabilities are generated from pooled time series and cross-section data on a large sample of "user" and "non-user" families. Chapter 5 contains an analysis of hospital utilization (volume, frequency, and duration) before and after copayment. These estimates are generated from pooled time series and cross-section data from relatively small sub-samples of user families. Chapter 6, the concluding chapter, presents a summary of the study, its main findings, and the implications for public policy.

CHAPTER II

PRICE EFFECTS IN HOSPITAL UTILIZATION: THE EVIDENCE FROM THE LITERATURE

A review of the extensive literature on factors affecting hospital use reveals that serious interest in the question of "price effects" is a comparatively recent phenomenon. Indeed, the two studies¹ widely regarded as pioneering attempts to quantify the price-use relationship were published in 1964, some thirty-five years after the (U.S.) Committee on the Costs of Medical Care inaugurated fundamental research in the economic and social aspects of personal health services.² Opportunities to appraise the relationship were of course severely limited over most of this period. They became apparent only after the dramatic post-war expansion in voluntary insurance and

¹Gerald D. Rosenthal, The Demand for General Hospital Facilities, (Chicago: American Hospital Association, 1964); and Paul J. Feldstein and Ruth M. Severson, "The Demand for Medical Care," Report of the Commission on Cost of Medical Care, Vol. I, (Chicago: American Medical Association, 1964).

²The seminal studies are: Isodore, S. Falk, Margaret C. Klem, and Nathan Sinai, The Incidence of Illness and the Receipt and Costs of Medical Care Among Representative Families: Experiences in Twelve Consecutive Months During 1928-31, Publication of the Committee on the Costs of Medical Care, Report 26 (Chicago: University of Chicago Press, 1933); and Isodore S. Falk, Rufus Rorem, and Martha D. Ring, The Costs of Medical Care (Chicago: University of Chicago Press, 1933).

prepayment arrangements, and the consequent accumulation and organization of the requisite price-use data.³ However, health economists and other social scientists oriented to health policy issues have wasted little time analyzing these data as they have become available. Spurred on by the intense debate over national health insurance in the United States, they have, over the last ten to twelve years, added no fewer than a dozen studies to the once meager literature on empirical price effects.

These studies, all of which employ U.S. data, generally indicate that such effects do indeed exist, and, in particular, that a higher net price reduces the consumption of hospital care. They also indicate that these effects are typically quite small, certainly well within the inelastic range where the ratio of proportionate change in use to proportionate change in price is less than one. Nevertheless,

³Data limitations did not, however, preclude speculation about the effects of hospitalization insurance on utilization. Within the private insurance industry, the belief was (and remains) widespread that deductibles, co-insurance, copayment, and other benefit limitations were necessary to prevent over-utilization and abuse of coverage. Outside the insurance industry, and especially among hospital administrators and members of the then very small research community, it was conventional wisdom that the demand for hospital care was perfectly price inelastic. Early (pre-1964) efforts to appraise these conflicting viewpoints were quite primitive, at least by today's standards. Surveys conducted in the mid-fifties by Anderson and others indicated that admission rates were higher among the insured than the uninsured, but more convincing support of the insurance industry position had to await the development and testing of more rigorous research designs which avoided the assumption, implicit in the survey results, that the comparison groups were similar in all respects other than insurance coverage. See Odin W. Anderson and Jacob J. Feldman, Family Medical Costs and Voluntary Health Insurance: A Nationwide Survey (New York: McGraw Hill Book Co., 1956). Subsequent, and slightly more sophisticated, studies also revealed greater utilization among insured individuals. See Milton I. Roemer and Max Shain, Hospital Utilization Under Insurance (Chicago: American Hospital Association, 1959); and Burton A. Weisbrod and Robert J. Fiesler, "Hospital Insurance and Hospital Utilization," American Economic Review, XLI (March 1961), pp. 126-143.

there are some significant inter-study differences³ in specific results which bear close scrutiny. These previous research efforts and the methodological details and problems pertinent to their appraisal will be selectively reviewed in this chapter.

Attention focuses first on three studies which exhibit some essential similarities in research strategy and design. The dominant common property is that of aggregation, specifically aggregation over elemental consumption units (i.e., patients, families, or households). Each employs state level data on hospital use and each asks the question "are there inter-state variations in hospital use that can be ascribed to inter-state variations in the price of inpatient care?" Moreover, each attempts to answer the question using multi-variate regression techniques in order to control for inter-state differences in population structure and other arguably relevant determinants of hospital use.

The prototypical study in this group is Rosenthal's The Demand for General Hospital Facilities.⁴ Following text-book theory, Rosenthal hypothesizes that the demand for inpatient care, like that for any other commodity or service, is a function of "tastes and preferences as well as price and income."⁵ He then develops a statistical model to test this hypothesis with cross-state data for 1950 and 1960. The dependent variable is defined as general hospital use in a given state during a given year, and is alternatively measured in patient-days per thousand population, admissions per thousand population, and average length of stay per admission. The independent variables fall into three

⁴Rosenthal, op. cit.

⁵Ibid., p. 20.

categories as indicated by the theory. With respect to "tastes and preferences", Rosenthal assumes that "the socio-demographic characteristics of an area reflect most of these tastes and preferences in a systematic way"⁶ and, shortly thereafter, admits that "particular variables were selected on the basis of their popularity in the literature."⁷ These include: age (proportions of the population over 64 and under 15 years of age); sex (proportion of the male population age 14 and over); marital status (proportion of the female population age 14 and over who are married); urbanization (proportion of the population residing in urban areas); education (proportion of the population over 25 with thirteen or more years of schooling); race (proportion of the population that is non-white); and housing (population per dwelling unit).⁸ As for income (proportions of the population with annual incomes under \$2,000 and over \$5,999), it is expected, again on the basis of earlier research, that both low and high income groups would use more care than middle income groups. Finally, and most importantly, the model incorporates two "price" variables, namely, the average (most frequent) per diem charge for a two-bed room across all general hospitals in a given state, and the proportion of the population covered by some form of hospital insurance. The expected signs on these variables are negative and positive respectively. That is, higher use is anticipated the lower the room charge, and the higher the proportion

⁶Ibid.

⁷Ibid., p. 23.

⁸The expected influence of each of these socio-demographic variables is discussed by Rosenthal, but need not be stated here, since the immediate concern is with his argument and evidence respecting price effects.

of the population covered by insurance.

For each of the three definitions of use, an equation in these twelve independent variables is fitted to the state data for 1950, 1960, and 1950-60 combined using the ordinary least-squares regression method. In brief, the results of this experiment indicate that: (1) the most important socio-demographic factors are "marital status" and "housing". Both variables - the proportion of the state's female population who are married, and population per dwelling unit respectively -- generally carry significant negative coefficients, suggesting that home care is an effective alternative to hospital care when domestic resources are available; (2) the least important socio-demographic factor is "age", possibly because, as Rosenthal suggests, states with a high proportion of aged may have a correspondingly high stock of alternative facilities, such as nursing homes;⁹ (3) states with a relatively high proportion of low income families tend to be low use areas, a paradoxical result which may reflect inter-state differences in bed availability that are not accounted for in the model; (4) states with a high average room rate and/or a low prevalence of hospital insurance tend to be low use areas, especially in 1960 when the "price" coefficients are significant for all three definitions of use. Estimated room-charge elasticities for patient-days, admissions, and mean stay are -0.65, -0.46, and -0.22 respectively, while estimated "insurance" elasticities are +0.39, +0.17, and +0.23 respectively.¹⁰

While Rosenthal regards the fourth finding as evidence that the

⁹Rosenthal, op. cit., p. 34.

¹⁰Ibid., p. 95, Appendix Table 1.

price elasticity of demand for beds is "consistent with the economists' expectations"¹¹ (i.e., non-zero), there are two weaknesses in the model which preclude uncritical acceptance of this conclusion. The first deficiency concerns the choice of the "price" variables. Neither singly nor in combination do they correspond to the theoretically appropriate price variable. The latter is given by the marginal point-of-service-price per unit of care, since this is the consumer-patient must pay "out-of-pocket," and hence, is the price most likely to affect his and his physician's behavior. For an uninsured individual, this price is likely to be understated by the daily room charge, since the latter excludes the costs of uninsured nursing care, lab tests, radiological exams, and a variety of ancillary services. For an individual with some degree of insurance coverage, the bias is almost certainly in the opposite direction, though the precise disparity between the point-of-service price and the (higher) daily room charge cannot be ascertained without detailed information on the benefit structure of the insurance policy.¹² Aggregating over individuals and insurance policies, as Rosenthal does, is tantamount to aggregating these biases, and is therefore no solution to the problem. Nor is it solved by including the insurance term. This variable, when combined

¹¹Ibid., p. 40.

¹²Recall that in many lines of insurance, including hospital insurance, the policy may feature, among other things, a deductible, coinsurance, or copayment. With a deductible, the marginal point-of-service price per unit equals the supply price until the deductible is met, and is zero thereafter; with coinsurance, the marginal point-of-service price per unit is a constant proportion of the supply price, where the proportion defines the marginal coinsurance rate; with copayment, the marginal point-of-service price per unit is a constant absolute charge per unit, and is therefore nominally independent of the supply price.

with the average daily room charge is, at best, a crude attempt to define the average net price of care facing consumer-patients in a given state. (In fact, since Rosenthal's insurance variable is simply the proportion of the population with some form of hospital insurance, it yields no information on the comprehensiveness of that insurance, and hence no information on the average coinsurance rate necessary to calculate the average net price.) Quite apart from the fact that such a price is itself theoretically suspect, approximating it in this indirect fashion is inferior to the more direct approach advocated by Paul Feldstein. He emphasizes that:

To estimate a price effect empirically, a "net" price variable must be used, that is, the "out-of pocket" price to the patient and not the stated price. To arrive at this estimate, the effect of health insurance, "free care" and the tax deductibility aspects would first have to be eliminated.¹³

Unfortunately, Rosenthal did not have the data necessary to perform these calculations.

The second deficiency in the model is the omission of a bed supply variable. In this connection, it should first be pointed out that the theoretical case for including this variable in a Rosenthal-type demand model rests on a rather unique assumption about the relationship between bed demand and bed supply. It is not merely that bed supply represents (as it ultimately must) a constraint on bed use, but rather that bed supply is itself a key determinant of bed use. On this assumption, a population's use of hospital facilities is governed (at least in part) by what Martin Feldstein has termed a "pure

¹³Paul J. Feldstein, "Research on the Demand for Health Services," Milbank Memorial Fund Quarterly, Vol. XLIV, No. 3 (July, 1966), p. 146.

availability effect."¹⁴ This, he rightly stresses, is not to be confused with the older and more popular proposition that the supply of beds "creates its own demand." The latter, sometimes referred to as "Say's Law of Hospital Beds," suffers from ambiguity in that it is consistent with both "pure" and "impure" availability effects. The "impure effect" is produced in standard fashion by a rightward shift in the supply curve with no corresponding shift in the demand curve; the "pure effect" is produced in non-standard fashion by a rightward shift in the supply curve which induces a corresponding rightward shift in the demand curve. In the former event "supply creates demand" only indirectly through its depressing effect on the price of hospital care, while in the latter event, "supply creates demand" by directly increasing the quantity of care demanded at every price. Clearly, as M. Feldstein points out, the operation of "a 'pure availability effect' would distinguish the market for hospital services from other markets."^{15,16}

¹⁴Martin S. Feldstein, The Rising Cost Hospital Care (Washington, D.C.: Information Resources Press, 1971), p. 31.

¹⁵Ibid.

¹⁶Speculation on the existence of a "pure availability effect" dates from the work by Roemer and Shain in the late 1950's. They presented evidence of an upward trend in the number of people admitted to hospital for diagnosis and treatment of conditions in the "low medical urgency" category. Thus,

"A half century ago only the most desperately ill were hospitalized; fractures were treated at home or in the doctor's office. Today not only are these cases hospitalized, but so are cases of multiple-tooth extractions, psychoneurosis, epilepsy, diabetes for insulin stabilization, or any obscure condition for diagnosis. All this is made possible by an increase in the relative supply of beds, and reciprocally it creates pressure for continual expansion of the bed capacity." Milton I. Roemer and Max Shain, Hospital Utilization Under Insurance (Chicago: American Hospital Association, 1959), p. 12.

Even if one agrees with Rosenthal that bed supply has no theoretical place in a bed demand function, the bothersome fact remains that inter-state differences in bed use are strongly correlated with inter-state differences in bed availability. On this basis alone, therefore, one can agree with M. Feldstein's charge that "the omission of bed supply from bed use ('demand') equations is a serious misspecification that is likely to have sizeable effects on each of the other coefficients."¹⁷ For present purposes, the coefficients of particular interest are those attached to the average daily room charge and the proportion of the population covered by some form of hospital insurance, since they provide the basis for Rosenthal's inferences about "price" effects.

It can be shown that the specification bias inherent in each of these coefficients will, in general, depend upon: (1) the regression coefficient of the bed supply variable in the "true" (as distinct from estimated) bed demand model; and (2) the regression coefficient of the "price" variable when bed supply is the dependent variable.¹⁸

This argument is based on some very casual empiricism. Not only did the authors fail to persuasively demonstrate that changes in bed supply triggered the increase in non-urgent admissions, but they also made no attempt to net out the effects on use of such arguably relevant factors as rising per capita incomes, changes in medical technology (and resultant changes in consumer attitudes toward care), and growth in hospital insurance coverage. Had these analytical gaps been filled, the study might well have yielded quite different conclusions.

¹⁷M. Feldstein, Economic Analysis for Health Service Efficiency (Amsterdam: North Holland Publishing Co., 1967), p. 193.

¹⁸For a compact and easy to follow treatment of the specification bias which results when the ordinary least squares method is used to estimate the parameters of a model with "left-out" variables, see Rao and R. L. Miller, Applied Econometrics (Belmont: Wadsworth Publishing Co. Inc., 1971), pp. 29-32.

Symbolically,

$$E(\hat{b}_{rc}) = \beta_{rc} + \beta_{beds} \cdot b_{beds.rc}$$

where: $\hat{\beta}_{rc}$ = the O.L.S. estimate of the room charge parameter.

β_{beds} = the regression coefficient of the bed supply variable in the "true" demand model.

$b_{beds.rc}$ = the partial regression coefficient of 'rc' when 'beds' is the dependent variable and all other variables are included in the equation.

and $E(\hat{b}_{ins}) = \beta_{ins} + \beta_{beds} \cdot b_{beds.ins}$

where: $\hat{\beta}_{ins}$ = the O.L.S. estimate of the insurance parameter.

$b_{beds.ins}$ = the partial regression coefficient of 'ins' when 'beds' is the dependent variable and all other variables are included in the equation.

The product term on the right-hand side of each of these expressions defines the bias afflicting the estimated "price" coefficients, b_{rc} and b_{ins} respectively. While Rosenthal does not provide the information necessary to fully evaluate these terms, it is possible to sign them, and thereby gauge the direction (if not the magnitude) of the biases. In the first expression, the bias term likely carries a negative sign, since such information as is provided indicates β_{beds} is positive and $b_{bed.rc}$ is negative.¹⁹ Accordingly, $\hat{b}_{rc}$ is more negative than β_{rc} , and hence, the "price" effect inferred from the room charge coefficient is likely overstated. In the second expression, the bias term likely carries a positive sign, since here there is reason to believe the

¹⁹Gerald D. Rosenthal, The Demand for General Hospital Facilities (Chicago: American Hospital Association, 1964), pp. 60-61.

component terms are both positive.²⁰ Accordingly, b_{ins} is more positive than β_{ins} , and hence the "price" effect inferred from the insurance coefficient is also likely overstated. Each of these statistical distortions weakens Rosenthal's conclusion that the own price elasticity of hospital demand is non-zero.

Notable among those who have continued the effort to detect price effects in the aggregate utilization data is Martin Feldstein.²¹ Like Rosenthal, he asks the question "are there inter-state differences in hospital use that can be ascribed to inter-state differences in the price of hospital care?" His approach to this question is also similar to Rosenthal's in that he makes extensive use of multi-variate regression techniques to distill off the influences of other parameters governing the use of inpatient facilities. However, he does effect certain refinements in theory and measurement which invite close examination.

To set the critical stage, brief mention should be made of the fact that Feldstein's interest in the price-use relationship derives from a wider concern with the determinants of hospital cost inflation. He notes that "most discussions of hospital cost inflation have focussed on how inflation has occurred (e.g., more staff, higher wages, more

²⁰ Ibid. It should be noted however that Rosenthal's regression of beds per 1000 population on per cent of population with insurance yielded an insignificant negative coefficient in 1950 and an insignificant positive coefficient in 1960.

²¹ Martin S. Feldstein, "Hospital Care Inflation: A Study of Nonprofit Price Dynamics," American Economic Review, LXI, No. 5 (December 1971), pp. 853-872.

equipment, etc.), rather than why it has."²² In his view, "rapidly increasing wages and the expansion of hospital employment can be viewed primarily as the results rather than the fundamental causes of this inflation."²³ The real culprit is ultimately identified as "the pressure of the rising demand induced by increases in insurance coverage, personal incomes, the availability of hospital specialists, etc."²⁴ (i.e., a rightwardly mobile market demand curve), but to reach this verdict he must first estimate a model of dynamic price adjustment to excess demand using a sample of observations for each state and year in the decade 1958-67.

Of the twelve equations comprising the model, only the three "demand relations" (two structural equations, one identity) are of immediate interest and concern.²⁵ These are specified in general terms as:

$$(1) \text{ ADM}_t = \text{ADM}(P_t, X_t)$$

$$(2) \text{ MS}_t = \text{MS}(P_t, X_t)$$

$$(3) \text{ BDD}_t = \text{ADM}_t \cdot \text{MS}_t$$

where: ADM = the admission rate (per capita) for
hospital inpatient services

MS = the mean duration of stay per case

BDD = the patient-day rate (per capita)

P = "Price"

²²Ibid., p. 853.

²³Ibid., p. 854.

²⁴Ibid., p. 870.

²⁵The rest of the model includes three "price adjustment" equations, four "components of cost" equations, and two "expansion of capacity" equations.

X = a vector of explanatory variables including income, the availability of hospital facilities and alternative sources of care, the demographic composition of the population, and general attitudes about hospital care. All variables are state and year specific.

In discussing these specifications, Feldstein forewarns the reader that "the precise definitions used for empirical estimations must reflect not only the theoretical relation but also the level of aggregation, (and) the availability of data."²⁶ This statement has important implications for his definition of "price".

For Feldstein, the theoretically appropriate price variable in equations (1) and (2) is "the cost (of care) to the patient net of insurance reimbursements and relative to the price of other consumer goods and services," i.e., the relative net price (RNP).²⁷ In attempting to operationalize this concept, he first assumes that insurance can be treated as a proportional reduction in the gross price of care (P), or, equivalently, the net price (NP) = $INS \cdot P$, where INS is the average (= marginal) coinsurance rate.²⁸ Hence,

$$RNP = \frac{NP}{P} \cdot \frac{P}{CPI} = INS \cdot \frac{P}{CPI} \quad \text{where CPI} = \text{consumer price index}$$

The difficult transition from concept to measurement is initiated when RNP is expressed for state i and year t as:

²⁶Ibid., p. 859.

²⁷Ibid., p. 859.

²⁸This assumption suppresses the fact that an insurance policy cannot be described solely by the coinsurance rate if it also features a deductible and/or other benefit controls.

$$RNP_{it} = \frac{NP_{it}}{P_{it}} \cdot \frac{P_{it}}{CPI_t} = INS_{it} \cdot RP_{it}$$

However, the data do not permit estimation of the state specific coinsurance rate in any year, so Feldstein is forced to assume that INS can be approximated via the relationship:

$$INS_{it} = w_{it} \cdot \mu_t + (1 - w_{it}) \cdot \lambda_t$$

where μ_t = the estimated national average coinsurance rate in year t among the insured population

λ_t = the estimated national average coinsurance rate in 1963 among the uninsured population and w_{it} and $(1-w_{it})$ weights based on year and state specific estimates of the insured and uninsured populations, and their corresponding utilization rates as estimated in a national survey conducted in 1963.

The questionable feature of the INS variable is that its measurement is based on an amalgam of state and national data for various years. Thus μ_t can vary over time but not over states, while λ_t can vary over neither. Couple this with the fact that the ratio of insured to uninsured utilization is assumed invariant over states and time, and it follows that all variation in INS across states is, in any given year, due solely to inter-state differences in the proportion of the population with some form of hospital insurance. This Rosenthal-type insurance variable is some distance from its price-theoretic origins. Even Feldstein admits it is "a very imperfect measure of hospital insurance."²⁹

Feldstein proceeds with the estimation of the (log-linear) demand relations. In his first pass at the data, he runs the admission (ADM) and mean stay (MS) equations without measures of bed supply or

²⁹Ibid., p. 860.

physician supply, and without lagged dependent variables, (i.e., without allowance for dynamic adjustments), and, in each case, he finds that RNP carries a significant negative coefficient. It indicates that the net price elasticities of ADM and MS are -0.63 and -0.49 respectively; together, and by virtue of their additivity, these imply a net price elasticity of -1.12 for patient days.³⁰ But, after augmenting the equations with a bed supply variable in order to correct for the initial specification bias (and, as well, to test for what turns out to be a significant "pure availability effect"), the relative net price (RNP) elasticity for patient-days falls to -0.67 [i.e., $\eta_{ADM} + \eta_{MS} = -0.44(+)$ -0.15].³¹ Explicit decomposition of the net price (RNP) elasticity into its component gross price (NP) and average coinsurance (INS) elasticities (for patient days) yields estimates of -0.71 and -0.17 respectively. The disparity between these two estimates makes Feldstein a bit uneasy, since, unlike the corresponding results from the basic equations (no bed supply variable), it contradicts his assumption that insurance is a proportional reduction in the gross price.³² He speculates that:

It may be a statistical artifact, reflecting the approximation in defining INS....It may however reflect a greater concern in admission decisions with the gross price than with the net

³⁰Ibid., p. 863. In a re-run of the basic admissions and mean stay equations with RP and INS substituted for RNP, he finds that the gross price (RP) elasticity and the coinsurance (INS) elasticity are roughly similar in magnitude to the original net price (RNP) elasticity. This he takes as evidence supporting his basic assumption that insurance is a proportional reduction in gross price, but this would seem unwarranted given that INS is a very poor proxy for the coinsurance rate.

³¹Ibid., p. 865, Table 2, equations 24 and 26.

³²See comment Supra, fn.30.

price, perhaps...because physicians influence admissions more than mean stay and are more concerned with gross resource use than with the net cost to patients. Only further studies of a more disaggregate kind will clarify this.³³

Further and final reductions occur in the estimated elasticities when allowance is made for dynamic effects by including a lagged dependent variable in the equations. In particular, the net price elasticity of patient days is estimated to be -0.27 in the short-run, and -0.55 in the long-run.³⁴ With these estimates in hand, Feldstein concludes "there is a substantial price elasticity of demand for hospital bed days".³⁵

Interestingly, Feldstein has since offered more extended discussion of these results.³⁶ First, he notes that his estimated net price elasticities are biased downward (i.e., less negative than their "true" values) if, as seems reasonable to assume, gross price and quality of care are positively correlated.³⁷ Second, he interprets the marked disparity between the estimated short-run (-0.27) and long-run (-0.55) price elasticities as a "delayed demand response ... evidence of

³³Ibid., p.866. More will be said of this in subsequent discussions.

³⁴Ibid., p. 867.

³⁵Ibid., p.866.

³⁶M. Feldstein, "Econometric Studies of Health Economics", in M.D. Intriligator and D.A. Kendrick (eds.), Frontiers of Quantitative Economics, Vol.II, (Amsterdam: North-Holland Publishing Co., 1974).

³⁷Ibid., p. 387.

the physician's role in modulating household behavior."³⁸ Third, he regards the disparity between the gross price elasticity (-0.71) and the "average coinsurance" elasticity (-0.17) as support for his generalized agency theory of household demand for hospital services."³⁹ Each of these points merits brief comment.

In respect of the bias afflicting the estimated net price elasticities, Newhouse and Phelps have shown that the (previously noted) assumptions and approximations used to construct the state coinsurance rate (INS) "appear to lead to bias in the coefficients away from zero."⁴⁰ i.e., in a direction opposite to that implied by the assumed positive correlation between price and quality. While the net effect of these opposing biases is unknown, it is clear that the observed elasticities are contaminated to a degree not recognized by Feldstein and others who have argued the case for "incomplete" national health.

³⁸Ibid.

³⁹This theory is essentially an attempt to integrate the behavior of the consumer-patient and the provider-physician using the concept of the "agency relationship." A "complete" agency relation exists whenever the physician acts solely to maximize the welfare of the patient, while an "incomplete" agency relation exists whenever the physician acts to maximize some (unspecified) combination of his own and his patient's welfare. A key prediction of the agency model is that the differential between the elasticity of use with respect to gross price (resource cost) and the elasticity of use with respect to insurance coverage (i.e., the coinsurance rate) depends upon the completeness of the agency relation. In particular, "the more complete the insurance coverage, the more the physician must depart from his role as the patient's agent and assume the responsibility for the allocation of a public resource; when the insurance coverage is nearly complete, the elasticity with respect to insurance would therefore be quite low." Ibid., p. 384.

⁴⁰Joseph P. Newhouse and Charles E. Phelps, "On Having Your Cake and Eating It Too: Econometric Problems in Estimating the Demand For Health Services." (Santa Monica: Rand Corporation, R-1149-NC, April 1974), p. 7, and Appendix B, pp. 38-39.

insurance on the basis that there is a "substantial" price elasticity of demand for care. Second, the interpretation of the short-run and long-run elasticities is as much rationalization as explanation. The rival hypothesis that elasticities are greater in the short-run due to the "shock-effect" of price changes on clinical and household decision-making has equal, if not greater, logical appeal.⁴¹ In any case, neither can be rigorously tested with state aggregate data. And finally, while the nascent "generalized agency" theory is a potentially fruitful one, the assumption that physicians are cost-conscious in their use of expensive hospital facilities is counter-intuitive? It may "explain" the disparity between the estimated gross price and coinsurance elasticities, but it is contrary to the working assumption of most health economists that physicians are able and willing to misallocate hospital resources whose costs they do not directly bear. In short, all the relevant incentives bearing on the physician are biased toward private cost-minimization and public cost-maximization. There can be no quarrel however with Feldstein's concluding observation that all his findings "must be interpreted cautiously."⁴²

An identical stricture applies in respect of the price effects presented by Davis and Russell.⁴³ They too use the state aggregate data

⁴¹It also finds some support in the Canadian data. See R. G. Beck, "Some Dynamic Effects of Copayment on Utilization of Medical Services in Saskatchewan," Health Economics Symposium (Queen's University, 1976), pp. 98-110.

⁴²M. Feldstein, Frontiers of Quantitative Economics, Vol. II, p. 388.

⁴³Karen Davis and Louise B. Russell, "The Substitution of Hospital Outpatient Care For Inpatient Care," The Review of Economics and Statistics, LIV, No. 2, (May, 1972), pp. 109-120.

as their inferential base. Their model, primarily developed "in order to ascertain the effect of a change in the price of hospital outpatient care on the demand for inpatient care,"⁴⁴ features two equations identical in all theoretical and practical aspects save for the definition of the dependent variable: it is defined in one as admissions per capita, and in the other as mean stay per admission. Beyond this it is sufficient to note that in its estimable form the model excludes a bed supply variable, and includes, *inter alia*, two measures of insurance coverage, and three alternative measures of the gross price of inpatient care. The insurance variables are defined as the proportions of the state population enrolled in Blue Cross and non-Blue Cross plans respectively. Underlying this distinction is the reasonable assumption that Blue Cross plans offer more comprehensive coverage of outpatient care, and hence are the plans least likely to promote substitution of inpatient care for outpatient care via a "relative price effect", i.e., by reducing the net price of inpatient care relative to the net price of outpatient care. For this reason, the value of the Blue Cross coefficient is expected to be smaller than the value of the non-Blue Cross coefficient. As for gross price, it is alternatively measured by inpatient revenue per patient day, inpatient revenue per admission, and the basic charge for a two-bed room.

Using the ordinary least squares estimation method, the log-linear equations are fitted to cross-state data specific to non-governmental, non-profit hospitals in 1969. This yields gross price elasticities for admissions ranging from -0.19 to -0.46, and gross price

⁴⁴*Ibid.*, p. 113.

elasticities for mean stay ranging from -0.30 to +0.35. In magnitude and statistical significance, these estimates are noticeably sensitive to both the definition of gross price and the specification of the full equations. To illustrate the first type of sensitivity, the elasticity of admissions is -0.19 (insignificant at the .05 level) when the gross price is measured by inpatient revenue per admission, and -.38 (significant at the .01 level) when it is measured by the room charge, ceteris paribus. To illustrate the second type of sensitivity, when state per capita income is deleted from the admission equation, the coefficient of inpatient revenue per admission changes from -0.19 (significant as above) to -0.32 (significant at the .01 level), ceteris paribus.

Similar sensitivities are evident in the mean stay equation. Thus, the elasticity of mean stay is -0.24 (insignificant at the .05 level) when gross price is measured by inpatient revenue per patient day, and +0.35 (also insignificant at the 0.5 level) when it is measured by the room charge, ceteris paribus. And, even more important, when both insurance variables (and all other non-price variables except income and occupancy rate) are deleted from the equation, the coefficient of inpatient revenue per patient day changes from -0.24 (insignificant as above) to -0.30 (significant at the 0.1 level). As between these latter two estimates, Davis and Russell choose -0.30 (it implies that a 10 per cent increase in revenue per patient day results in a 3 per cent decrease in average length of stay), and conclude "stays

are strongly influenced by inpatient prices per day."⁴⁵ However, the decision to delete variables should not be taken lightly, especially when it leads, as it does in this case, to substantially different conclusions.⁴⁶ By deleting the insurance (and other) variables, and thereby suppressing their underlying collinearity with the gross price variable, Davis and Russell improve the precision (reduce the variance) of the gross price coefficient. This explains why the coefficient turns significant, but it does not explain why variables are deleted which have a "rightful" place in the specification. Moreover, the non-Blue Cross variable is deleted notwithstanding that it carried a significant (but perversely negative) coefficient. Without the insurance variables, the demand equations are misspecified, and, as such, afford no opportunity to test whether or not there are net price effects governing the demand for inpatient care.

Attention now shifts to four studies where the search for price effects has been concentrated in various types of disaggregate data using micro-analytic research designs. Each features the individual "household" (family and/or family member) as the basic unit of observation, and each asks some variant of the question "are there inter-household variations in hospital use that can be ascribed to inter-household differences in the net (point-of-service) price of inpatient care?" Moreover, each attempts to answer the question using multi-variate regression techniques in order to control for inter-

⁴⁵Ibid., p. 116.

⁴⁶For a clear statement of this position and the associated econometric theory see Rao and Miller, op.cit., pp. 35-38, and pp. 56-59.

household differences in socio-demographic and other arguably relevant factors and characteristics governing the use of inpatient services. This statistical methodology is, of course, similar to that employed by the studies just reviewed, the difference being that the variables are (for the most part) household specific rather than population specific. While this strategy avoids much of the aggregation bias which afflicted attempts to isolate state level price effects, it is by no means a problem free one. Thus, in addition to the endemic difficulty of operationalizing the net price concept, it must contend with the strong possibility that, in a voluntary insurance system, households may purchase policies providing coverage consistent with their expectations of illness and attendant demands for care. Known in the insurance literature as "advance selection," and in the econometric literature as "simultaneity bias," this imparts a definite, albeit often indeterminate, inconsistency away from zero in price elasticities estimated from "micro" data.⁴⁷ These and other problems will be more fully explored in the context of the studies slated for review and comment.

The first study justifiably rates its "pioneering" status, not only because it was published in the same year (1964) as Rosenthal's, but also because its authors, Paul Feldstein and Ruth Severson, seized the initial opportunity to quantify the price-use relationship using

⁴⁷Put another way, any negative correlation between net price and hospital use will be overstated to the extent "high use" ("low use") households are those who have correctly anticipated their illness experiences and chosen policies offering "low" ("high") point-of-service prices. This problem is not of course unique to the household level data, but its severity is likely greater there than in the state level data.

family-level data.⁴⁸ These data were obtained from a nation-wide survey of the U.S. civilian, non-institutionalized population in 1958, conducted jointly by the Health Information Foundation (HIF) and the National Opinion Research Center (NORC) of the University of Chicago. In addition to information on the extent of their health insurance and how this insurance helped finance their hospital, medical, dental, and drug expenditures, respondents provided details on age, income, education, marital status, family size, location, and free care situations pertinent to appraising their family's demographic, social, and economic status. Taken together, these data facilitated the specification and estimation of multi-variate demand relationships, and the subsequent testing of numerous hypotheses concerning "inter-family"⁴⁹ variations in health services use and expenditure.

At the risk of doing some injustice to the scope of the study, only those results specific to families' hospitalization experiences will be examined here. These are presented in Table 2.1, where, in particular, ordinary least squares estimates of the log-linear equations relating admissions and inpatient days to a common vector of explanatory variables are shown. In each equation, the two variables whose measurement and performance command exclusive attention are "Price" and "Percent Insurance." Due to data limitations, Feldstein and Severson

⁴⁸Paul J. Feldstein and Ruth Severson, "The Demand for Medical Care," Report of the Commission on Cost of Medical Care, (Chicago: AMA, 1964).

⁴⁹Inter-family is placed in inverted commas, since Feldstein and Severson first formed (in unspecified fashion) sixty-six primary sampling units, and then averaged the individual family data within each of these groups. Hence, it is strictly inter-group, not inter-family variation in use and expenditure which they set out to explain.

Table 2.1: Feldstein and Severson's Log-Linear Regression
Equations Based on 1958 HIF-NORC Survey Data

Dependent Variables: Admissions per Family (EQN.Va); and
Patient-Days per Family (EQN.VIa)

Independent Variables	<u>Regression Coefficients*</u>	
	EQN.Va	EQN.VIa
Constant	+2.70	-8.20
Family Income	-0.76 (.36)	+0.47 (.43)
Age of Head	+0.16 (1.44)	+3.49 (1.76)
% 65 Years Plus	-0.13 (.22)	-0.04 (.26)
% Under 5 Years	+0.45 (.21)	+0.46 (.24)
% Single Person Families	-0.02 (.13)	+0.02 (.15)
Education	-0.01 (.15)	-0.24 (.17)
Free Care	+0.04 (.09)	-0.10 (.11)
% Urban	+0.00 (.06)	+0.06 (.07)
Price	+0.11 (.10)	+0.02 (.10)
% Insurance	+0.45 (.11)	+0.25 (.13)
R ² (unadjusted)	0.42	0.32
n	66	66

*Standard Errors in Parentheses

Source: Paul Feldstein and Ruth Severson,
The Demand For Medical Care,
Equations Va and VIa, pp. 67-68.

were unable to measure net price directly, so they generated the price variable by dividing net family expenditures on inpatient care by the quantity of care, i.e., via the relationship

$$\frac{NEX}{Q} = \frac{P_n \cdot Q}{Q} = P_n$$

where NEX is net expenditure (after insurance reimbursements, etc.), Q is quantity, and P_n is the average net price per unit of care. As for the insurance variable, (termed a "pseudo-price" by the authors), it is defined and measured as the "proportion of hospital bills covered by insurance,"⁵⁰ i.e., as,

$$INS = 1 - \frac{NEX}{GEX} = 1 - \frac{P_n Q}{P_g Q} = 1 - \frac{P_n}{P_g} = 1 - k$$

where GEX is gross expenditure (before insurance reimbursements, etc.), P_g is the average gross price per unit of care, k is the average coinsurance rate, and NEX, P_n , and Q are as above.

While these are acceptable definitions in principle, in practice of course, they are subject to the weaknesses of household interview data, especially measurement error. Moreover, given that the data are averaged over the families' observations within each of the sixty six primary groups,⁵¹ the assumption that the "average unit price is the same for all families"⁵² in a given group is questionable; it should not have been adopted without evidence that the groups were defined so as to minimize inter-hospital variation in gross price and inter-policy

⁵⁰Ibid., p. 60.

⁵¹This averaging procedure perhaps explains why the equations fit the data as well as they do; the R^2 values of .32 and .42 are atypically high for a cross-section of household observations.

⁵²Ibid., p. 60.

variation in bill coverage (i.e., the coinsurance rate). The key question, however, is why the equations were specified and estimated with both variables included. Theoretically, there is no place for the insurance term, since the price term is by itself an estimate of the net price of care. As Paul Feldstein has himself argued, it is only when the data do not permit the net price to be constructed in this fashion that one must resort to more indirect approximations using a combination of gross price and insurance coverage a la Rosenthal.

This perhaps explains why the "price" variable performs so perversely. It indicates, contrary to expectation, that a 10 per cent increase in the net price of inpatient care leads to a 1.1 per cent increase in admissions, and a .2 per cent increase in patient-days. These positive (but statistically insignificant) price effects might well be due to multi-collinearity between the price and insurance terms, and, in particular, to a strong negative correlation between the net price and the per cent of the bill covered by insurance. Unfortunately, this reasonable assumption cannot be tested with the data provided, but there is every reason to believe that had the insurance term been omitted from the equations, the resultant price effects would have differed substantially. However, there would still be problems in observing the true elasticities, ~~given~~ the strong likelihood of measurement errors in the net expenditure and quantity data used to generate the net price variable.⁵³

⁵³Newhouse and Phelps have shown that these errors will, in general, have a negative covariance with the true net price, and, in consequence, will produce (contrary to the "standard" errors in variables case) an estimated price effect with an inconsistency away from zero and/or a sign opposite to the true price effect. Hence, some combination of measurement error and multi-collinearity between the

Interestingly, the insurance variable is the most significant of any included in either equation. It indicates that a 10 per cent increase in the proportion of the bill paid by insurance leads to a 4.5 per cent increase in admissions and a 2.5 per cent increase in patient-days. Feldstein and Severson proffer that these estimates "are likely to be biased downward because of errors in the measurement of insurance."⁵⁴ This may be so, but there are also distortions produced by the inclusion of what are, in effect, two "net price" variables.

Finally, it should be noted that Feldstein and Severson are appropriately circumspect in their overall conclusions. They evidently believe that the price elasticity of hospital use is very low, but caution the reader that their study has limitations which must "serve as a guide to those who might give undue weight to certain findings."⁵⁵ They see "the main area where further research would greatly increase the accuracy of the results is in the data to be collected."⁵⁶

Next on the agenda are two micro-analytical studies very similar to one another in several key respects. Each has as its principal objective the determination of the price elasticity for only one component of hospital use, namely, the duration of stay per admission, and each attempts to do this using a disaggregation scheme featuring groups of patients heterogeneous with respect to insurance coverage, but

price and insurance terms may well account for the perversely positive price coefficient in both the admission and patient-day equations. See Joseph P. Newhouse and Charles E. Phelps, op.cit., pp. 2-13, and Appendix D, pp. 45-47.

⁵⁴P. Feldstein and R. Severson, op.cit., p. 67.

⁵⁵Ibid., p. 76.

⁵⁶Ibid.

homogeneous with respect to such basic characteristics as age, sex, and medical diagnosis. This strategy enables the investigators to test the specific hypothesis that the price elasticity of hospital stay is not a constant, but a variable whose magnitude differs across diagnostic and/or illness categories.⁵⁷

In his study of the price-stay relationship, Gerald Rosenthal uses data drawn from a sample of 68 non-federal, short-term general and special hospitals in the five-state New England area. Representing 15,685 admissions, or approximately one per cent of all admissions to these hospitals in 1962, these data are grouped into 28 categories defined by the patient's age, sex, and diagnosis. Variations in length of stay within each of these "homogeneous" groups are hypothesized to be "a function of some price-payment measure."⁵⁸ The equations on which the analysis is based are:

⁵⁷Underlying this hypothesis is the reasonable assumption that the discharge decision is more discretionary for some "case-types" than for others. Accordingly, comparatively high (low) elasticities are "expected" for those cases where the discharge decision is governed less (more) by strict medico-technical requirements than by patient preferences and/or the availability of alternative treatment and recovery facilities. "Expected" is placed in inverted commas, since neither investigator attempts an a priori classification of cases in terms of discretionary vs non-discretionary discharge decisions. This necessarily weakens the logic of the exercise (though not the test of the hypothesis that the price elasticity varies), insofar as it denies them an independent standard of discretion against which the observed pattern of elasticities might be compared. Stated differently, they are at risk of committing a tautological error by defining discretionary (non-discretionary) cases as those with comparatively high (low) price elasticities. The problem is largely unavoidable in the absence of a well-defined medical model which clearly specifies the indications for discharge for all classes of patient and for all illnesses and conditions. Such a model does not exist.

⁵⁸Gerald D. Rosenthal, "Price Elasticity of Demand for Short-Term General Hospital Services," in Herbert E. Klarman (ed.) Empirical Studies in Health Economics (Baltimore: The John Hopkins Press, 1970) p. 105.

$$(1) Y = aX_1^{b_1}$$

$$(2) Y = aX_2^{b_2}$$

where Y is length of stay, X_1 is the patient's cash outlay as a per cent of the total bill, X_2 is the average daily room charge, and b_1 and b_2 are the elasticities of length of stay with respect to each of the two alternative price-payment measures. Notice that the first measure is equivalent to the average coinsurance rate, and that the second measure is equivalent to the average gross price per day averaged over the entire stay. Notice also that neither measure corresponds to the price actually paid by the patient, and hence neither corresponds to the theoretically appropriate net price variable. More will be said of this very shortly.

Estimating the log-linear equations using the ordinary least-squares regression method, Rosenthal finds that the two price-payment measures perform rather differently. In particular, the coefficient of the average coinsurance rate is -0.06 or less (negative) and statistically insignificant in the majority (24/28) of cases, while the coefficient of the average room charge is -0.20 or more (negative) and significant in the majority (15/28) of cases.⁵⁹ Curiously, for this reason, he judges the average daily room charge to be the most

⁵⁹Ibid., Tables 3, 4, and 5, pp. 110-113. Across the 20 medical and 8 surgical categories b_1 ranges in value from -0.065 (for males, age 0-4, with acute upper respiratory infection) to +0.079 (i.e., perversely positive for males, over age 64 with arteriosclerotic and degenerative heart disease). Correspondingly, b_2 ranges in value from -0.973 (for post-operative length of stay for females, age 0-4, undergoing tonsillectomy and adenoidectomy) to +0.440 (i.e., perversely positive for females, over age 64, with cholelithiasis and cholecystitis).

"persuasive" price, and concludes there is "evidence of significant price elasticity of demand."⁶⁰

The obvious deficiency is, of course, the choice of the price variables. As mentioned above, neither qualifies as the appropriate net price, and hence their coefficients provide no basis for inference regarding the true price elasticity. As others have pointed out,⁶¹ this problem could have been avoided had Rosenthal estimated the equation $Y=a(X_1X_2)^b$, since X_1X_2 provides an estimate of the true net price. However, he pre-empted his critics to some extent by aptly describing his results as "not determinate".⁶²

⁶⁰Ibid., p. 111.

⁶¹Paul B. Ginsburg and Larry M. Manheim, "Insurance, Copayment, and Health Services Utilization: A Critical Review," Journal of Economics and Business, Vol. 25, No. 3, Spring-Summer 1973, p. 150.

⁶²Rosenthal, op.cit., p. 117. In addition to the specification error, the study has other weaknesses. Thus, the estimated elasticities are likely to be biased upward due to advance selection. More important, the elasticity pattern across age categories is disconcerting. For example, within certain medical categories (e.g., arteriosclerotic and degenerative heart disease), the room charge elasticity is higher for the group aged 45-64 than for the group over 64. To Rosenthal, this suggests that the aged are subject to more serious illness episodes, and are therefore less able to adjust their lengths of stay in response to price changes. Perhaps, but at the same time he offers no rationalization for the fact that within certain surgical categories (e.g., curettage of the uterus), the elasticity is higher for older adults than for younger. Equally important, he makes little attempt to appraise his hypothesis that "certain diagnostic categories will be more likely to show a higher price elasticity of length of stay than others." Ibid., p. 109. The reader is left the difficult task of determining whether there is any systematicity in the overall pattern.

Hyman Joseph's results are equally indeterminate.⁶³ Using patient-discharge data from twenty-two Iowa hospitals for the period November 1965 to April 1966, he attempts to establish the "price-stay" relationship in each of twenty-two diagnostic categories by regressing length of stay on a dummy variable describing the patient's insurance status, and other control variables.⁶⁴ He finds that the insurance variable (equal to 1 if the patient received some third-party payment from government, Blue Cross, or private hospital insurance, and 0 otherwise) carries a positive and statistically significant coefficient in seven of the twenty-two categories. At this point it is reasonably concluded that "third party payments tend to induce patients to stay longer in hospital for some illnesses and conditions and have no statistically discernible effect for other illnesses and conditions."⁶⁵ The analysis subsequently falters, however, when "price elasticities of demand" are calculated by dividing the insurance coefficient by the constant term in each equation. This procedure yields estimates of the difference in mean stay between patients with and patients without some

⁶³Hyman Joseph, "Hospital Insurance and Moral Hazard," The Journal of Human Resources, VII, 2, Spring 1972, pp. 152-161.

⁶⁴The other variables, all of which are also of the dummy or categorical type, describe the patient's age, sex, accommodation type (standard, semi-private, or private ward), separation status (dead or alive), and case complications (present or absent). Diagnosis is, of course, controlled for in the initial grouping of the data.

⁶⁵Ibid., p. 156. Interestingly, the insurance effect is seen by Joseph to be present (absent) among the least (most) "serious" and/or the least (most) "psychological unpleasant" conditions, as his a priori reasoning would suggest. However, included among the seven diagnoses where the effect is present are cholelithiasis (gall-stones) and duodenal ulcer, while included among the fifteen where the effect is absent are hemorrhoids and hernia of the abdominal cavity without mention of obstruction. Few would judge the former to be less serious than the latter.

form of insurance, but it does not yield estimates of the price elasticity because the insurance variable contains no information on the net prices faced by the two groups of patients. As Newhouse and Phelps put it, "there is no measure of price (coinsurance) change in Joseph's data," and hence "his estimates cannot be considered as elasticities."⁶⁶ Clearly, therefore, nothing in the data warrants his claim that the elasticities are "generally low."⁶⁷

Much more compelling evidence in support of such a claim is to be found in Newhouse and Phelps' own study of "price effects."⁶⁸ They find that the relevant price elasticities of hospital admissions and length of stay are -0.17 and -0.06 respectively, or very near zero, the

⁶⁶Charles E. Phelps and Joseph P. Newhouse, Coinsurance and the Demand for Medical Services, (Santa Monica: Rand Corporation, R-964-1-OEO/NC, October 1974), p. 28.

⁶⁷Joseph, op.cit., p. 160. Subject to this rather binding constraint, it may be noted that he calculates an elasticity greater than one for two categories, an elasticity less than one for thirteen categories, and a perversely positive elasticity for seven categories. Of the fifteen negative elasticities, only seven are statistically significant, but neither their magnitude nor diagnostic identity are revealed.

⁶⁸Joseph P. Newhouse and Charles E. Phelps, "New Estimates of Price and Income Elasticities of Medical Care Services," in Richard N. Rosett (ed.), The Role of Health Insurance in the Health Services Sector (New York: National Bureau of Economic Research, 1976), pp. 261-313. This is the third and most comprehensive of several closely related studies published by these investigators. The other studies are: "Price and Income Elasticities for Medical Care Services," in Mark Perlman (ed.), The Economics of Health and Medical Care (London: Macmillan, 1974), pp. 139-161; and Price and Income Elasticities for Medical Care Services (Santa Monica: Rand Corporation, R-1197-NC/OEO, June 1974), 32 p.

point of perfect inelasticity.⁶⁹ These estimates are notable not only because they are the smallest of any in the published health economics literature, but also because they are free of the major defects and distortions which marred those reviewed thusfar.

The study is unique in several respects, the most fundamental of which⁷⁰ is its application and generalization of the conceptual framework developed by Michael Grossman.⁷⁰ Grossman's capital theoretic approach to the demand for health is unquestionably one of the most notable additions to the health economics literature in recent years, and, because of its prominent place in Newhouse and Phelps' research, the opportunity will be taken at this juncture to review its essentials.

At the outset of his study, Grossman makes clear his disaffection with traditional demand theory and its basic assumption that the consumer's well-being depends directly upon goods and services purchased in the market. Acceptance of and reliance upon this paradigm has led to research emphasizing the demand for medical services, notwithstanding that "most students of medical economics have long realized that what consumer's demand when they purchase medical services are not these services per se but rather good health."⁷¹ For this reason, "it seems logical to study the demand for medical care by first constructing a model of the demand for health itself," and, accordingly, this becomes Grossman's primary aim.

⁶⁹Ibid., p. 283.

⁷⁰Michael Grossman, The Demand for Health: A Theoretical and Empirical Investigation (New York: National Bureau of Economic Research, Occasional Paper 119, 1972), 115 p.

⁷¹Ibid., p. xiii.

The basic distinction between health⁷² and medical services is formalized with reference to the new approach to consumer behavior developed by Becker and Lancaster.⁷² Health is viewed as one of those fundamental objects of choice or "commodities" which enters the consumer's utility function, while medical services, including hospital care, are "market goods" which, together with the consumer's own time, are viewed as inputs in the health production function. Thus, the individual is said to have a fundamental demand for good health and a derived demand for medical services.

The capital-theoretic flavour of Grossman's model stems from the proposition that the commodity, good health, is a durable item. Individuals are assumed to "inherit an initial stock of health that depreciates over time -- at an increasing rate, at least after some stage in the life cycle -- and can be increased by investments."⁷³ Gross investments in the health stock take place via a production process in which variable combinations of "own time, medical care, diet, exercise, housing, and other market goods" are direct inputs.⁷⁴ Thus, the health stock is not exogenous, but depends at least in part, on the resources

⁷²See Gary S. Becker, "A Theory of the Allocation of Time," Economic Journal Vol. 75, No. 299, September 1965; and Kelvin J. Lancaster, "A New Approach to Consumer Theory," Journal of Political Economy, Vol. 75, No. 2, April 1966.

⁷³Michael Grossman, op.cit., p. xv.

⁷⁴Ibid.

allocated to its production by the individual.⁷⁵

For the welfare-maximizing individual, the basic choice-theoretic problem is to determine the optimum quantity of health at any given age during his lifetime. As a rational and fully informed consumer-producer of health, he does this by comparing the costs and benefits associated with incremental units of health capital. Owing to the non-marketability of health -- it cannot be sold in the capital market -- the relevant cost is the "shadow price" per unit of health capital. This price in turn depends on the price of medical care, the value of time, and the efficiency of the health production process, i.e., the marginal cost of gross investment. Multiplying the "shadow price" by the own rate of interest and the age-specific rate of depreciation, yields the rental price or user cost per unit of health capital on which the age-specific gross investment decision is based. The benefits relevant to this decision are those which accrue to the individual from the ownership of health capital. In the most general case, they include both the consumption and investment benefits of the "healthy time" outputted by the health stock. This time, the self-evident complement of "sick time," is a direct source of utility (the consumption benefit), and a means of increasing the total time available for market and non-market activities (the investment benefit). When

⁷⁵This view -- that the individual has some control over his health status -- is clearly complementary to the discussion of volitional "life-style" differences and their positive and negative effects on personal health contained in a recent, and very readable, analysis of health care problems and policies in Canada. See Marc Lalonde, A New Perspective on the Health of Canadians (Ottawa: Information Canada, 1975), 76 p.

expressed in monetary terms,⁷⁶ these benefits define the marginal value of the output produced by age-specific increments in the health stock, the requisite concept for benefit enumeration purposes. Armed with this information on costs and benefits, the individual is in a position to determine his optimum health stock, and, by extension, the optimum quantity of gross investment in any given period.⁷⁷ Equilibrium is achieved when the total rate of return to an investment in health equals the user cost of health capital in terms of the price of gross investment.

Beyond this, and for purposes of reviewing Newhouse and Phelps' research, it is sufficient to note that Grossman penultimately specifies a demand for medical care function in which the individual's age, education, and wage rate are the principal arguments. Age is taken to be a proxy for the rate of depreciation on health capital. It is expected to carry a positive coefficient on the twin assumptions that depreciation increases with age, and that the individual always offsets

⁷⁶In particular, the marginal utility of healthy days is "monetized" in terms of the marginal utility of wealth, while the marginal product of the stock of health in the production of healthy days is "monetized" in terms of the wage rate. See Grossman, op.cit., p. 4.

⁷⁷Formally, this requires that the individual maximize his inter-temporal utility function (in which health and all other commodities are arguments) subject to the constraints given by: (1) the definition of net investment, viz., gross investment minus depreciation; (2) two sets of production functions, one for investment in the production of health, the definition of "full wealth," viz., initial assets plus the present value of the earnings an individual would obtain if he spent all his time at work. For the complete mathematical derivation, see Ibid., Appendix A, pp. 84-86.

at least some of the implied health stock loss with gross investment.⁷⁸ Education, an "environmental" variable, improves the individual's efficiency as a producer of health by increasing the marginal products of the direct inputs in the gross investment production function. This reduces the quantity of medical care and own-time required to produce a given amount of health capital and, with no change in the prices of these inputs, reduces the marginal cost of gross investment. While the resultant increase in the rate of return to an investment in health induces the more educated individual to demand a larger health stock, Grossman shows that the efficiency gains from education may lead the individual to demand a lower quantity of medical care.⁷⁹ Hence, education is expected to carry a negative coefficient in the specified medical care demand function. For its part, the wage rate measures the price of time, and, specifically, the hourly return to time spent in market activity, and the opportunity cost of time spent in non-market activity. It is hypothesized that an increase in the wage rate increases the demand for medical care since: (1) it increases the value of "healthy time" outputted by each and every unit of health capital, and, ~~inso facto~~, increases the stock demand for health and the derived demand for medical care; and (2) it increases the price of time relative

⁷⁸In Grossman's model, the individual "chooses" not to indefinitely maintain his health stock due to the inexorable increase in its supply price occasioned by the rising rate of age-related depreciation. Indeed, as one reviewer of the study wryly observed, "at a certain stage, death occurs because it is too expensive to go on living!" See Robert Ferber, "Consumer Economics, A Survey," Journal of Economic Literature, Vol. XI, No. 4, December 1973, p. 1325.

⁷⁹The precise requirement for this to occur is that the elasticity of the marginal efficiency of health capital schedule (i.e., the stock demand curve for health) be less than one. For a rigorous proof, see Grossman, op.cit., p. 27.

to the price of medical care, and, ipso facto, induces the substitution of medical care for non-market time in the production of health capital, i.e., induces the least-cost producer of health to "pay the doctor, rather than play the doctor."⁸⁰

In its final estimable form, this specification is augmented to include family income, family size, and the sex of the adult member whose demand for medical care constitutes the dependent variable. Family income, a proxy for family wealth, assumes its conventional role as a measure of the family's pooled ability to command real resources, and is therefore expected to exert a positive effect on the demand for medical care. Family size is included for two reasons, one that the number of children and the health stocks of adult members might be complements in the utility function, and the other that the income effect might be a function of family size, i.e., that per capita and not family income might be the appropriate measure of the adult individual's ability to pay for medical care. The adult member's sex is included "to test the hypothesis that females might be more efficient producers of health than males (or vice versa)," and hence less reliant upon medical care.

After adopting this as their basic specification, Newhouse and Phelps effect two major modifications designed to clarify the empirical nature of the price elasticity for medical services. First, they relax Grossman's assumption (hitherto unstated) that these services constitute a homogeneous aggregate whose components have fixed relative prices and

⁸⁰ Obviously, if medical care and own time are assumed not to be substitutable factors of health production, the positive effect of a wage rate increase on the demand for medical care is due solely to its positive effect on the demand for health.

are consumed in fixed proportions. Thus, they distinguish between hospital care and ambulatory physician care, and estimate separate demand equations for each. It should be noted that this disaggregation scheme does not represent a total relaxation of Grossman's homogeneity assumption. It rests itself on the strong assumption that hospital care and ambulatory physician care are legitimate composite commodities. This not only suppresses the manifest heterogeneity of diagnostic and treatment services provided within each setting, but it also precludes assessment of price effects at a level where policy relevant distinctions might well be drawn, for example, between in-hospital medical and surgical services, or between referred and unreferred physician office visits. Hence, if there is a major weakness in the Newhouse and Phelps study, it is that disaggregation is not carried far enough. Second, they relax Grossman's assumption (also hitherto unstated) that the prices of these services are invariant across individuals. On the basis of this assumption Grossman excludes the price of medical care from the medical care demand function, but, as Newhouse and Phelps correctly argue, individuals do face different net prices of care depending on their choice of insurance policy. Moreover, to the extent their choice is made on the basis of expected medical care consumption, the net price elasticity estimated via ordinary least squares will be biased upward (i.e., more negative than its true value). This, the problem of advance selection, Newhouse and Phelps anticipate by treating net price as an endogenous variable, and, in their econometric analysis, by experimenting with a more appropriate (two-stage least squares) estimation procedure.

Other notable features of Newhouse and Phelps' model are best

illustrated with reference to the subset of equations designed to explain inter-individual variation in the use of hospital services.

These can be specified in general terms as:

$$(1) \quad P(A)_i = f(P_i, Z_i)$$

$$(2) \quad LS_i = g(P_i, Z_i)$$

where, for the i^{th} individual:

$P(A)$ = the probability of admission

LS = the length of stay conditional on admission

P = the "price" of hospital care

Z = a vector of explanatory variables including inter alia, the individual's wage rate, education, age, sex, family size, and family income.

Equation (1) relates exclusively to the admission decision. This decision is of course an "all of nothing" one in the sense that the individual is either admitted or not admitted to hospital during any arbitrary time period. Correspondingly, the observed probability of admission is necessarily limited to two values, 1 if admission occurs, and 0 if it does not. To accommodate this fact when estimating equation (1), Newhouse and Phelps employ a dichotomous or binary dependent variable which takes on the value 1 for a "user" and 0 for a "non-user." As for the price variable, they note that the theoretically appropriate price in a probabilistic equation of this sort is the "expected" out-of-pocket price, i.e., a price which itself reflects the essential uncertainty of the event. Price uncertainty "occurs because the consumer is not given a price per unit and asked how many units he wants, but rather an uncertain price for an entire stay (depending on

length, tests ordered, etc.) and asked whether he wants exactly one unit (admission)."⁸¹ While the expected net price is not directly observable in any data set, Newhouse and Phelps show that it is proportional to the coinsurance rate in policies that do not contain a deductible.⁸² Hence, they use the coinsurance rate as the "price" variable in equation (1).

Equation (2) relates exclusively to the discharge decision, or equivalently, to the length of stay decision. Since this decision can only be taken in respect of "users", the dependent variable is defined as length of stay conditional on admission. The theoretically appropriate "price" in this equation is the marginal point-of-service price, and this Newhouse and Phelps operationally and accurately define as the marginal coinsurance rate times the gross price per day of care.

To estimate these (and their model's other) equations, Newhouse and Phelps use data drawn from the 1963 household interview survey conducted by the National Opinion Research Center (NORC) and the Center for Health Administration Studies (CHAS) of the University of Chicago. The NORC-CHAS Survey is a probability sample of the civilian, non-institutionalized population, and provides detailed health care utilization and insurance coverage data on 2,367 families and 7,803 family members. From this sample, Newhouse and Phelps identify 4,522 individuals who satisfy various criteria appropriate to the estimation

⁸¹Newhouse and Phelps, op.cit., p. 271.

⁸²Ibid., p. 272.

of equation (1).⁸³ The vast majority (92 per cent) of this group are "nonusers," i.e., individuals who did not experience any hospitalization during 1963, and they in turn are excluded from the sample for purposes of estimating equation (2); there are 364 individuals in the "user" sub-sample.

With the relevant use/nonuse data in hand, Newhouse and Phelps proceed with the empirical analysis of admission probabilities. Specifically, they fit these data to a non-linear form of equation (1) using the maximum likelihood estimation method.⁸⁴ The results of this experiment indicate that: (1) the wage, sex, age, education, family size, and family income elasticities of admission are uniformly insignificant, contrary to expectations based on Grossman's theory; (2) the individual's race (white, non-white) and self-perceived health status (poor, fair, good, excellent) are the most important predictors of admission; (3) the cross-price elasticity of admission, i.e., the elasticity with respect to the coinsurance rate on physician office visits, is significantly negative (-.57), suggesting that hospital and

⁸³Consistent with their operational definition of the expected net price, Newhouse and Phelps restrict the sample to individuals whose hospital insurance policies did not contain a deductible in 1963. To further ensure accurate calculation of the marginal coinsurance rate, they also include only those individuals whose policies could be verified with the insurance company. For more details on the sample, see *Ibid.*, p. 274.

⁸⁴In its estimated form, equation (1) is expressed as a logit function, i.e., as: $P(A) = 1/(1+e^{Bx})$, where x is a vector of explanatory variables. Being non-linear in B , this regression model has the desirable property of constraining the estimated probabilities of admission to lie within the a priori correct unit (0,1) interval.

physician services are "strongly complementary;"⁸⁵ (4) the probability of admission is slightly higher for individuals resident in communities where hospital beds are relatively abundant, a finding consistent with the operation of a "pure" availability effect; and (5) the coinsurance elasticity of admission is -0.17 and statistically significant. This last finding is, of course, the one most relevant to this review. It indicates that a 10 per cent increase in the (room and board) coinsurance rate leads to a very small 1.7 per cent decrease in the probability of admission.

Equation (2), the length of stay equation, is fitted in linear form to the "user" sub-sample via the ordinary least squares (OLS) procedure, and alternatively, the two-stage least squares (TSLS) procedure. This experiment indicates that: (1) length of stay increases by approximately one day for each 10 year increase over the age range 25-65+ years; (2) length of stay is invariant to the individual's sex, education, wage rate, family income, family size, and self-perceived health status, but does vary with the individual's race (non-whites stay 3.6 days longer than whites); (3) length of stay is unaffected by the local availability of beds and physicians; (4) the cross-price elasticity of length of stay is insignificant, i.e., an increase in the point-of-service price of a physician's office visit has no effect on length of stay; and, most important, (5) the net price elasticity of length of stay is $+0.05$ and insignificant when estimated via TSLS; and

⁸⁵Ibid., p. 273. They note, however, that the sample contains "so few individuals with insurance for physicians' services that it is hard to place a great deal of weight on this conclusion."

-0.06 and significant when estimated via OLS.⁸⁶

As mentioned at the outset of this review, Newhouse and Phelps' estimates of the price elasticity of hospital use (-0.17 for admissions and -0.06 for length of stay) are the lowest of any in the literature. When appraising this fact, the authors fairly claim to have achieved a greater degree of accuracy than their fellow health economists. They note (as we have) that data deficiencies, especially errors in measuring the price variable, militate against acceptance of the elasticities generated, for example, by M. Feldstein, and Davis and Russell.⁸⁷ At the same time, however, they duly acknowledge the weaknesses of their survey data. These data are

Far from ideal ... Insurance coverage is measured at the end of the year, and there is no guarantee that the insurance was in force throughout the year ... There is variation in [the benefit structure of] insurance policies that we have not measured ... Most important, the TSLS results are almost never significant, so the possibility that all elasticities are in fact zero cannot be dismissed on the basis of these data. Thus the estimates are fragile.⁸⁸

Quality differences aside, the studies reviewed thusfar are similar in that they all reflect a methodological preference for isolating price effects through multi-variate regression techniques. Lest the impression be conveyed that this is standard research practice, brief reference should be made to a third and final group of studies

⁸⁶In this connection, the authors note that "neither the OLS nor the TSLS are ideal; the OLS estimates ... are inconsistent away from zero, if there is advance selection of insurance, but the TSLS estimates are very imprecise and may in fact have larger mean square errors than the OLS estimates. We have solved this problem by accepting the OLS estimates, but increasing their standard errors so that a 95 per cent confidence interval would include zero." Ibid., p. 282.

⁸⁷Ibid., p. 285.

⁸⁸Ibid.

wherein price effects are inferred by simply comparing the mean utilization levels of differentially insured populations. To the extent the comparison populations are similar in all respects other than insurance coverage, this strategy obviates the need for formal multi-variate analysis. In practice, however, these comparisons are more often than not invalidated by data inadequacies, and by the pervasive and previously noted problem of advance selection. A small, but nevertheless representative sample of studies will illustrate the unsophisticated flavour of this approach.

Using data from a survey of geographically homogeneous Blue Cross Plans in 1964, Williams compared the hospital use of enrollees in a plan featuring full coverage (i.e., no point-of-service charges) with the hospital use of enrollees in a plan featuring a flat \$4 per day copayment charge.⁸⁹ After adjusting for age and sex differences, he found that the full coverage group had an admission rate of 156 per 1000 enrollees, while the copay group had an admission rate of 135 per 1000. (The lengths of stay were virtually identical.) This suggests that the copayment charge, roughly equivalent to a twelve per cent coinsurance rate, reduced utilization by 15.5 per cent. The implied arc price elasticity of admissions over this range of coinsurance is, at -0.07,⁹⁰ well within the range of weak price effects. There is, however, a definite upward bias in this estimate due to the fact enrollees were

⁸⁹R. Williams; "A Comparison of Hospital Utilization and Costs by Type of Coverage," Inquiry, 3:28, September 1966.

⁹⁰Williams does not express the result in these terms, but it is easily calculated using the arc elasticity formula

$$\eta = (Q_1 - Q_0 / Q_1 + Q_0) \cdot (P_1 + P_0 / P_1 - P_0)$$

and the values:

$Q_1=135; Q_0=156; P_1=4.0; P_0=0.0$. Thus $\eta = -27/291 = -0.07$.

able to select into the plan of their choice.

In a similar study, Hardwick *et al.* compared the 1969 inpatient claims experience of two Pennsylvania Blue Cross groups, one with full-coverage and the other with a \$5 per day copayment charge for the first fifteen days hospitalization.⁹¹ They found no significant difference between the two groups for five age-sex adjusted measures of utilization, namely, average length of stay, average benefits per admission, average benefits per day, admissions per 1000 enrollees, and patient-days per 1000 enrollees. However, as two critics of the study have noted, these results "should probably be ignored," since no allowance was made for the disproportionate number of rural (urban) residents in the copay (non-copay) group.⁹² A more accurate assessment of the copayment effect would require a multi-variate analysis in which residence-related differences in bed supply, income, education, and other determinants of use are held constant.

More, albeit indirect, evidence on price effects is contained in a study by Heaney and Riedel.⁹³ They analyzed utilization patterns over the 1966-68 period during which the Connecticut Blue Cross changed half its group plans from indemnity insurance (paying a cash benefit of \$15 per day of hospital care) to full semi-private coverage. Expectations were that subscribers who voluntarily switched from the indemnity to

⁹¹C. Patrick Hardwick, Larry Shuman, and Shlomo Barnoon, "Effects of Participatory Insurance on Hospital Utilization," Health Services Research, Vol. 7, No. 1, Spring 1972.

⁹²Paul B. Ginsburg and Larry M. Manheim, *op.cit.*, p. 146.

⁹³Charles T. Heaney and Donald C. Riedel, "From Indemnity to Full Coverage: Changes in Hospital Utilization," Blue Cross Reports, Research Series No. 5, (Chicago: The Blue Cross Association, October, 1970).

full coverage plan would experience a price-induced increase in admissions and length of stay. To test this hypothesis, comparisons were made of their claims data for the six months before and the six months after the change in insurance coverage. For five of seven size specific "switcher" groups, it was found that the admission rate increased following the increase in coverage, but in no instance was the change statistically significant. Nor was the overall increase in admissions from 4.56 to 5.09 per 100 enrollees statistically significant. Correspondingly, there was an insignificant increase in average length of stay from 7.26 to 8.17 days. Evidently, therefore, the implicit net price elasticities of admissions and lengths of stay were effectively zero in these data.⁹⁴ It should be noted, however, that no attempt was made to adjust the utilization data for seasonal and other extraneous factors. Had this been done, the results might well have differed substantially.

Conclusion

Taken together, the studies reviewed in this section illustrate the "state of the art" of applied research on price effects in hospital use. On the surface, the impression is one of considerable progress and rapid achievement within a relatively short time span of some ten to twelve years. At this level there is no denying M. Feldstein's summary

⁹⁴Newhouse and Phelps have shown that the (insignificant) arc price elasticities are respectively -0.04 and -0.03 over the relevant (zero to 25 per cent) range of coinsurance rates. See Charles E. Phelps and Joseph P. Newhouse, Coinurance and the Demand for Medical Services (Santa Monica: Rand Corp. R-964-1-OEO/NC, October 1974), pp. 24-25.

judgment that the accumulated effort provides "much more information and a better sense of the issues than just a few years ago."⁹⁵ At another, more critical level, however, the dominant impression is much less charitable. Faulty theoretical formulations, inappropriate methodologies, and inadequate data have all too often been the rule rather than the exception. Paradoxically, the prevalence of these deficiencies makes final appraisal of the evidence that much easier.

Central to this assessment is the ineluctable conclusion that none of the observed price effects can be regarded as unbiased estimates of the true price elasticity of demand for hospital care. The reason for this is quite straight-forward and, it should be stressed, has nothing to do with the statistical methods employed. Rather, it follows from the fact that no study has successfully disentangled the preferences of the consumer-patient and the provider-physician in such a way as to reveal the identity of the demand and "need" curves with which, in theory, these preferences are respectively associated. In the absence of such efforts, there can be no claim that the observed relationships neatly coincide with either one of these distinctly different concepts. To be sure, solution of this "identification problem" was not a stated objective of any of the studies, nor were any claims to solution made, but this only reinforces the case for cautious interpretation of results. The appropriately neutral term to describe price effects in the all-too-common instances where the principal decision-maker cannot be identified is the price elasticity of use.

Equally ineluctable is the conclusion that methodological

⁹⁵M. S. Feldstein, "Econometric Studies of Health Economics," in M. D. Intriligator and D. A. Kendrick, (eds.), op.cit., p. 388.

problems and data deficiencies seriously erode the credibility of observed price effects even when interpreted as use elasticities. Most vulnerable to this criticism are those studies which have sought price effects in macro-data sets. This review has challenged the case for further interrogation of these data. Research based on micro-data sets affords greater opportunity to establish the price-use relationship(s), but investigators who have chosen this strategy have also encountered methodological problems, particularly in operationally defining the relevant net price.

To date, the most successful micro-analytic effort in terms of both theoretical and statistical rigour is the Newhouse and Phelps study. Their evidence indicates that admission and discharge decisions are weakly sensitive to variations in the coinsurance rate and marginal net price of care respectively. Thus, in the literature, the best available answer to the question "Are there price effects governing hospital use?" is "Yes, but they are much smaller than many have hitherto believed."

Of course no study is totally immune from criticism and the Newhouse and Phelps research is no exception. Mention was made of the fact that their omnibus definition of hospital use is one which makes no allowance for variations in price effects across different diagnostic and/or treatment categories, or across different institutional settings. The implicit assumption that all utilization decisions are equally (non) discretionary, and hence equally price (in)sensitive is clearly unrealistic, given the wide differences in case severity manifest over the entire spectrum of medical problems for which hospitalization may be indicated. It follows, as two reviewers of the coinsurance literature

concluded,

There is a need for further studies of the problem ... As there are many price elasticities instead of just one, a particularly important need is for disaggregated studies.⁹⁶

The present study represents an attempt to fill this void. It conducts the search for price effects at a highly disaggregate level in a body of uniquely appropriate micro-data. In addition, it employs a measure of net price which more closely approximates the theoretical construct than those in the majority of previous studies. To the details of this research strategy and design, the discussion now turns.

⁹⁶Ginsburg and Manheim, op.cit., p. 152.

CHAPTER III

THE MICRO-ANALYTICAL FRAMEWORK

The overall purpose of this chapter is to present and discuss the elements of the micro-analytical framework on which the remainder of this investigation is based. Specific objectives are: in section (1) to describe the nature and sources of the data that are available to this study; in section (2) to explain the method by which these data are sampled, and to examine the representativeness of the resultant family sample; in section (3) to describe and define the component variables of the summary record constructed for each of the sample families; and, in section (4), to explicate the statistical model and associated estimation procedures used to ascertain the hypothesized influence of price and other factors governing families' use of inpatient facilities. This sets the stage for chapters 4 and 5 where the observed price-use and other relevant relationships are reported and analyzed.

The Data Base

It will be recalled from Chapter 1 that this study owes its feasibility to the "copayment experiment" conducted in the Province of Saskatchewan over the period 1968-1971. The details of that experiment

and the programmatic setting within which it took place were noted in that chapter, as were the broader geo-demographic features of the province itself. Attention now shifts to the information system maintained by the Saskatchewan Hospital Services Plan (SHSP), the agency which has had administrative responsibility for the province's hospital insurance program since its inception in January 1947. Relevant details of the agency and its information system will be explained in this section, since together they represent the primary source of data for this study.

The present SHSP information system can, to a very large extent, be seen as the product of the government's original decision to partially finance the hospital insurance program from the proceeds of a "hospitalization tax." This tax took the form of an annual premium levied upon all residents of the province. Payment of the premium was compulsory (as was participation in the Plan itself), and was the means by which eligibility for benefits was determined. Collection of this tax became one of the major administrative responsibilities of SHSP, and pursuant to this function, the agency developed a central registry of program beneficiaries.¹ To this day, and notwithstanding the abolishment of the premium in 1974, the Master Registration File (MRF) remains one of the key elements in the SHSP information system.

¹Records of registration and tax collection were channelled to SHSP through municipal offices located in cities, towns, villages, and rural areas. As compensation for this service, SHSP paid the local taxation offices a percentage commission on the monies collected. A former senior SHSP official cites this arrangement as one factor which has tended, over time, to keep tax delinquency to a minimum, since it gave the local governments a financial stake in maintaining adequate surveillance. See J. D. Ramsay, "The Hospital Discharge Abstract System of Saskatchewan Hospital Services Plan," Medical Care, Vol. VIII, No. 4, Supplement (July-August 1970), p. 146.

The structure of the MRF is best illustrated with reference to its handling of a given "family." For administrative reasons, a family is defined as a self-supporting person of any age, together with his spouse and any dependents under the age of eighteen. A single person over eighteen (whether self-supporting or not) is also defined as a family. Upon registration, an individual is assigned an eight-digit identification number, of which the first six digits are his family number, and the last two digits are his beneficiary number. This procedure guarantees that every member of a multiple-person family receives a unique registration-beneficiary number. For example, if, in a four person family consisting of husband, wife, and two dependent children, the assigned family number is 123456, the assigned beneficiary numbers would be 123456 00 (husband), 123456 01 (wife), 123456 02 (first child), 123456 03 (second child). The "family record" portion of the MRF includes the six-digit family number and, as well, information on family residence and family size. The "beneficiary records" which follow contain information specific to each family member. Included here are the individual's age, sex, and marital status, and an indicator of his or her "activity status." The latter item designates the individual's eligibility for benefits; "inactive status" applies to individuals who die or leave the province. Supplementary and more detailed information on eligibility for benefits is contained in the "coverage entries." These were particularly important during the premium period (1947-1974), when they indicated whether or not the individual was responsible for payment of the tax (welfare recipients were among those exempted from the tax), and the dates from which and to which the taxes had been paid.

Mention should also be made of the fact that under this registration system certain individuals will receive more than one identification number over the course of their enrollment. Thus, any individual who is initially registered as a dependent under the age of eighteen will, upon turning eighteen, be defined by SHSP as a separate "family," and, accordingly, be assigned a new eight-digit number. In the case of the four person family illustrated above, re-registration of the two children on their eighteenth birthdays would entail assigning two new and unique numbers, for example, 654321 00 (first child), and 754321 00 (second child), where in each case the double-zero beneficiary number indicates the individual's status as an SHSP "family head." If both these individuals are males, they will retain their new numbers as long as they remain enrolled in the program. However, if one (or both) of them is a female, she will retain her new number only for as long as she remains single. When she marries, she will be re-registered for a second time under her husband's family number and assigned an appropriate (01) beneficiary number. To accommodate these number change eventualities, the beneficiary record opposite the new (old) registration number contains a cross-reference to the old (new) registration number, and the date of re-registration. As well, the coverage entries in the old and new beneficiary records indicate the (non-overlapping) time periods during which the individual was eligible for benefits under each registration number.

In its most comprehensive version, the registration file contains a cumulative list of all family and beneficiary numbers assigned since the start of the hospital insurance program. Inspection of this file and the associated coverage entries would, of course,

reveal a wide variety of enrollment histories. Some individuals - those who were residents of the province in 1947 and every year thereafter - will have records indicating their enrollment in the program during each and every year of its operation. Others will have records indicating their enrollment for variably shorter periods of time. Included here are all those individuals who took up residency in the province and/or who ceased residency in the province anytime after 1947. Their initial dates of registration will vary with their dates of in-migration or birth, while their terminal dates of registration will vary with their dates of out-migration or death. Still others, who could belong to either one of the preceding groups, will (for reasons mentioned above) have records under more than one registration number by virtue of their having turned eighteen and/or having married during the course of their enrollment. Thus, the most comprehensive version of the MRF contains detailed demographic data on the insured population for every year since 1947.

It must be hastily added, however, that this complete file of historical number assignments is not maintained by SHSP in a form suitable to applied research. It was not until 1962, the year when the province implemented its medical care insurance program ("Medicare"), that the registration records were converted to machine-readable formats. To expedite the conversion process and to minimize its costs, the records of the then "non-active" beneficiaries, i.e., those who had died or left the province anytime between 1947 and 1961, were left in their original state, and hence were not included on the computerized MRF. When generating this smaller file, the coverage entries of the "active" beneficiary records were revised to include information on the

individual's eligibility for ~~benefits~~ under both the hospital and medical care insurance programs. By this relatively simple procedure, the Medical Care Insurance Commission (MCIC), the agency set up to administer the Medicare program, was able to avoid the costly exercise of assigning its own registration numbers and developing its own master registration file. In the years since 1962, this dual-purpose MRF has been updated on a monthly basis to include the family and beneficiary records of all new SHSP/MCIC "families." Thus, the population list available for practical research purposes includes only those individuals who have been eligible for benefits under either or both of the two insurance programs at any time since 1962. This constitutes one of several minor constraints imposed upon the present study.

Attention now shifts to another component of the SHSP information system central to the development and execution of this study. Known informally as the Hospital Claims File, this data set contains records of the inpatient experiences of all SHSP beneficiaries in any specific year or years since the inception of the program. For statistical reporting and cost-accounting purposes, the agency stores these records in separate annual files. Within each file, the records are identified by the beneficiary's (eight-digit) registration number, and are ordered sequentially on this basis. No records appear on this file for any beneficiary who did not experience at least one hospitalization during the period to which the file refers. Only "users," i.e., those individuals who have had inpatient status during the period, will have records on this file.

Each record is compiled from the Admission-Separation form submitted to the agency by the hospital within forty-eight hours after

the discharge or death of the patient. This document contains all the information required by SHSP for its eligibility checking and coding procedures.² Included are: the patient's name, address, age and date of birth, sex, marital status, and registration number; the admitting hospital; details of the admission history (date and type of admission, provisional diagnosis, name of admitting physician, etc.); details of the separation history (date and type of separation, primary and secondary discharge diagnoses, nature of any operation or procedure performed, name of attending surgeon and surgical assistant, etc.); and details of the accounting record (type, quantity and cost of accommodation payable by SHSP, utilization fees payable by patient, etc.). A facsimile of the Hospital Admission-Separation Record is shown in Figure 3.1.

At SHSP these forms are routinely subjected to a variety of checks designed to ensure their completeness, accuracy, and reliability. In the event of unsatisfactory or missing information, they are

Returned to the hospital for amendment, and a copy of the deficient document is sent to the Medical librarian consultant employed by the Regional Hospitals Council. The consultant visits the hospital and advises on the situation. In this way, a continual quality check is maintained.³

Needless to say, this surveillance enhances the attractiveness of the claims file for research purposes. It minimizes the probability of

²Prior to 1969, the eligibility check was done by a manual comparison of the form with the file copy of the Medical and Hospital Identification Card carried by the beneficiary. Since 1969, verification of the patient's eligibility, age, sex, residence, and marital status has been done with reference to the Master Registration File. Source: Personal communication with Mr. D. Crighton, Claims Section, Taxation and Claims Division, S.H.S.P., October 2, 1973.

³Ramsay, op.cit., p.148.

Hospital Admission — Separation Record

SASKATCHEWAN HOSPITAL
SERVICES PLAN — REGINA, SASK.

HEALTH STATUS
DATE: 9/10/77

NAME OF HOSPITAL: _____

NAME OF PATIENT — SURNAME: _____ GIVEN NAME: _____ INITIAL: _____

ADDRESS: _____

IF PATIENT IS DEPENDENT — GIVE SURNAME, GIVEN NAME AND INITIAL OF HEAD OF FAMILY: _____

RESPONSIBLE FOR PAYMENT OF PUBLIC WARD CARE:
☐ S.H.S.P. ☐ NON-RESIDENT OF SASKATCHEWAN
☐ W.C.B. ☐ UNINSURED SASKATCHEWAN RESIDENT
☐ D.V.A. ☐ OTHER AGENCY: _____

ADMISSION HISTORY
 DATE OF ADMISSION: _____ TIME: _____
 MARITAL STATUS: _____
 TRANSFERRED FROM — NAME OF HOSPITAL OR INSTITUTION: _____
 DAYS CHARGED FOR UTILIZATION FEES: _____
 TYPE OF ADMISSION: EMERGENCY ☐ URGENT ☐ ELECTIVE ☐
 ACCIDENT OR INJURY REPORT — THE UPPER COPY OF THIS ACCIDENT REPORT MUST BE COMPLETED ON EACH ACCIDENT OR INJURY CASE.
 OCCUPATION OF HEAD OF FAMILY: _____
 EMPLOYER'S NAME AND ADDRESS: _____
 NEAREST RELATIVE OR FRIEND: _____ RELATIONSHIP: _____
 ADDRESS: _____ PHONE NO.: _____
 ADMITTED BY: _____
 I HEREBY REQUEST _____ NURSE
 ACCOMMODATION AND AGREE TO ASSUME RESPONSIBILITY FOR CHARGES NOT COVERED BY THE SASK. HOSPITAL SERVICES PLAN OR ANY OTHER AGENCY.
 I CERTIFY THAT THE ABOVE PATIENT RECEIVED SERVICES INDICATED

SEPARATION HISTORY
 DATE OF DISCHARGE: _____ TIME: _____
 DIAGNOSIS ON DISCHARGE — PRIMARY: _____
 SECONDARY DIAGNOSIS OR COMPLICATIONS: _____
 NATURE OF OPERATIONS OR PROCEDURES: _____
 DATE PERFORMED: _____
 NAME OF SURGICAL ASSISTANT: _____ NAME OF ANESTHETIST: _____
 TYPE OF SEPARATION:
☐ MEDICAL ☐ DISCHARGE ☐ DISCHARGE ON ORDER ☐ DEATH ☐ TRANSFERRED TO NAME OF HOSPITAL OR INSTITUTION: _____
 ACCOUNTING RECORD

	DAYS	RATE	AMOUNT
PUBLIC WARD OR NURSERY			
SEMI-PRIVATE OR PRIVATE			
UTILIZATION FEES — PAYABLE BY PATIENT			
UTILIZATION FEES — PAYABLE BY P.C.B. NO. OF DAYS			
NON-PHENOL DRUGS			
OTHER CHARGES			
GRAND TOTAL			
LESS CHARGES TO OTHER AGENCY			
LESS UTILIZATION FEES — PAYABLE BY P.C.B. NO. OF DAYS			
PAYABLE BY PATIENT			

 DO NOT WRITE BELOW THIS LINE

DATE OF BIRTH	DATE OF DISCHARGE	PRIMARY DIAGNOSIS	SECONDARY DIAGNOSIS	OPERATION PERFORMED	ATTENDING SURGEON
23 25 27 29 31	01 03 05 07 09	01 03 05 07 09	10 12 14 16 18	20 22 24 26 28	30 32 34 36 38
33 35 37 39 41	43 45 47 49 51	53 55 57 59 61	63 65 67 69 71	73 75 77 79 81	83 85 87 89 91

CONTROL REFERENCE NUMBER: _____

Fig.3.1. Page 1 of the Admission/Separation record of the Saskatchewan Hospital Services Plan. Page 2 of this four-part form duplicates the top third of page 1, with the remainder specially designed for reporting accidents or injuries to SHSP, to facilitate the detection of possible third-party liability. Pages 3 and 4 are identical to page 1 except for the final section at the bottom of the page, and are retained by the hospital as part of the patient's business office record and medical record, respectively. See Ramsay, *op.cit.*

observational error.

Subsequent to its verification, the Admission-Separation Record is translated into a machine-readable "claim-card" by SHSP's data processing division. This essentially entails the coding of selected items of written information into numeric formats. Some notable examples are: (1) the names of the attending physician and attending surgeon shown on the record are converted into the four-digit numbers under which each physician is uniquely identified in the central (MCIC) register of practitioners licensed and resident in the province; (2) the primary and secondary diagnoses written in the "separation history" portion of the record are coded according to the International Statistical Classification (ISC) of Diseases, Injuries, and Causes of Death,⁴ (these basic data are, as will be subsequently shown, an integral part of the disaggregation scheme employed in this study); (3) the accident information reported on page 2 of the A-S Record (not shown in Figure 1) is also coded according to the ISC (ICDA-8) codebook; and (4) the surgical or other procedure (if more than one, "the most

⁴Originally limited to 179 causes of death identified by the International Statistical Institute in 1900, this system has evolved into a comprehensive list of well over 1000 rubrics for the reporting of both mortality and morbidity statistics. Under the auspices of the World Health Organization, it has been revised on two occasions since its first adoption by SHSP in 1949. Written diagnoses were coded according to the ISC, 6th Revision (1948) until January 1, 1958 when the agency began coding according to the ISC, 7th Revision (1955). Ten years later, on January 1, 1969, the agency implemented the ISC 8th Revision (1965) or more accurately, the American Hospital Association's adaptation of the 8th Revision, known colloquially as the "ICDA-8" system. For further historical details of the International Classification of Diseases and the complete tabular list of the four-digit subcategories included in the ICDA-8 system, see U.S., Department of Health, Education and Welfare, Eighth Revision International Classification of Diseases Adapted for Use in the United States, Vol. 1, Tabular List, Public Health Service Publication No. 1963 (Washington: U.S. Government Printing Office, 1972).

important") described in the separation history is (since January 1, 1969) coded according to the Classification of Surgical Operations, Diagnostic and Other Therapeutic Procedures prepared by the American Hospital Association for the United States Public Health Service.⁵ Other items routinely coded on the claim card include the patient's registration number, age, sex, residence code, family size,⁶ and length of stay, and the hospital's location and three-digit identification number.

By preserving the salient details of each admission in this way, SHSP has, over ~~the~~ years, generated an accurate and comprehensive record of the annual utilization of hospital facilities by the insured population. Retrospective analyses of these data are, however, complicated by two factors. First, the roughly two hundred thousand claim cards contained in each of the annual files are retained in the SHSP statistical library for a maximum of ten years. This implies that any research requiring a time series of observations on hospital use at the beneficiary and/or family level is unavoidably constrained to the ten annual files which exist under the retention schedule prevailing at

⁵For the complete tabular list of procedures and their corresponding three-digit codes, see Ibid. From 1947 to 1961 SHSP coded procedures according to a system developed by the Saskatchewan Department of Public Health; from 1962 to 1968 it used a system known as the American Classification of Operations and Treatments. In detail and comprehensiveness, this latter system was very similar to the one implemented in 1969. Only the codes were changed to frustrate the researcher!

⁶Prior to 1969, the patient's residence code (three digits for municipality, two digits for place within municipality) and family size were written on the Admission-Separation Record by personnel in SHSP's Registration Division during their manual check of the patient's eligibility. Since 1969 this information has been added during computer processing of the claim card against the Master Registration File.

the time the research is initiated. Second, it has not been standard SHSP practice to maintain these annual files in a form consistent with least-cost computer retrieval and processing techniques. Understandable as this is, it does impose another constraint on the present study. Thus the only claim files to which ready access has been gained are those for the six year period 1966-71.⁷ Notice, however, that two of these years (1966-67) precede the copayment experiment (1968-71), so there is an opportunity to sift these data (or samples thereof) for "price effects."

To summarize, the SHSP information system consists of two basic data files, one a master registry of program beneficiaries, and the other a claims file of the inpatient experiences of these beneficiaries. The Master Registration File (MRF) contains socio-demographic information on each beneficiary and his or her family including age, sex, marital status, social assistance status, residence, and family size. The claims file details for each hospitalization such facts as type of admission, length of stay, principal and secondary diagnoses, surgical procedure, and attending physician and/or surgeon. An individual identified on the registration file may or may not be identified on the claims file. It all depends on whether or not he was a user of insured hospital services during the period in question. By linking the two files via the registration-beneficiary numbers, it is thus possible to divide the insured population into "user" and "non-user" groups. This property of the SHSP information system will be

⁷This study was initiated in the summer of 1972. At that time, and due to the sampling strategy employed (about which more will be said in the next section), these were the only claim files available for analysis.

further discussed in subsequent sections of this chapter. It figures prominently in the overall research design.

Finally, brief mention should also be made of another data set, unrelated to the hospital insurance program, on which this study will place considerable, albeit secondary, reliance. These are the income tax files compiled annually for the Saskatchewan Treasury Department by the Federal Department of National Revenue. The representative record in these files preserves most of the information contained on the "T-1" form filed by individuals. Included are the individual's name, age, address, total income, personal exemptions, taxable income, and tax liability, etc.⁸ Records are identified by surname and are stored sequentially on magnetic tapes by the first five characters of the surname. Access to these records is of course strictly controlled. The Statistics Act requires that no record shall be released in a form which enables identification of the individual to whom it relates. Thus, at this early juncture, it is well to stress very emphatically that, while access to these data has been gained, no breach of the confidentiality requirement has occurred in their retrieval, assembly, and processing.

⁸The reliability and validity of the income information depends upon: (1) the frequency and magnitude of unintentional and intentional reporting errors; and (2) the definition of income. Surveillance and audit techniques used by the Department of National Revenue have traditionally acted as disincentives to under-reporting of income. More problematic is the "narrowness" of the income definition. It excludes income from capital gains (in all years prior to 1972 when the capital gains tax was introduced), and, as well, any income-in-kind. Both omissions are conceptually significant (especially the latter, since income-in-kind has doubtless been important among Saskatchewan families for some time), but there is, unfortunately, no relevant, extraneous information available which might be used to adjust the reported income figures. In any event, the income data on the tax files is arguably superior in quality to any that might be obtained by alternative, and more costly methods such as mailed questionnaires or household interview surveys.

All linking of the records with the SHSP registry of program beneficiaries has taken place under strict government supervision.

Sampling Method and Sample Data

As mentioned above, the data available to this study consist of the registration records and the inpatient histories for all individuals insured by the Saskatchewan Hospital Services Plan at any time during the six year period 1966-71. While it is technically possible to base the study upon the entire data set, resource constraints militate against this strategy. Accordingly, the study is based upon a random sample of the available records. In this section, the sampling methodology and the representativeness of the resultant sample are discussed in turn.

The population frame for sampling purposes is the Master Registration File. The rationale for choosing this file over the functionally independent Hospital Claims File is quite straightforward. The MRF is the only one of the two which identifies all (past and present) SHSP "families," and, hence, is the only one which permits the selection of a truly representative sample of the provincial population. The claims file, it will be recalled, relates only to the sub-population of "users," i.e., those who experience hospitalization. To facilitate a full and complete analysis of the effects of copayment on families' access to and use of hospital services, it is desirable that the sample include both "users" and "non-users."

The following method was employed to select the sample for this study. First, from the cumulative list of all family registration

numbers contained on the machine-readable MRF, a ten per cent random sample was drawn using a "count and random skip" algorithm.⁹ Second, this overall sample was sorted into annual sub-samples by consulting the coverage entries in the beneficiary record(s). Recall these entries indicate the time periods from which and to which the individual family members were "active," i.e., eligible for benefits. Allocation of the overall sample into nine annual sub-samples was thus effected by identifying those families having one or more beneficiaries eligible for benefits during any one or all of the twelve month periods defined by the calendar years 1963-71. Of these sub-samples, only those for the years 1966-71 are used in this study, since, as mentioned in the previous section, the corresponding hospital claims data are not available in machine-readable form prior to 1966.

⁹In point of fact, the sample is an updated version of the one earlier employed by Professor R.G. Beck, University of Saskatchewan, to analyze the effects of utilization fees on family utilization of physicians' services. The original sample was drawn on June 26, 1969 from the cumulative list of all family registration numbers assigned since the inception of the medical care plan in 1962; at the time of sampling, the list contained approximately 540,000 registration numbers, and from these a random sample of 54,000 registration numbers was selected. For further details see R.G. Beck, "Analysis of the Demand for Physicians' Services in Saskatchewan," (unpublished doctoral dissertation), University of Alberta, 1971. This sample was subsequently updated on February 11, 1972, immediately prior to the initiation of the present study. The update procedure involved selecting a (10 per cent) random sample from the cumulative list of the registration numbers assigned by SHSP/MCIC between June 26, 1969 and February 11, 1972, i.e., between the dates of the first and second sampling; at the time of second sampling, this much shorter cumulative list contained approximately 78,000 family registration numbers, and from these a random sample of 7,800 numbers was selected. Thus, the updated sample contains approximately 61,800 family registration numbers, (comprising 54,000 "original selects" and 7,800 "new selects.") As subsequently noted in the text, however, this random sample contains families with a wide variety of enrollment histories. To determine their "activity" status over the study period, it is necessary to consult the coverage entries where the enrollment history of each family member is described.

2

It bears emphasizing that, under the above random select and sort procedure, a given family is included in as many of the annual sub-samples as are consistent with the enrollment histories of its individual members. If, for example, the coverage entries indicate that all (or some) of the beneficiaries were active throughout the entire six year study period, then the family will be included in each and every one of the annual sub-samples. If, on the other hand, the coverage entries indicate that the family was registered for less than the full six years, it will only be included in the annual sub-samples corresponding to those discrete twelve month periods when it was active. It could, for example, be in the 1969, 1970, and 1971 sub-samples, but not in the 1966, 1967, and 1968 sub-samples by virtue of its "non-existence" during the first three years of the study period. Non-existence could mean that the family was not resident in the province prior to 1969, or it could mean that the individual did not become an SHSP "family" until 1969, the year of his or her eighteenth birthday. In general, all of the factors (birth, death, in-migration, out-migration) determining activity status described in the previous section will dictate the entry and exit of families to and from the annual sub-samples. Thus it is a property of the annual samples that they do not contain a common set of families (or individuals). Rather they reflect the demographic dynamics of the parent provincial population from which they have been drawn.

A schematic representation of the annual samples is shown in Figure 3.2. Comparing the size of any two consecutive samples reveals the net effect of all intervening changes in family status from inactive to active, and vice versa. Each sample contains approximately 38,000

Figure 3.2
Schema of Samples

families, and represents a cross-section random sample of all SHSP families active in each of the respective years.¹⁰

The representativeness of the annual samples can be evaluated along any one of several demographic dimensions for which comparable population data are available. Table 3.1 presents the distribution of active families by size of family for each of the annual samples and for the corresponding SHSP insured population. Inspection of the table suggests that the samples are representative of their respective parent populations. This has been confirmed using the Chi-Square test for "goodness-of-fit." In each case, the calculated value of the Chi-Square statistic is well below its critical value, and hence the null hypothesis that the samples are representative cannot be rejected.¹¹ Additional tests using other information contained in the registration records (e.g., age of head, family residence, marital status, social assistance status, etc.) would doubtless yield very similar results. However, in the absence of the requisite population distributions, it was not possible to perform these supplementary determinations.

There is, of course, no reason to base the tests of representativeness solely upon information contained in the registration records. These records are after all only one of several types of data

¹⁰The samples are not, however, strictly equivalent to the independent random samples that would have resulted had the Master Registration File been repeatedly sampled at the end of each year, beginning in 1966. On this hypothetical procedure, a family that was active for the entire period would not necessarily appear in all of the sub-samples. Hence, while the actual samples are truly random, they are not truly independent.

¹¹The calculated values of Chi square for the years 1966-71 are 0.46, 0.13, 0.14, 0.15, 0.16, and 0.13 respectively. For the appropriate degrees of freedom and the 5 per cent level of significance, the critical value of Chi square is 11.07.

TABLE 3.1
DISTRIBUTION OF ACTIVE FAMILIES* BY SIZE OF FAMILY
SAMPLE AND POPULATION DATA, 1966-71

Family Size	1966		1967		1968	
	Sample %	Pop. %	Sample %	Pop. %	Sample %	Pop. %
1	40.25	42.91	42.40	43.61	42.33	43.57
2	19.74	19.96	21.61	20.16	21.85	20.43
3	10.40	10.05	9.86	9.91	10.06	10.09
4	11.10	10.57	10.32	10.34	10.38	10.36
5	7.86	7.53	7.15	7.34	6.94	7.19
6+	10.44	8.97	8.66	8.64	8.44	8.35
Total No.	37,854	359,799	38,138	367,586	38,348	371,239

*Excludes Indians and War Veterans' Allowance Recipients.

Source: Population data from Saskatchewan Hospital Services Plan, Covered Population, 1966-71 (Regina: Queen's Printer).

TABLE 3.1 - Continued

Family Size	1969		1970		1971	
	Sample %	Pop. %	Sample %	Pop. %	Sample %	Pop. %
1	42.22	43.73	42.17	43.69	42.77	44.11
2	22.24	20.84	22.61	21.32	22.82	21.59
3	10.07	10.07	10.23	10.08	10.10	10.10
4	10.53	10.34	10.54	10.30	10.51	10.21
5	6.85	7.02	6.60	6.92	6.47	6.73
6+	8.09	8.00	7.84	7.68	7.34	7.27
Total No.	38,853	375,790	37,973	372,242	37,826	373,033

*Excludes Indians and War Veterans' Allowance Recipients.

Source: Population data from Saskatchewan Hospital Services Plan, Covered Population, 1966-71 (Regina: Queen's Printer).

collected for the families in the annual samples. The other integral components of the sample information system are the inpatient records drawn from the Hospital Claims File and the income records from the Income Tax File. These basic data are destined to play major analytical roles, so it is well to include them in any overall appraisal of representativeness. Each will be examined in turn.

The inpatient records contain many items of information which might, in principle, be used to examine the representativeness of the sample. In practice, however, the choice is limited to a few items (such as length of stay, principal diagnosis on discharge, surgical procedure) for which SHSP routinely compiles population data. Table 3.2 has been prepared using the diagnostic information. It shows the distributions of inpatient care by the broad diagnostic groupings known colloquially as the "18 break" of the International Classification of Diseases.¹² As expected, the sample and population distributions for each are very similar. Indeed, there are no statistically significant differences between the distributions,¹³ and hence it can be concluded that the sample inpatient data are representative.

The same conclusion applies in respect of the sample income

¹²Reference to this system of classifying diseases and conditions was made in Section 1. Here it should be noted that the data for 1966-68 are grouped according to the ISC-7th Revision (1955), while the data for 1969-71 are grouped according to the ISC-8th Revision (1965). Because these two editions of the International Classification involve different definitions of the "18 break", the 1966-68 distributions are not comparable to the 1969-71 distributions. The sample and population data are, however, consistently and comparably grouped in any given year.

¹³The calculated values of Chi square for the years 1966-71 are 0.45, 0.28, 0.29, 0.37, 0.34, and 0.62 respectively. For the appropriate degrees of freedom and the 5 per cent level of significance, the critical value of the Chi square is 27.59.

TABLE 3.2

DISTRIBUTION OF INPATIENT GENERAL HOSPITAL CARE*
BY DIAGNOSIS (ICD "18 BREAK")

DIAGNOSTIC GROUP**	1966		1967		1968	
	S*** %	Pop. %	S*** %	Pop. %	S*** %	Pop. %
Infective & Parasitic Diseases	1.2	1.4	1.4	1.4	1.1	1.2
Neoplasms	4.3	4.3	5.0	4.5	4.9	4.6
Allergic, Endocrine System, Metabolic & Nutritional Diseases	3.1	3.0	3.5	3.2	2.9	2.9
Diseases of the Blood & Blood-Forming Organs	0.7	0.6	0.6	0.6	0.5	0.6
Mental, Psychoneurotic, & Personality Disorders	2.3	2.3	2.4	2.4	2.4	2.3
Diseases of the Nervous System & Sense Organs	4.7	4.4	4.5	4.3	4.3	4.4
Diseases of the Circulatory system	7.9	7.8	9.0	8.2	8.8	8.4
Diseases of the Respiratory System	18.1	18.9	16.2	16.7	16.3	17.2
Diseases of the Digestive System	12.7	12.7	13.3	13.4	12.6	12.5
Diseases of the Genito-Urinary System	7.8	7.4	7.7	7.7	7.8	7.6
Deliveries & Complications of Pregnancy, Childbirth & the Puerperium	12.7	11.9	11.2	11.5	12.0	11.6
Diseases of the Skin & Cellular Tissue	1.8	1.9	1.7	1.9	1.6	1.9
Diseases of Bones & Organs of Movement	3.1	2.8	2.9	3.0	3.3	3.0
Congenital Malformations	0.6	0.6	0.6	0.6	0.6	0.6
Certain Diseases of Early Infancy	0.2	0.3	0.3	0.4	0.4	0.4
Symptoms, Senility & Ill-defined Conditions	2.9	3.0	3.0	3.1	3.0	3.1
Accidents, Poisonings & Violence	8.2	7.9	8.1	8.6	8.3	8.5
Supplementary Classifications	7.7	9.0	8.6	8.7	9.0	9.2
TOTAL SEPARATIONS (Thousands)	20.1	206.3	20.1	201.6	19.3	200.9

*Includes adult, child, and newborn care covered by SHSP; excludes services provided by Chronic and Convalescent Hospitals, and services provided to Indians and War Veterans' Allowance Recipients.

**Groups are titled and defined according to the International Classification of Diseases, 1955 Revision, Volume I, (Geneva: World Health Organization, 1957).

***S = Sample

TABLE 3.2 - Continued

DISTRIBUTION OF INPATIENT GENERAL HOSPITAL CARE*
BY DIAGNOSIS (ICD "18 BREAK")

DIAGNOSTIC GROUP**	1969		1970		1971	
	S*** %	Pop. %	S*** %	Pop. %	S*** %	Pop. %
Infective & Parasitic Diseases	3.5	4.2	3.6	4.2	3.4	4.2
Neoplasms	4.3	4.1	4.5	4.2	4.1	4.1
Endocrine, Nutritional & Metabolic Diseases	2.0	2.0	2.1	2.0	2.0	2.1
Diseases of the Blood & Blood-Forming Organs	0.9	0.9	0.8	0.9	0.9	0.9
Mental Disorders	2.5	2.7	3.2	3.0	3.2	3.1
Diseases of the Nervous System & Sense Organs	3.5	3.4	3.6	3.5	4.0	3.6
Diseases of the Circulatory System	9.4	8.8	9.2	8.6	9.4	9.0
Diseases of the Respiratory System	16.7	17.8	16.8	17.7	15.7	17.3
Diseases of the Digestive System	11.3	10.7	10.9	10.8	10.7	10.7
Diseases of the Genito-Urinary System	8.3	8.0	8.5	8.2	8.3	7.9
Complications of Pregnancy, Childbirth & the Puerperium	11.1	11.1	10.2	10.4	10.6	10.2
Diseases of the Skin & Subcutaneous Tissue	2.1	2.1	2.1	2.1	1.7	2.0
Diseases of the Musculoskeletal System & Connective Tissue	3.3	3.0	3.4	3.2	3.7	3.3
Congenital Anomalies	0.9	0.8	1.0	0.8	1.0	0.8
Certain Causes of Perinatal Morbidity & Mortality	0.7	0.6	0.6	0.6	0.9	0.7
Symptoms & Ill-Defined Conditions	3.0	3.0	3.6	3.5	3.8	3.6
Accidents, Poisonings & Violence	8.9	8.9	9.0	8.9	9.2	9.1
Supplementary Classifications	7.6	7.9	7.0	7.5	7.4	7.3
TOTAL SEPARATIONS (Thousands)	20.5	200.8	20.7	204.9	20.8	208.7

*Includes adult, child, and newborn care covered by SHSP; excludes services provided by Rehabilitation and Extended Care Hospitals, and services provided to Indians and War Veterans' Allowance Recipients.

**Groups are titled and defined according to the Eighth Revision International Classification of Diseases, Adapted for Use in the United States, Volume 1 (Washington: U.S. Department of Health, Education, and Welfare, Public Health Service Publication No. 1693, n.d.).

***S = Sample.

data. This is especially comforting in view of the unique procedure employed in their collection. Unlike the inpatient records, the income tax returns, from which the income data originate, do not contain the individual's SHSP registration-beneficiary number, so it was necessary to search the provincial tax files for the sample records using other appropriate, but less precise identifiers. Specifically, the search was based on the name, age, and address information drawn from the family and beneficiary records in the SHSP Master Registration File. Executed under strict official supervision so as not to violate the statutes governing access to these data, this procedure is fully described in another study.¹⁴ Here it suffices to note that the distributions shown in Table 3.3 are: (1) based on taxable returns for individuals and not for families; and (2) in the case of the "population" data, the Department of National Revenue's estimates based on a random sample of provincial income tax returns. Chi-Square tests confirm there are no significant differences between the sample and "population" distributions.¹⁵ Thus it may be concluded that the search and link procedure was successfully executed, and that the resultant sample income data are indeed representative.

Towards An Estimable Model of Hospital Use:

The Family Summary Record.

¹⁴R.G. Beck, op.cit., Appendix A.

¹⁵The calculated values of Chi square for the years 1966-71 are 1.02, 0.34, 0.40, 0.40, 0.41, and 1.48 respectively. For the appropriate degrees of freedom and the 5 per cent level of significance, the critical value of Chi square is 28.9.

TABLE 3.3

DISTRIBUTION OF TOTAL INCOME BY INCOME GROUP

Income Class	1966		1967		1968	
	Sample	Pop.	Sample	Pop.	Sample	Pop.
	%	%	%	%	%	%
\$ 0- 1,000	.12	.57	.19	.44	.19	.34
1,000- 1,500	5.84	5.10	4.95	4.78	5.06	4.93
1,500- 2,000	8.47	7.11	7.17	6.65	7.29	6.89
2,000- 2,500	8.85	8.31	7.64	7.70	8.09	7.50
2,500- 3,000	9.28	8.75	8.37	8.69	8.03	7.72
3,000- 3,500	9.65	9.45	8.94	8.53	8.62	8.79
3,500- 4,000	8.16	8.39	7.90	8.05	7.94	7.75
4,000- 4,500	7.62	8.21	7.50	7.44	7.16	7.06
4,500- 5,000	7.33	7.31	7.25	6.86	7.01	6.36
5,000- 5,500	6.13	6.61	6.36	6.21	6.12	6.38
5,500- 6,000	5.45	5.50	5.89	5.69	5.26	5.59
6,000- 7,000	7.70	8.23	9.32	9.39	9.19	9.82
7,000- 8,000	5.03	5.23	5.49	5.83	6.35	6.27
8,000- 9,000	3.14	3.39	4.00	4.02	4.11	4.43
9,000-10,000	2.14	2.23	2.61	2.70	2.70	2.77
10,000-15,000	3.45	3.99	4.52	4.97	5.06	5.38
15,000-20,000	.92	.94	1.06	1.19	.91	1.15
20,000-25,000	.31	.33	.38	.40	.47	.41
over 25,000	.41	.38	.45	.45	.45	.46
Total Number	25,702	253,011	26,318	268,511	26,762	269,725

Source: Department of National Revenue, "All Returns by Province and Income," Taxation Statistics, 1966, 1967, 1968, 1969, 1970, 1971.

TABLE 3.3 - Continued

DISTRIBUTION OF TOTAL INCOME BY INCOME GROUP

Income Class	1969		1970		1971	
	Sample %	Pop. %	Sample %	Pop. %	Sample %	Pop. %
\$ 0- 1,000	.26	.55	.35	.34	.17	.14
1,000- 1,500	4.91	4.47	5.21	4.93	4.80	.15
1,500- 2,000	7.53	7.52	7.50	6.79	7.44	4.25
2,000- 2,500	8.02	7.40	7.41	7.20	7.52	6.22
2,500- 3,000	8.19	7.95	8.23	7.54	7.37	6.90
3,000- 3,500	8.37	8.15	7.97	8.12	7.78	7.53
3,500- 4,000	7.58	7.51	7.87	7.77	7.40	7.63
4,000- 4,500	7.04	6.72	6.61	6.24	6.64	7.13
4,500- 5,000	6.11	6.04	5.75	5.88	5.69	6.12
5,000- 5,500	5.82	5.87	5.16	5.37	5.02	5.80
5,500- 6,000	5.25	5.65	5.04	5.06	4.75	5.33
6,000- 7,000	8.84	9.48	8.76	8.61	8.50	10.06
7,000- 8,000	6.92	6.99	6.98	7.47	7.03	8.76
8,000- 9,000	4.71	4.56	5.27	5.28	5.30	6.23
9,000-10,000	2.94	3.09	3.38	3.81	4.08	4.69
10,000-15,000	5.61	6.05	6.37	7.32	7.94	9.93
15,000-20,000	.98	1.07	1.15	1.24	1.48	1.89
20,000-25,000	.39	.42	.40	.44	.48	.54
over 25,000	.51	.52	.59	.57	.61	.69
TOTAL NUMBER	25,793	256,840	25,178	256,749	27,942	245,853

Having discussed the sampling methodology and the representativeness of the sample data, it is now appropriate to explain the form in which these data have been assembled and organized for subsequent analytical purposes. Essentially this entails describing the "family summary record" constructed for each of the sample families in each of the annual sub-samples. This record derives from the year specific registration, hospitalization, and income records collected for each family member, and, as its name suggests, it is a family level summarization of the pertinent data extracted from these records. Thus it includes information on all of the variables judged relevant to modelling the effect of copayment on families' access to and use of inpatient facilities.

Details of the model in its estimable form, and the theoretical considerations affecting the choice and definition of particular variables are discussed in the next section. Suffice it to say at this point that the family summary record reflects two quite deliberate decisions regarding the appropriate modelling strategy. These are: (1) the decision to employ the family, rather than the individual family member as the elementary unit of observation and analysis; and (2) the decision to employ a disaggregation scheme consistent with both the essential richness of the inpatient data, and the obvious heterogeneity of the hospitalization events described by these data. Neither of these decisions was made in haste. The former was the most difficult one to make, since valid arguments are easily adduced in favour of analyzing hospital use at the individual family member level. Intra-family aggregation was, however, the strategy judged best suited to the

specific policy-analytic orientation of this study. It permits appraisal of the copayment experiment in terms identical to those under which the experiment was officially monitored, namely, the SHSP "family unit" (i.e., a family head, his spouse, and non-taxable dependents). Moreover, as Andersen points out, "the family seems particularly appropriate for analysis of health service utilization because of changing needs for medical care through the family life cycle and because the family as a unit often decides what medical care a family member will receive."¹⁶ The rationale for the second decision should be more obvious. It stems from the conviction, acquired after careful appraisal of previous attempts to quantify price effects in hospital use,¹⁷ that thusfar there has been insufficient allowance for the possibility that user charges may ultimately affect only a subset of admission and/or discharge decisions. In these conceptual circumstances, disaggregation of the inpatient data is clearly indicated.

The reader is now invited to peruse Table 3.4 wherein a list of the summary record variables and their associated definitions is shown. Notice first that the variables are classified into two general categories, those that describe the family's socio-economic and demographic characteristics, and those that describe the family's

¹⁶Ronald Andersen, A Behavioral Model of Families' Use of Health Services (Chicago: Center for Health Administration Studies, Research Series, No. 25, University of Chicago, 1968), p.5. For a more extended discussion of the role of the family in health and illness see Theodor J. Litman, "The Family as a Basic Unit in Health and Medical Care: A Socio-Behavioral Overview," Social Science and Medicine, Vol.8, Nos.9 and 10, September 1974, pp. 495-519.

¹⁷See Chapter 2 for a review of this literature.

TABLE 3.4

LIST OF VARIABLES IN FAMILY SUMMARY RECORD

I SOCIO-ECONOMIC and DEMOGRAPHIC VARIABLES

SINGLE MALE	Equals 1 if the family head is a single male; otherwise zero.
SINGLE FEMALE	Equals 1 if the family head is a single female; otherwise zero.
C1	Equals 1 if the family has one child; otherwise zero.
C2	Equals 1 if the family has two children; otherwise zero.
C3	Equals 1 if the family has three children; otherwise zero.
C4	Equals 1 if the family has four children; otherwise zero.
C5+	Equals 1 if the family has five or more children; otherwise zero.
A1	Equals 1 if the age of the family head is 25-34 years; otherwise zero.
A2	Equals 1 if the age of the family head is 35-44 years; otherwise zero.
A3	Equals 1 if the age of the family head is 45-54 years; otherwise zero.
A4	Equals 1 if the age of the family head is 55-64 years; otherwise zero.
A5	Equals 1 if the age of the family head is 65-74 years; otherwise zero.
A6	Equals 1 if the age of the family head is 75 years or more; otherwise zero.
URBAN	Equals 1 if the family lives in an urban centre with a population of 10,000 or more; otherwise zero.
WELFARE	Equals 1 if the family receives social assistance payments; otherwise zero.
SIZE	The number of adults and children (under the age of eighteen) in the family.
AGE	The age of the family head (in years).
INCOME	Total family income before taxes.

II UTILIZATION VARIABLES

DAYDTX	The number of inpatient days for <u>all separations</u> (discharges, D; transfer, T; and deaths, X) from Saskatchewan general hospitals experienced by all family members except newborn.
ADMDTX	The number of admissions represented in DAYDTX above.
BINDTX	Equals 1 if DAYDTX is positive; otherwise zero.
DAYDIS	The number of inpatient days for <u>all discharges</u> (DIS) from Saskatchewan general hospitals experienced by all family members except newborn.
ADMDIS	The number of admissions represented in DAYDIS above.
BINDIS	Equals 1 if DAYDIS is positive; otherwise zero.
DAYTYPE _i	The number of inpatient days for all discharges in the <u>ith diagnostic category</u> from Saskatchewan general hospitals experienced by all family members except newborn.
ADMTYPE _i	The number of admissions represented in DAYTYPE _i above.
BINTYPE _i	Equals 1 if DAYTYPE _i is positive; otherwise zero.
DAYMED	The number of non-obstetrical <u>medical</u> (MED) inpatient days for all discharges from Saskatchewan general hospitals experienced by all family members except newborn.
ADMED	The number of admissions represented in DAYMED above.
BINMED	Equals 1 if DAYMED is positive; otherwise zero.
DAYSUR	The number of non-obstetrical <u>surgical</u> (SUR) inpatient days for all discharges from Saskatchewan general hospitals experienced by all family members except newborn.
ADMSUR	The number of admissions represented in DAYSUR above.
BINSUR	Equals 1 if DAYSUR is positive; otherwise zero.
DAYSGP	The number of non-obstetrical inpatient days for all discharges from Saskatchewan general hospitals attended by <u>solo general practitioners</u> (SGP) experienced by all family members except newborn.
ADMSGP	The number of admissions represented in DAYSGP above.
BINSGP	Equals 1 if DAYSGP is positive; otherwise zero.

DAYGGP	The number of non-obstetrical inpatient days for all discharges from Saskatchewan general hospitals attended by <u>group general practitioners</u> (GGP) experienced by all family members except newborn.
ADMGGP	The number of admissions represented in DAYGGP above.
BINGGP	Equals 1 if DAYGGP is positive; otherwise zero.
DAYMSP	The number of non-obstetrical inpatient days for all discharges from Saskatchewan general hospitals attended by <u>practitioners of medical or surgical specialties</u> (MSP) experienced by all family members except newborn.
ADMMSp	The number of admissions represented in DAYMSP above.
BINMSP	Equals 1 if DAYMSP is positive; otherwise zero.
DAYCOM	The number of non-obstetrical inpatient days for all discharges from Saskatchewan <u>Community</u> (COM) general hospitals experienced by all family members except newborn.
ADMCOM	The number of admissions represented in DAYCOM above.
BINCOM	Equals 1 if DAYCOM is positive; otherwise zero.
DAYREG	The number of non-obstetrical inpatient days for all discharges from Saskatchewan <u>Regional</u> (REG) general hospitals experienced by all family members except newborn.
ADMREG	The number of admissions represented in DAYREG above.
BINREG	Equals 1 if DAYREG is positive; otherwise zero.
DAYBAS	The number of non-obstetrical inpatient days for all discharges from Saskatchewan <u>Base</u> (BAS) general hospitals experienced by all family members except newborn.
ADMBAS	The number of admissions represented in DAYBAS above.
BINBAS	Equals 1 if DAYBAS is positive; otherwise zero.

III TIME-COPAYMENT VARIABLES

D67	Equals 1 if summary record year = 1967; otherwise zero.
D68	Equals 1 if summary record year = 1968; otherwise zero.
D69	Equals 1 if summary record year = 1969; otherwise zero.
D70	Equals 1 if summary record year = 1970; otherwise zero.
D71	Equals 1 if summary record year = 1971; otherwise zero.

hospitalization experiences. (A third group of variables -- the "time-copayment" variables -- identify the annual sub-sample to which the family in question belongs and to which the summary record relates.) Included in the first group are descriptors of the age, sex, and marital status of the family head, number of children, family residence, family income, and social assistance status. All these variables, with the exception of family income, have been compiled from information contained in the registration records drawn from the Master Registration File; the income variable was, as previously mentioned, compiled from the tax returns filed by individual family members.

Notice also that all but three of these variables (SIZE, AGE, and INCOME) are defined in dichotomous, categorical terms. Marital status is divided into SINGLE MALE and SINGLE FEMALE, a scheme under which both variables are "zeroed-out" for a family in which both adult spouses are present. Age of head is divided into seven groups: under 25 years (the "omitted" age group for purposes of this dummy variable scheme), 25-34 (A1), 35-44 (A2), 45-54 (A3), 55-64 (A4), 65-74 (A5), and 75 and over (A6).. Family size is divided into six groups: no children (the "omitted" category), one child (C1), two children (C2), three children (C3), four children (C4), and five or more children (C5+). (For methodological reasons, to be later explained, both age of head and family size are alternatively defined by the continuous variables, AGE and SIZE respectively.) Finally, family residence and social assistance status are respectively defined by the dummy variables, URBAN and WELFARE. These variables "turn-on" (i.e., equal 1) only when the family being considered resides in an urban area and/or receives social assistance payments. A feature of the overall dummy variable scheme is

that the marital status, age, size, residence, and welfare variables all assume zero values for a rural couple headed by a male under 25 years, having no children, and receiving no social assistance payments. In subsequent analysis, this category of family will be known as the "reference" family.

The second group of variables describe the family's hospitalization experiences. Compiled from the sample records drawn from the SHSP Hospital Claims File, these variables invite a number of general comments. First, like those in the preceding group, the time period over which observations on these variables are defined is one year. Second, each variable has been compiled by aggregating the relevant data over all family members except newborn. Contrary to what might be assumed, this procedure does not materially reduce the completeness of the utilization history, since the pertinent use data relating to the birth of new family members are contained in the mother's admission-separation record.¹⁸ Third, they refer only to inpatient care provided in Saskatchewan general hospitals and insured by SHSP. The resultant exclusion of a very small number of SHSP insured

¹⁸While her occupancy of a bed in the obstetrical ward and her infant's occupancy of a bassinet in the newborn nursery are clearly two discrete hospitalization events, and are routinely recorded as such, the latter is conceptually unimportant to the present study, since newborn care was not subject to the per diem utilization fee during the copayment period. Curiously, the feminists apparently overlooked the implied opportunity to argue for the exemption of "confinements" (a medical term to which they would surely relate) from copayment on the grounds that it merely added "economic insult to biological injury!"

out-of-province hospitalizations¹⁹ experienced by members of the sample families can be justified on purely pragmatic grounds: the admission-separation records documenting these events do not in general provide sufficient information to classify the attending physician and/or surgeon by mode of practice (solo general practitioner, group general practitioner, medical or surgical specialist), or the host institution by size and service capability (community, regional, or base hospital). The machine-readable doctor and hospital "dictionaries" via which these determinations have been made in this study include only in-province practitioners and in-province general hospitals.²⁰ Third, each variable relates to a particular definition of the hospitalization event. Some "broadly" define the event as admission to a general hospital regardless of the (adult or child) patient's illness or condition, the treatment provided, the admitting physician's practice mode and/or specialty, or the host institution's size and service capability; others "narrowly" define the event in terms of these various case-specific characteristics.

Included in the sub-group of "broadly defined use" variables are DAYDTX, ADMDTX, BINDTX, DAYDIS, ADMDIS, and BINDIS. Of these, the first three define the event as one involving admission to and subsequent separation from hospital, where the term separation definitionally includes any one of medical discharge, transfer to another institution,

¹⁹Specifically, out-of-province claims represent only 2.5 per cent (or less) of total sample claims in each of the study years, 1966-71.

²⁰These two dictionaries were constructed specifically for this study from information contained in MCIC's historical register of Saskatchewan physicians and surgeons and SHSP's historical register of Saskatchewan general hospital respectively.

or death (with or without autopsy). DAYDTX and ADMDTX are continuous variables measuring the family's total number of days and separations respectively; BINDTX is a binary or dichotomous variable which assigns the family "user" or "non-user" status depending on whether or not any of its members experienced hospitalization as defined by DAYDTX (ADMDTX). The last three broadly defined use variables define the event as one involving admission to and subsequent discharge from hospital, where the term discharge definitionally excludes transfer to another institution and death. DAYDIS and ADMDIS are continuous variables measuring the family's total number of days and discharges respectively; BINDIS is a binary or dichotomous variable which assigns the family "user" or "non-user" status depending on whether or not any of its members experienced hospitalization as defined by DAYDIS (ADMDIS). Notice that this strategy of constructing two continuous variables and one dichotomous variable to alternatively measure use (non-use) is common to each and every one of the "broad" and "narrow" definitions of hospitalization.

Included in the sub-group of narrowly defined use variables are:

- (1) the diagnostic-specific measures, DAYTYPE_i, ADMTYPE_i, and BINTYPE_i;
 - (2) the treatment-specific measures, DAYMED, ADMMED, BINMED, DAYSUR, ADMSUR, and BINSUR;
 - (3) the practitioner-specific measures, DAYSGP, ADMSGP, BINS GP, DAYGGP, ADMGGP, BINGGP, DAYMSP, ADM MSP, and BINMSP; and
 - (4) the institution-specific measures DAYCOM, ADMCOM, BINCOM, DAYREG, ADMREG, BINREG, DAYBAS, ADMBAS, and BINBAS.
- The diagnostic categories associated with the variables DAYTYPE_i, ADMTYPE_i, and BINTYPE_i are those colloquially referred to as the "18 break" of the International Classification of Diseases, 7th Revision (1955). These categories are:

(1) infective and parasitic diseases; (2) neoplasms; (3) allergic, endocrine system, metabolic and nutritional diseases; (4) diseases of the blood and blood-forming organs; (5) mental, psychoneurotic, and personality disorders; (6) diseases of the circulatory system; (8) diseases of the respiratory system; (9) diseases of the digestive system; (10) diseases of the genito-urinary system; (11) deliveries and complications of pregnancy, childbirth, and the puerperium; (12) diseases of the skin and cellular tissue; (13) diseases of the bones and organs of movement; (14) congenital malformations; (15) certain diseases of early infancy; (16) symptoms, senility, and ill-defined conditions; (17) accidents, poisonings, and violence; and (18) supplementary classifications for special admissions, livebirths, and stillbirths. It should be noted that these categories have been adjusted for a major revision in the International Classification of Diseases which would have otherwise affected their composition, and hence their comparability between 1966-68 and 1969-71.²¹

²¹Compilation of the three use measures for each of these diagnostic categories was easier for the 1966-68 summary records than for the 1969-71 summary records. In the former years, when SHSP coded diagnoses according to the I.S.C., 7th Revision (1955), it was simply a matter of summing the days and admissions over all claims falling within each of the eighteen diagnostic categories. Allocation of the claims to each category was based on the principal discharge diagnosis (a four digit I.S.C., 7th Revision code) and its corresponding "18 break" code. In the latter years, when SHSP coded diagnoses according to the I.S.C., 8th Revision (1965), it was first necessary to translate the reported ICDA-8 codes (four digits) for principal discharge diagnosis into their equivalent I.S.C., 7th Revision codes (four digits), since, as previously mentioned (See page 86, fn. 12), the 7th and 8th Revisions of the ICD codebook involve different definitions of the "18 break." To ensure a consistent allocation of claims by diagnosis, and hence a consistent definition of the diagnostic-specific use variables across all the annual sub-samples, a uniform definition of the "18 break" had to be employed. The 7th Revision definition of the "18 break" was chosen due to the relative ease of the requisite recode operation. The 8th Revision contains a more detailed tabular list of (four digit)

3 Definitions of the other "narrowly defined" use variables are, for the most part, self-explanatory. The treatment-specific measures (DAYMED through BINSUR in Table 3.4) distinguish hospitalizations for non-obstetrical, medical reasons from hospitalizations for non-obstetrical surgical reasons. The variables DAYMED, ADMED, and BINMED were compiled from claims where the principal discharge diagnosis was non-obstetrical (i.e., where the I.S.C. code was not among those included in category [11], "deliveries and complications of pregnancy, childbirth, and the puerperium"), and where there was no indication that a surgical procedure had been performed. Correspondingly, the variables DAYSUR, ADMSUR, and BINSUR were compiled from claims containing a non-obstetrical primary discharge diagnosis, and an indication that the patient had undergone some surgical procedure during his or her hospital stay. As mentioned above, compilation of the practitioner-specific measures (DAYSGP through BINMSP in Table 3.4) entailed sorting the claims on the basis of information contained in the specially constructed "doctor dictionaries." These dictionaries - one for each of the study years 1966-71 - were the devices by which the numeric code of the admitting physician identified on each claim was translated into a mode of practice and/or specialty indicator. The dictionary definitions of solo general practice, group general practice, and specialty practice are identical to the official M.C.I.C. definitions. Thus, solo general practice is a setting in which the general practitioner does not work in association with another physician; group general practice is a setting

✓ illnesses and conditions than does the 7th Revision, and, in consequence, many I.S.C.-7 codes do not have unambiguous counterparts in the ICDA-8 codebook. Hence the rationale for the particular recode procedure adopted in this study.

in which two or more general practitioners work in association by virtue of their having access to one another's medical records of patient encounters; specialty practice is a setting restricted to those physicians listed by the College of Physicians and Surgeons of Saskatchewan as eligible to receive M.C.I.C. payments for insured services at specialist rates.²² Finally, the institution-specific measures (DAYCOM through BINBAS in Table 3.4), are based on official definitions of service capability. Thus, according to SHSP: a "community" hospital is a general hospital that serves the population of the local area by providing a smaller range of services than a regional hospital; a "regional" hospital is a general hospital which serves as a district or community hospital for the local population, but because it provides a wider range of medical and surgical services than the limited function community hospital, it also serves as a first referral center for persons outside its local area; a "base" hospital is a general hospital which serves as a district or community hospital for the local population, but because it provides a wider range of services than a regional hospital, it serves as an ultimate referral center, and, as well, performs a major role in medical research and education.²³ For a more detailed description of the clinical and other services provided by the representative institution in the community, regional and base

²²Medical Care Insurance Commission, Annual Report 1971 (Regina: Queen's Printer, 1972), p.21. It should be noted that for the very small number of physicians who experienced a change in practice setting and/or specialty during any given study year, dictionary assignments were based on the setting and/or specialty in which the physician practised for seven or more months of the year.

²³Saskatchewan Hospital Services Plan, Annual Report 1971 (Regina: Queen's Printer, 1972), p.1.

categories, along with the pertinent data on the distribution of Saskatchewan hospitals by size and service capability, see Appendix B.

In summary, the sampling procedures described in Section 2, together with the assembly strategy described in this section, yield a basic data matrix for each of the six study years 1966-71. Each annual data matrix contains summary records for as many families as are indicated in Figure 3.2, and hence approximately 38,000 observations on all of the variables named and defined in Table 3.4. It is this set of six annual data matrices upon which all subsequent analyses are based. Details of the conceptual and statistical model to be confronted with these data are discussed in the next section.

Formulation of the Model

The model via which the effects of copayment on families' use of hospital services are analyzed in this study essentially consists of three basic equations. Specified in their most general form they are:

$$(1) \text{Pr}(H)_{it} = f(P_{it}, Z_{it})$$

$$(2) \text{DAY}_{it} = g(P_{it}, Z_{it})$$

$$(3) \text{ADM}_{it} = h(P_{it}, Z_{it})$$

where, for the i^{th} family in year t :

$\text{Pr}(H)$ = the probability of hospitalization.

DAY = the number of patient-days, given hospitalization occurs.

ADM = the number of admissions, given hospitalization occurs.

P = the "point-of-service price" of inpatient care.

Z = a vector of variables describing the family's socio-economic and demographic status.

While this statement of the model is intended primarily to orient the reader to subsequent discussion, it is well at this early juncture to briefly comment upon several of its more prominent attributes. First, it is micro-analytic in that it treats the family as the elemental unit of observation. As mentioned in the previous section, this strategy permits price effects to be searched for at a policy-relevant level and, as well, facilitates conceptually important distinctions between different categories of program beneficiaries. Equally important, as will be shortly shown, it does not preclude the formulation and testing of hypotheses concerning the effects of physician-related and hospital-related factors on families' use of inpatient facilities. Second, each equation involves a unique concept and measure of hospital use, and thereby, a unique implicit phrasing of the central price-theoretic question. Equation (1) focuses upon the probability of hospitalization, and addresses the question "is this probability reduced when the user price is increased?"; equation (2) focuses upon the volume of hospitalization among "user" families, and addresses the question "is the annual number of patient-days per family reduced when the user price is increased?"; equation (3) focuses upon the frequency of hospitalization among "user" families, and addresses the question "is the annual number of admissions per family reduced when the user price is increased?" Moreover, and equally important, equations (2) and (3) together define a fourth concept and measure of use, namely, the duration of stay per admission, and hence they together address the question "is the average stay per admission per family reduced when the user price is increased?" Third, the model

is multi-variate in that each question is specified to include not only the "user price" of inpatient care, but also a vector of socio-demographic variables. These latter variables are inserted to test a number of hypotheses concerning inter-family variation in hospitalization experiences, and, more important, to test the specific hypothesis that price effects vary in magnitude across different types and/or classes of families. Finally, the model is practical in that it respects the limitations of the available Saskatchewan data; it includes only those a priori important variables for which operational measures can be generated.

Having said this much, other more specific properties of the model and the modelling strategy invite discussion. In order of presentation the topics are: (a) the dependent variables and the logic of the disaggregation scheme they are designed to accommodate; (b) the independent variables and their hypothesized relationships to the dependent variables; and (c) the estimating forms of the three basic equations, and the methods of empirical implementation.

(a) The Dependent Variables and the Disaggregation Scheme

As noted above, the dependent variables in the model are the probability of hospitalization, the volume of hospitalization, and the frequency of hospitalization. Taken together these variables capture three salient dimensions of inter-family variations in hospital use and, as such, are self-evidently complementary to the micro-analytic character of the research design. Equally obvious, however, is the fact that in their present definitional states, they fail to convey any information on the nature of the hospitalization events to which they

refer. More specifically, they do not bear any direct correspondence to the wide variety of utilization measures contained in the family summary record (see Table 3.4). Thus the immediate task is to clarify the logic of the disaggregation scheme used to generate the complete list of utilization variables shown in the summary record.

The general rationale for the scheme is very simple. It is to avoid the assumption, implicit or explicit in the majority of price-use studies reviewed in Chapter 2, that all admission and discharge decisions are equally (non)discretionary, and hence, equally (in)sensitive to point-of-service charges on inpatients. This assumption invites relaxation not only because it contradicts the reality of the decision-making process, but also because it needlessly restricts the range of testable hypotheses on price effects to those featuring an omnibus definition of hospital use and/or the hospitalization event. Evidence from two studies outside the price effects literature clearly indicates that the reality involves substantial variation in the "necessariness" of hospitalization. Thus, Anderson and Sheatsley found that only 70 per cent of some 2000 non-obstetrical admissions to Massachusetts hospitals in 1960-61 were "absolutely necessary" in the opinion of the physicians who made these decisions,²⁴ while Fitzpatrick et al. found that "extra-medical"

²⁴Odin W. Anderson and Paul B. Sheatsley, Hospital Use: A Survey of Patient and Physician Decisions (Chicago: Center for Health Administration Studies, Research Series No. 24, University of Chicago Press, 1967), p.85. Seventy per cent is almost certainly an overestimate of the true incidence of "absolute necessary" hospitalization, since physicians are educated or otherwise wont to justify their own decisions respecting the use of expensive hospital facilities in terms of some medico-technical imperative, else they reinforce the widespread belief that costs can be reduced by eliminating "unnecessary" admissions and "over-long" stays.

factors not normally associated with "absolutely necessary" hospitalization were reported by Michigan physicians as having affected their decisions in over 19 per cent of the (1751) cases selected for review.²⁵ Moreover, as Anderson and Sheatsley rightly emphasized, the concept of necessary hospitalization is itself "very flexible": physicians' decisions concerning admission and length of stay are generally affected by a variable combination of medical and extra-medical factors depending on time, place, and circumstance, with the result that hospital use judged necessary in one context might well be judged unnecessary in another.²⁶ This strongly suggests that the search for price effects ought not to be constrained by an assumption which, in effect, suppresses the contextual heterogeneity of hospitalization events, and thereby, denies any possibility of identifying these effects in those subsets of events where the requisite element of choice is arguably greatest.

²⁵Thomas B. Fitzpatrick, Donald C. Riedel, and Beverly C. Payne, "Character and Effectiveness of Hospital Use" in W.J. McNerney et al. (eds.) Hospital and Medical Economics, Vol. 1 (Chicago: Hospital Research and Educational Trust, 1962). The authors emphasize that the use of the term 'extra-medical' in no sense

"indicates that the factor does not adequately explain the hospital use in question, nor is it equated with any judgment that the use was To divide all factors influencing use of the hospital into 'justified' and 'not justified' and further to class all 'justified factors as medical,' as is the tendency, would completely obscure the causative factors relating to use of the hospital." Ibid., p.468.

Among the extra-medical factors identified in the study as affecting admission and/or discharge decisions were: the availability of care outside the hospital, financial deterrents, financial incentives, convenience to physicians, and convenience to patients. For a detailed classification of the extra-medical factors and related discussion of the results from the physician interview, see Ibid., pp. 496-509.

²⁶Anderson and Sheatsley, op. cit., pp. 112-113.

Progress in this intuitively appealing direction is impeded however by one major problem. It is that the degree of discretion in any given utilization decision is, as implied above, an item of information which defies direct observation, and hence, direct measurement. Thus it is impossible to directly test the key hypothesis that price effects are greater (smaller), the more (less) discretionary the admission and discharge decisions. Nor is it possible to directly test a variant of this hypothesis suggested by Berki. He speculates that

When a condition in both the patient's and physician's estimation is emergent in the sense of being traumatic or life-threatening, the physician's medical decision is strictly condition (or patient) focused and hence the economic algorithm is not applicable,²⁷

i.e., there is no price (or income) effect. The price effect can occur only "when in the physician's estimation the disequilibrium (in the patient's health status) is not emergent."²⁸ Again the problem is that the emergent vs. non-emergent character of the condition is not rigorously identified in most data-sets, least of all in claims records routinely compiled by public and private insurance agencies. Evidently then, data limitations effectively rule out any chance of implementing what might be termed the "ideal" disaggregation scheme, one incorporating explicit and unambiguous distinctions between discretionary and non-discretionary utilization decisions.

The obvious alternative is to adopt a "second-best" scheme based on certain reasonable assumptions about the distribution of discretion

²⁷Sylvester E. Berki, Hospital Economics (Lexington: D.C. Heath and Co., 1972), p. 132.

²⁸ Ibid.

across well-defined hospitalization events. Indeed, in view of the discussion in Section 3, it should not surprise the reader to learn that it is this approach on which the present study places major reliance. Recall that the set of utilization variables contained in the family summary record is dominated by what were there referred to as "narrowly-defined" use measures. Unlike those which are "broadly-defined," these measures explicitly preserve selected factual aspects of the event, including the diagnostic and treatment status of the patient, the practice setting and specialty of the admitting physician, and the size, location, and service capability of the host active-treatment institution. Since these are the primary vehicles via which the search for differential price effects will be conducted, the arguments germane to their selection merit careful statements and consideration.

Consider first the diagnostically defined measures $DAYTYPE_i$, $ADMTYPE_i$, and $BINTYPE_i$. Two basic assumptions underlying the adoption of these measures are: (i) that the discretionary component of the physician's decision to recommend admission and subsequent discharge depends in part upon his diagnosis of the patient's illness or condition; and (ii) that some of the eighteen major diagnostic groups contained in the International Classification of Diseases (ICD) can be proximately classified as either "relatively discretionary" or "relatively non-discretionary." The first assumption requires only brief comment, inasmuch as it corresponds to one obvious fact of the clinical reality, namely, that indications for admission and discharge vary widely depending upon the nature of the patient's health problem. For some signs, symptoms, illnesses or conditions, the admission and discharge decisions are dominated by fairly strict medico-technical

criteria, and hence are quite unequivocal. The patient suspected of having an acute myocardial infarction, acute appendicitis, acute peritonitis, or other potentially life-threatening disorder would, under prevailing standards of "good" medical practice, place the physician in this "little or no-choice" situation. For other, less serious problems, however, such as hernia of the abdominal cavity or hypertrophy of the tonsils and adenoids, indications are more equivocal, and, in consequence, the timing of hospitalization decisions might well be influenced by extra-medical criteria, including convenience to the patient and/or the physician. This reasoning is, of course, very similar to that employed by Berki, the only refinement being that his conceptual distinction between "emergent" and "non-emergent" conditions is specifically expressed in the diagnostic dimension.²⁹

The second assumption is much stronger than the first, since the eighteen major diagnostic groups of the ICD are primarily arranged according to principal anatomical site (i.e., diseases of the digestive system, circulatory system, genito-urinary system, etc.), and are not designed to support inferences about inter-group differences with respect to the degree of technical determinism in the admission and discharge decisions. Nevertheless, it is possible to single out certain groups whose component illnesses and conditions can be presumed to present the physician with the least opportunity for discretionary decision-making, and, by extension, the least opportunity to consciously respond to an extra-medical factor like the user-cost of inpatient care.

²⁹It must be emphasized that diagnosis is here assumed to be only a first approximation of the patient's location on the spectrum of case severity. More precise determinations of severity would require detailed information from medical records not available to this study.

Thus, "Neoplasms" (Major Group II), "Deliveries and complications of pregnancy, childbirth and puerperium" (Major Group XI), and "Accidents, poisonings, and violence" (Major Group XVII), are three categories which may be classified as "relatively non-discretionary," since cancer patients (especially those with malignant neoplasms), expectant mothers, and accident victims are three general case categories where indications for admission and, to a lesser extent, discharge are dominated by fairly strict medico-technical criteria. It is therefore hypothesized that there will be no significant price effect when the basic equations in the model are alternately specified and estimated with these three diagnostically defined dependent variables.

Expectations in respect of the other categories are less easily formulated, given that they each contain a wide variety of specific diagnoses which vary markedly in severity and/or complexity. For example, diseases of the circulatory system encompass many different heart, vein, and artery diseases, some of which are "very serious" (e.g., acute myocardial infarction and cerebral hemorrhage), and some of which are not (e.g., hemorrhoids and varicose veins). The obvious alternative is to further disaggregate each of these categories using, for example, the detailed list of three-digit ICD rubrics. However, in the absence of a generally accepted a priori scheme for allocating each and every one of the more than 1000 discrete diagnoses, and equally important, in the presence of computer constraints on the number of variables that can be constructed and analyzed for each of the sample families, this strategy is not viable. Again, a "second-best" strategy is in order, and, in this particular instance, it is to select the one major diagnostic category whose component illnesses and conditions are

concentrated in the "relatively discretionary" category. The category which comes closest to satisfying this criterion is Major Diagnostic Group VIII, "Diseases of the respiratory system." It includes acute nasopharyngitis (common cold), acute tonsillitis, influenza, pneumonia, bronchitis, and others, the majority of which are identified in one clinical taxonomy as conditions "not generally requiring hospitalization."³⁰ Accordingly, it may be hypothesized that this category is the one most likely to show evidence of a significant price effect, though it is well at this point to emphasize that expectations of this sort rest on the additional assumption that the doctor-patient relationship is not totally dominated by provider preferences unrelated to the patient's economic welfare.

The second dimension along which the search for differential price effects will be conducted is complementary to the first in that it also relates to the nature of the patient's health problem. In particular, the "treatment specific" measures (DAYMED, ADMMED, BINMED, DAYSUR, ADMSUR, BINSUR) distinguish between patients who are admitted to hospital for medical treatment of non-obstetrical conditions and patients who are admitted to hospital for surgical treatment of non-obstetrical conditions.³¹ Evidence strongly supportive of this strategy is presented in the study by Anderson and Sheatsley. They found that

³⁰Arnold V. Hurtado and Merwyn R. Greenlick, "A Disease Classification System for Analysis of Medical Care Utilization, with a Note on Symptom Classification," Health Services Research Vol. 6, (Fall 1971), pp. 235-250.

³¹Consistent with the preceding argument that obstetrical cases are "relatively non-discretionary," they are excluded from these and all other "narrowly defined" use measures to be subsequently discussed.

Physicians exercise much greater discretion in the hospitalization of medical patients than of surgical patients. It is noted that physicians felt that for 46 per cent of the medical patients it was impossible to treat them outside of the hospital (compared with 74 per cent of the surgical patients).³²

It follows, in theory at least, that price effects should be detectably greater among medical patients than among surgical patients.

Evidence from other studies also indirectly suggests that price effects may systematically vary in relation to certain characteristics of the admitting physician. Thus a number of investigators have shown, albeit with varying degrees of clarity, that salaried physicians in group practice (and particularly in group practices whose revenues are annually pre-determined through a capitation arrangement with subscribers) are less likely to hospitalize and less likely to do controversial surgical procedures (e.g., tonsillectomy, cholecystectomy, hysterectomy, herniorrhaphy) than fee-for-service physicians in solo practice.³³ Quite apart from the issue of whether this finding supports a quality of care argument for group practice, there is some question in these studies whether the salutary influence on the physician's hospitalization decisions is primarily economic or medical in origin.

³²Anderson and Sheatsley, op. cit., pp. 92-93.

³³The "classic" studies are: P.M. Densen, E. Balamuth, and S. Shapiro, Prepaid Medical Care and Hospital Utilization, Hospital Monograph Series No. 3, (Chicago: American Hospital Association, 1958); and P.M. Densen et al., "Prepaid Medical and Hospital Utilization in a Dual Choice Situation," American Journal of Public Health, Vol. 50, November 1960, pp. 1710-1726. See also G.T. Perrott and J.C. Chase, "The Federal Employees Health Benefits Program," Report of the National Advisory Commission on Health Manpower, Volume 1, Washington, D.C., 1967.

For evidence that education of both the patient and physician in the conservative use of hospital services can minimize differences in admission rates between prepaid group and fee-for-service solo practice, see P.M. Densen, E.W. Jones, and I. Baldinger, "Prepaid Medical Care and Hospital Utilization," Hospitals, Vol. 36, November 16, 1962.

If it is economic, that is, if it is the absence of fee incentives in the group practices, and their presence in solo practices which explains the hospitalization differential, then one would not expect differential price effects across practice settings in the present study for the simply reason that the solo/group dichotomy here employed relates only to the number of general practitioners per clinic (one vs. more than one), and not to any distinction in the method of physicians' remuneration. On the other hand, if the disparity in hospitalization patterns stems from differences in the organization and conduct of medical practice, that is, if professional controls (e.g., peer review, collegial interaction, and medical audit) normally associated with group practice, but not normally associated with solo practice account for the utilization differences, then the following relationship may be conjectured between copayment and practice setting in the Saskatchewan data: the probability, volume, frequency, and duration of hospitalization is likely to be more price sensitive among families attended by solo general practitioners than among families attended by group general practitioners. Underlying this hypothesis is the (untestable) assumption that in the absence of copayment, there is a concentration of "questionable" hospital use among solo GPs and hence greater scope (if not inducement) for them to become more conservative in their use of inpatient facilities when copayment is introduced. To test this hypothesis, the family summary record includes the "practitioner-specific" use measures, DAYSGP, ADMSGP, BINS GP (the solo GP variables) and DAYGGP, ADMGGP, BINGGP (the group GP variables). As well, it includes the variables DAYMSP, ADMMSP, BINMSP to test the related hypothesis that medical and surgical specialists located in the

province's urban areas are less likely to alter their practice styles with the introduction of copayment, since their patients are typically seen on a referred basis, and are typically those with the more serious and/or complex health problems.

The last group of narrowly defined use measures are "institution-specific." They differentiate between patients admitted to community hospitals (DAYCOM, ADMCOM, BINCOM), regional hospitals (DAYREG, ADMREG, BINREG), and base hospitals (DAYBAS, ADMBAS, BINBAS). In the Saskatchewan context these three types of active treatment institutions differ not only in terms of service capability (i.e., the range of diagnostic and therapeutic services offered) but also in terms of rated bed capacity, medical staff organization, and population served.³⁴ The community hospital is a small "low" technology institution (10-24 beds), staffed by one or two general practitioners, and populated by patients from the local town or village and adjacent rural municipalities. In sharp contrast, the regional and base hospitals tend to be much larger institutions (150-300 and 300-500 beds respectively) with formally organized medical staffs consisting of both general practitioners and specialists. These institutions provide primary services to the population of the immediate urban centers (population 10,000 or more) and adjacent rural municipalities in which they are located, and as well act as referral centres for the population of the region in which they are located. The regional hospitals, (located in

³⁴ Selected characteristics of these institutions are described in Appendix B. For further details see Saskatchewan Hospital Survey and Master Plan 1961, Part I (Regina: Department of Public Health, 1962); and Directory of General Hospital Facilities in Saskatchewan, (Regina: Saskatchewan Hospital Services Plan, December 31, 1971).

North Battleford, Prince Albert, Yorkton, Moose Jaw, and Swift Current) are "medium" technology institutions offering a basic range of specialty medical services, while the base hospitals (located in Regina and Saskatoon) are "high" technology institutions offering a wide range of specialty medical services.

From these scale and organizational distinctions, a number of "performance" propositions immediately follow, the most important of which for present purposes is the one associating the "low" technology community hospitals with a relatively liberal admission and discharge policy, and the "medium to high" technology regional and base hospitals with a relatively strict admission and discharge policy.³⁵ Anecdotal evidence on the relatively high incidence of "so-called" "social admissions" to small rural hospitals in Saskatchewan and elsewhere is clearly consistent with this reasoning. In any event, the hypothesis that price effects will be differentially greater among patients admitted to community hospitals than among patients admitted to regional and/or base hospitals is eminently reasonable, and accordingly, will be tested with the "institution-specific" dependent variables described above.³⁶

³⁵Another not unrelated proposition, which has received some support and attention in the literature, is that the quality of services provided in or by small community hospitals is below attainable standards. See Walter O. Spitzer, "The Small General Hospital: Problems and Solutions," Milbank Memorial Fund Quarterly, Vol. XLVIII, October 1970.

³⁶One notable departure from this strategy became necessary during the data analysis phase of the study. In particular, since core constraints in the computer dictated that the correlation matrices generated from the family summary records could not exceed 132 x 132, some difficult decisions had to be made, one of which was to suppress the distinction between regional and base hospitals. The dependent variables defined for the resultant institutional group are DAYREB,

This completes the discussion of the dependent variables and the logic of the disaggregation scheme they are designed to accommodate. While no contention is made that the scheme is an ideal one, it can be fairly claimed that it satisfactorily incorporates many of the key dimensions along which diverse price effects may theoretically occur. Thus, by successive substitution of the various "broad" and "narrow" definitions of hospital use into the basic three equation model, it will be possible to extend the search for price effects into areas which have hitherto received little or no serious attention.

(b) The Independent Variables and Associated Hypotheses

Thusfar, the independent variables in each of the three basic equations have been identified in general terms as the point-of-service price of inpatient care, and the vector of socio-demographic and economic variables. Further comment on these variables will focus on their theoretical status in the model and their operational counterparts in the family summary record. At the conclusion of this sub-section, the stage is set for a full viewing of the model in its various estimable forms and the associated methods of empirical implementation.

Present purposes are served by making repeat mention of the fact that the vast majority of previous attempts to estimate price effects in hospital use have encountered problems in operationally defining the relevant net price variable, i.e., the "out-of-pocket" price per unit actually paid by the consumer-patient. The one notable exception was

ADMREB, and BINREB, where the prefixes DAY, ADM, and BIN denote the alternative measures of use/nonuse, and the common suffix REB denotes use/nonuse of regional and base hospitals.

the Newhouse and Phelps study where care was taken to obtain accurate estimates of the marginal coinsurance rate and gross price per day of care facing individual members of their cross-sectional sample.³⁷ Most investigators have, however, resorted to some very crude and/or inappropriate proxy measure of net price which seriously compromised both the logic of their exercise and the quality of their resultant elasticity estimates. Fortunately, this particular methodological problem does not arise in the present study. Recall that the data slated for analysis span the six year period 1966-71 and that two of the years, 1966-67, predate the introduction of the copayment policy. This immediately suggests that price effects can be isolated by comparing (suitable adjusted) estimates of families' use of inpatient facilities in the non-copayment years with corresponding estimates in the copayment years. An economical procedure to effect these "before-after" comparisons is to pool the six annual random samples into one overall sample containing dated observations on all of the variables in the model.³⁸ Dating itself is easily accomplished via the conventional "dummy" variable technique, i.e., defining a variable with the value 1 for all observations in a given year and zero for all observations in other years. Four such variables, one for each of the copayment years 1968-71, are slated for inclusion in each of the three basic equations; they are identified in the family summary record (Table 3.4) as D68,

³⁷Newhouse and Phelps, in Richard N. Rosett (ed.) The Role of Health Insurance in the Health Services Sector, pp. 261-320.

³⁸This procedure was employed by Phelps and Newhouse in their analysis of the medical services utilization data from the Palo Alto Group Health Plan. See Charles E. Phelps and Joseph P. Newhouse, "Effect of Coinsurance: A Multivariate Analysis," Social Security Bulletin, June 1972.

D69, D70, and D71. Subject to one caveat to be subsequently noted, the coefficients of these variables will be interpreted as the impact of copayment on, variously, the probability, volume, frequency, and duration of hospitalization.

Expectations are of course that the magnitudes of the coefficients will vary across different "narrow" definitions of the hospitalization event. This position was made clear in preceding discussion of the disaggregation scheme, so no further comment is required other than to emphasize that for any given definition, e.g., use in the i^{th} diagnostic category, it is reasonable to expect: (a) relatively small price effects in equations (1) and (3) where the dependent variable is the probability of hospitalization ($\text{Pr}[H]$) and the frequency (ADM) of hospitalization respectively; (b) relatively large price effects in equation (2) where the dependent variable is the volume (DAY) of hospitalization; and (c) relatively large price effects in the length of stay equation implicitly defined by equations (2) and (3). The underlying behavioural assumption here is that the admission decision is less discretionary than the discharge decision, and hence, is less likely to be affected by an "extra-medical" factor like copayment.³⁹ Thus, the copayment elasticities with respect to volume

³⁹This assumption was central to Rosenthal's analysis of the relationship between "price-payment" factors and length of hospital stay. He noted that

"Recent studies of the British experience support the hypothesis that length of stay is highly sensitive to economic characteristics, such as price and insurance, while the admissions rate is far more likely to reflect other differences in the characteristics of the population and the medical care system."

Rosenthal, in H.E. Klarman (ed.) Empirical Studies in Health Economics,

(i.e., patient-days) and duration (i.e., length of stay) ought to be larger than the copayment elasticities with respect to probability and frequency of admission.

Modelling the copayment experiment in this fashion does not of course pre-empt the case for cautious interpretation of results, a case that was developed with some care in the critique of the price effects literature. On the contrary, it reinforces the case, since the D68-D71 variables will capture effects unique to any given year that are not explicitly controlled for by the vector of socio-demographic and economic variables. These "other" effects - whose identity cannot be fully anticipated at this point - could result in one or more of the "price" coefficients being biased upward (i.e., more negative than its true value) or downward (i.e., less negative than its true value) depending upon whether they systematically decrease or increase the probability, frequency, volume, or duration of hospitalization. Thus, every effort must and will be made to resist indiscriminate labelling of the coefficients as "price effects" where and when rival labels might be applied with equal or greater conviction.

Having said this, it must be hastily added that this strategy is less problematic than it may appear. Thus, there is no threat of any upward bias due to advance selection of insurance, i.e., the assumed, though never proven, tendency in a voluntary insurance system for individuals and families to purchase policies providing benefits consistent with their expectations of illness and attendant demands for

p.102. In point of fact, the study to which Rosenthal refers presents no evidence on the price and insurance elasticities of admissions and lengths of stay. See M. S. Feldstein, Economic Analysis for Health Service Efficiency. Nevertheless, the assumption is reasonable.

care. Unlike their U. S. counterparts, whose utilization and insurance coverage data have thusfar been the basis for estimating net price effects, Saskatchewan families do not have this option. It, along with the benefit ("experience rating") principle on which the private insurance premia were structured, was eliminated with the introduction of the public universal hospital insurance program (SHSP) in 1947. Moreover, with the introduction of copayment in 1968, Saskatchewanians were not given the option of choosing between SHSP policies with and without the copayment clause nor were they able to purchase private insurance against copayment charges since such insurance was prohibited by law. Hence, the point-of-service price for inpatient care can be assumed exogenously given to all SHSP beneficiaries, and, by econometric extension, the price variables can be assumed to be free of "simultaneity bias." Second, the price coefficients will not contain any downward bias due to capacity constraints on hospital bed-use. This might have been a problem had the study period been characterized by copayment in the "before" years and non-copayment in the "after" years (i.e., by an initial sub-period of some price rationing followed by a sub-period of no price rationing), since in such quasi-experimental circumstances the hypothesized price-induced shifts in utilization might not have occurred due to full or near full occupancy of the available bed stock during the copayment years. A supply effect of this sort would clearly confuse the picture by producing artifactually small "price" coefficients. But again this is not a problem. Nor, thirdly, is there any reason to believe that the copayment and non-copayment years were fundamentally dissimilar in epidemiological terms. To be sure, the true incidence and prevalence of various illnesses and

conditions cannot be ascertained or even inferred from any existing morbidity and mortality data, but it seems extremely doubtful that Saskatchewan families were suddenly more (less) healthy during the copayment period. Hence, the coefficients are not likely to be biased upward (downward) on this account. Fourth, it is possible to evaluate the impact of copayment by examining the behaviour of a control group, namely a sub-sample of socially assisted families who were exempted from payment of the utilization fees. Their utilization experiences before and after copayment provide a standard for appraising the extent to which the D68-D71 variables capture non-price effects, and thereby a means of reducing the risk of labelling error. Finally, it can be confidently re-asserted that the socio-demographic and economic variables serve not only as substantively important variables in their own right, but also as methodologically important "controls" via which some of the more obvious ceteris are rendered paribus. It is to these variables that the discussion now turns.

As previously named and defined in the family summary record, the socio-demographic and economic variables include indicators and/or measures of family size and composition, family residence, family income, and social assistance status. The assumptions and hypotheses relevant to each of these family descriptors will be discussed in turn.⁴⁰

⁴⁰In what follows, no attempt will be made to comprehensively assess the status of these variables in the vast, and still rapidly growing multi-disciplinary literature on health services utilization. The reader unfamiliar with this literature is referred to the following very useful reviews: J.B. McKinlay, "Some Approaches and Problems in the Study of the Use of Services - An Overview," Journal of Health and Social Behavior, Vol. 13, No. 2, June 1972, pp. 115-152; L.A. Aday and R.L. Eichhorn, The Utilization of Health Services: Indices and

The hypothesized effects of family size and composition on hospital use derive from various assumptions. Thus, following Andersen, the fundamental assumption is made that these socio-demographic characteristics are not primary determinants or direct causes of hospitalization, but rather are factors which "predispose" families toward greater (lesser) use of inpatient facilities.⁴¹

As he puts it

Some families have a propensity to use more health services than other families. This propensity toward use can be predicted by family characteristics which exist prior to the onset of specific episodes of illness. Families with these characteristics are more likely to use health services even though the characteristics are not directly responsible for health service use. Such family characteristics are defined as predisposing conditions.⁴²

Four measurable characteristics which occupy this conceptual role in Andersen's model, and which are slated for empirical duty in this study are the age, sex, and marital status of the family head, and the number of dependent children.

Expectations in respect of each of these variables are based on a variety of ancillary assumptions. Age of the family head is assumed correlated with the ages of other family members, and is therefore taken as a proxy for the age structure of the family, or equivalently, as one indicator of the family's "life cycle" position. Add to this the fact

Correlates, A Research Bibliography (Washington: U.S. Dept. HEW Publication No. (HSM) 73-3003, December 1972); J.G. Anderson, "Health Services Utilization: Framework and Review," Health Services Research, Vol. 8, No. 3, Fall 1973, pp. 184-199.

⁴¹Ronald Andersen, op.cit., p.15.

⁴²Ibid. Andersen conceptualizes a three stage utilization model consisting of predisposing, enabling, and need components. The latter two components relate to families' ability to secure services, and their perceived need for such services respectively.

that many illnesses and conditions are known to be age-related, and it follows that the age effect ought to account for a significant proportion of the observed inter-family variations in hospital use. It is also expected that the age-use (nonuse) relationships will exhibit variations across different "narrow" definitions of the hospitalization event. For example, it is commonsensical (and epidemiologically sound) to anticipate an inverted U-shaped relationship in the "use/nonuse" equation (i.e., basic equation [1]) defined for obstetrical diagnoses, since childbearing reaches a bio-cultural peak in the 25-35 age range. The corresponding relationship for circulatory diagnoses is, in contrast, likely to approximate positive linearity, since the risk of serious heart, vein, and artery diseases is known to increase with age. Needless to say, it is quite impractical, indeed quite unnecessary, to fully catalogue expectations of this sort in each and every "basic" equation for each and every "narrow" definition of hospitalization. Instead, it suffices to acknowledge that the disaggregate data may contain a diversity of age-use relationships - some approximately linear, some distinctly non-linear - and to allow for this in the statistical model via alternative specifications of the age effect. Pursuant to the implementation of this strategy, the family summary record includes a continuous age variate, AGE, and a set of categorical age variates, A1 through A6.⁴³

Categorical variates are also employed to describe the marital

⁴³The continuous variate imposes linearity on the age-use relationships, while the categorical variates allow for non-linearities over ten-year age intervals. Further discussion of these variates and their role in testing the hypothesis that "older" families are more affected by copayment than "younger" families is deferred until the next sub-section where the details of the statistical model are unveiled.

status and sex of the family head. SINGLE MALE and SINGLE FEMALE identify SHSP families where, respectively, the female and male spouses are absent. On the reasonable assumption that absence of either spouse limits the opportunity for home care and recovery, it may be hypothesized that "singles" are more predisposed toward hospital use than "couples".⁴⁴ Unfortunately, however, the model is not well-suited to detecting this effect in the data, since intra-family aggregation of the inpatient data definitionally increases (decreases) the probability, volume, and frequency of use in two (one) spouse families. In consequence of this size effect, the coefficients of the SINGLE MALE and SINGLE FEMALE variables are likely to be downward biased (i.e., less positive than they would be in a model featuring the individual family member as the unit of observation), perhaps to the point of carrying "perversely" negative signs.⁴⁵ For this reason, these variables are assigned the less onerous, but nevertheless important task of controlling for this particular size effect.

There is of course another size effect of which account must be taken, namely, the effect on use (nonuse) of the number of children in the family. Due again to intra-family aggregation, this effect will doubtless "explain" an appreciable fraction of the total variance in

⁴⁴While Andersegg treats marital status and sex as "predisposing" conditions, he neither formulates nor tests any specific hypothesis regarding the nature of utilization differentials across one and two-spouse families. The "home care" assumption clearly enhances the logic of the exercise.

⁴⁵Be it noted, however, that in the average length of stay equation implicitly defined by basic equations (2) and (3), the coefficients should be unaffected by this bias, since average length of stay is not "definitionally" affected by the number of adult family members.

each of the basic utilization variables. Equally likely is the prospect of considerable variation in the strength of this effect across the alternative "narrow" definitions of these variables, especially those which are diagnostic specific. Thus, for obvious epidemiological reasons, relatively weak effects are anticipated when the basic equations are referrent to diseases of the circulatory, digestive, or genito-urinary systems, while a relatively strong effect is anticipated when the basic equations are referrent to diseases of the respiratory system.⁴⁶ Furthermore, for any given definition of hospitalization, the incremental effect of additional children may be constant or variable over the relevant range of family sizes. Previous studies of the relationship between family size and the utilization of ambulatory medical services per child generally indicate that marginal service use declines as family size increases, a finding consistent with the assumption that large families include a number of adolescents whose health status is "good" and whose service use is consequentially "low".⁴⁷ To determine if similar non-linearities are present in the Saskatchewan data, the basic equations will be specified and estimated with a set of categorical size variates; they are identified in the

⁴⁶As in the case of the age effect, a priori differences of this sort need not be fully enumerated in each and every basic equation for each and every "narrow" definition of hospitalization. Where and when they surface in the empirical analysis, appropriate interpretative comment will be made.

⁴⁷For reviews of these studies see: Litman, op. cit.; and Ronald Andersen and Judith Kasper, "The Structural Influence of Family Size on Children's Use of Physicians' Services," Journal of Comparative Family Studies, Vol.IV, No.1, Spring 1973, pp.116-130. For indirect evidence that the marginal use of hospital services declines as family size increases see Grover Wirick, Hospital Use and Characteristics of Michigan's Blue Cross Subscribers (Ann Arbor: University of Michigan, Bureau of Hospital Administration, 1972), p.55.

family summary record as C1 through C5+. In addition, experiments will be conducted with a continuous size variate (SIZE) to determine if there is any information loss ascribable to a linear specification of the size effect.

The URBAN variable categorically differentiates family resident in cities (population > 10,000) from families resident in towns, villages, and rural municipalities. Its primary role in the model is to control for known urban-rural disparities in the utilization of active treatment facilities. More specifically, it will accommodate the fact that, over the study period, rural residents experienced approximately 46 per cent more separations per 1000 and 25 per cent more patient-days per 1000 than did urban residents.⁴⁸ Being representative of these two sub-populations, the sample data will almost certainly contain directionally similar residence effects. However, it would be presumptuous to expect effects of correspondent magnitude in all prospective analyses of the sample data, since: (a) the figures just cited are based on utilization rates per 1000 beneficiaries, not per 1000 families; (b) the figures are unadjusted for age, sex, and other arguably relevant differences in the rural and urban populations (of

⁴⁸The following data for 1966 and 1971 illustrate these disparities in greater detail.

	Separations per 1000		Patient-Days per 1000		Average Length of stay	
	1966	1971	1966	1971	1966	1971
Rural	232	249	2113	2170	9.1	8.7
Urban	162	169	1717	1700	10.6	10.0
Rural/Urban Ratio	1.45	1.47	1.23	1.28	.86	.87

Source: Saskatchewan Hospital Services Plan, Annual Reports, 1966 and 1971 (Regina: Queen's Printer).

which account will be taken in the model); and (c) the figures relate to a single, "broad" definition of hospitalization. Thus, while the coefficients of the URBAN variable will likely carry a negative sign in the three basic equations (and a positive sign in the implicit average length of stay equation),⁴⁹ it will, at the same time, likely exhibit substantial size variation across the family-aggregate, "narrow" definitions of hospitalization.

It would also be presumptuous to view the URBAN variable as a good proxy for geographic differences in the supply of hospital beds, or, by extension, as an appropriate vehicle for testing the hypothesis that a Feldstein-type "pure availability effect" is operative in Saskatchewan.⁵⁰ The reasons for this are very simple. In the first place, the URBAN variable only identifies the residence of the family and not the location of the hospital(s) where the family receives its inpatient care. Second, the dependent variables do not differentiate between families' use of local and non-local hospitals, i.e., between use of rural (urban) hospitals by rural (urban) families or use of urban (rural) hospitals by rural (urban) families.⁵¹ Third, the provincial data show quite clearly that the relatively high incidence of hospitalization among rural residents has less to do with their access

⁴⁹As suggested by the length of stay data reported in footnote 48, *supra*.

⁵⁰For a discussion of this effect, see Chapter 2, pp. 23.

⁵¹To be sure, the institution-specific variables, DAYREB, ADMREB, and BINREB will capture most of the "cross-haul" involved when rural areas export patients to urban (regional and base) hospitals, but this is hardly equivalent to capturing a "pure availability effect." Indeed, as the text subsequently notes, it illustrates how easily the latter can escape detection.

to rural beds, and more to do with their access to both rural and urban beds. In 1971, for example, the nominal stock of rural beds available to rural residents was 5.4 per 1000, but due to "cross-haul" of rural residents to urban hospitals, the stock of rural and urban beds effectively available to them was 6.9 per 1000.⁵² In consequence of this phenomenon, the stock of urban beds available to urban residents was reduced from a nominal value of 10.1 per 1000 to an effective value of 5.9 per 1000.

That these rural-urban discrepancies between nominal and effective bed stocks complicate the bed-use/ bed-supply relationship should be obvious. They imply a negative correlation between total use and nominal bed availability, and a positive (but largely tautological) correlation between total use and effective bed availability.⁵³ And of course, they also militate against searching for a pure availability

⁵²The number of beds at a given hospital effectively available to residents of a given area (defined by SHSP's five digit geo-code system), is calculated by multiplying the hospital's rated bed capacity by the area's share of the hospital's total annual case load. Repeating this for each and every hospital to which area residents have been admitted during the year, and summing the resultant figures yields the total number of beds effectively available to the area's residents during the year. Aggregating these bed counts across all rural (urban) areas, and dividing by the appropriate rural (urban) population count yields the "effectively available" bed-population ratio for rural (urban) residents.

⁵³Failure to control for "cross-haul" marred an early attempt to implicate the nominal bed stock in areas of "high" and "low" hospitalization. See F.B. Roth, *et al.*, "Some Factors Influencing Hospital Utilization in Saskatchewan," Canadian Journal of Public Health, Vol.46, No.8, August 1955, pp.303-323.

effect armed only with the URBAN variable.⁵⁴ All that can and will be reasonably asked of this variable is to capture the net effect of rural vs. urban residence on the probability, volume, frequency, and duration of hospitalization. In performing this function, it will not only add new and useful information on the distribution of insured hospital care among Saskatchewan families, but will also render methodological assistance to identifying the copayment effect.⁵⁵

Finally, attention shifts to two variables - family income and social assistance status - whose incorporation in the model affords a unique opportunity to explore the relationship between families' socio-economic position and use (nonuse) of inpatient services with data

⁵⁴Mention should be made of the fact that URBAN is a much better proxy for the availability of physician service. Thus, over the study period, the active general practitioner-population ratio was approximately 44 per cent higher in urban areas than in rural areas. (The total active practitioner-population ratio was approximately 235 per cent higher due to the concentration of specialists in urban areas.) Medical Care Insurance Commission, Annual Reports, 1966-71 (Regina: Queen's Printer). In light of this information, it is not surprising that urban families utilize ambulatory medical services to a significantly greater extent than rural families. (See R.G. Beck, "Analysis of the Demand for Physicians' Services in Saskatchewan," [unpublished doctoral dissertation], University of Alberta, 1971.) Combining this result with the fact that the nominal number of urban beds per urban physician is at 7.4 (1971) much lower than the corresponding rural ratio of 13.0 (1971), it might be argued that relatively "high" hospital use among rural families is indeed a "supply effect" (if not a pure availability effect) due to the representative rural physician having more beds at his disposal than his urban counterpart, and a consequent greater propensity to hospitalize his patients. For indirect evidence supportive of this reasoning, see Roth, op.cit.; and Milton Roemer, "Hospital Utilization and the Supply of Physicians," Journal of the American Medical Association, Vol. 178, No. 10, Dec.9, 1961, pp. 125-129.

⁵⁵In this latter connection, it should be noted that omission of the URBAN variable would almost certainly impart an upward bias to the price coefficients, since the proportion of "low use" urban families in both the population and the sample was marginally higher in the copayment years than the non-copayment years.

generated under a public universal health insurance program. The reader may be surprised to learn that the extensive literature on health services utilization contains few relevant precedents for this exercise. Granted, many researchers have examined some variant of the income-use relationship,⁵⁶ but their almost total reliance on the U.S. voluntary private and non-universal public (Medicare, Medicaid) insurance data leaves students of the Canadian health care system little choice but to relegate their findings, and in some instances their underlying conceptual frameworks, to the "interesting but uninformative" category. Evidence in eight studies that "the higher the income, the greater are the expenditures for hospital services"⁵⁷ must be treated this way, as must related hypotheses based on the assumption that income constrains (or, a la Andersen, is a factor "enabling")⁵⁸ the purchase of both health insurance and health services. Text-book assumptions of this sort are doubtless reasonable in the U.S. setting where the ability-to-pay principle is alive and well, but they are decidedly counter-intuitive in the Canadian setting where it is not. If there are income effects on use in a system where hospital and medical services are "freely" and universally distributed on the philosophical basis of ability to benefit, they are surely more indirect than their U.S. counter-parts. More will be said of this very shortly.

⁵⁶According to one non-exhaustive, and now somewhat dated bibliography, "income effects" of one sort or another have been analyzed in no less than seventy studies. See L.A. Aday and R.L. Eichhorn, op.cit., p.7.

⁵⁷ Ibid., p.23.

⁵⁸Is it not a little curious that what is "enabling" to the sociologist is "constraining" to the economist?

As for the more relevant precedents, there are, so far as can be determined, only three: the study by Enterline, et. al., of physician utilization in Montreal before and after Medicare;⁵⁹ the study by R.G. Beck of physician utilization in Saskatchewan over the post-Medicare period 1963-1968;⁶⁰ and the study by P. Manga of hospital and physician utilization in Ontario over the twelve-month Medicare period, April 1974 to March 1975.⁶¹ In each of these studies, attempts were made to identify family income-related variations in the utilization of selected publicly insured health services. Marked inter-study differences in data and method do, however, preclude any direct comparison of results. Suffice it to say that: (1) Enterline's cross-tabulation of household survey data yielded an inverse, but not monotonically decreasing

⁵⁹P.E. Enterline, et.al., "The Distribution of Medical Services Before and After 'Free' Medical Care - The Quebec Experience," New England Journal of Medicine, Vol.289, No.22, November 29, 1973, pp. 1174-78.

⁶⁰Beck, op.cit. For other related studies of the Saskatchewan experience see: R.G. Beck, "Economic Class and Access to Physician Services Under Public Medical Care Insurance," International Journal of Health Services, Vol.3, No.3, 1973; R.G. Beck, "The Effects of Copayment on the Poor," Journal of Human Resources, Vol.IX, No.3, 1973; R.G. Beck and J.M. Horne, "Economic Class and Risk Avoidance: Experience Under Public Medical Care Insurance," Journal of Risk and Insurance, Vol.XLIII, No.1, March 1976; and R.G. Beck, "Some Dynamic Effects of Copayment on Utilization of Medical Services in Saskatchewan," in R.D. Fraser (ed.) Health Economics Symposium - Proceedings of the First Canadian Conference (Kingston: Industrial Relations Center, Queen's University, 1976).

⁶¹P.Manga, A Benefit Incidence Analysis of the Public Medical and Hospital Insurance Programs in Ontario (Toronto: Ontario Economic Council, June 1976), (draft copy, 51 pages).

relation between annual family income and frequency of physician visits per family member per year;⁶² (2) Beck's multi-variate regression analysis of the Saskatchewan (MCIC) claims data yielded a statistically significant but very weak direct relation between family income and use of all physicians' services per family per year;⁶³ and (3) Manga's cross-tabulation of the Ontario claims data (augmented with census tract income information) yielded an inverse (but, like Enterline's, non-monotonically decreasing) relationship between annual family income and frequency of physician encounters per family member per year, and, as well, evidence that the poor (income < \$4,000) experienced significantly more hospital days and longer lengths of stay than did other income classes.⁶⁴ Based on these and other pertinent results contained in each study, it may be tentatively concluded that the indirect effects of family income on use of "free" health services are quantitatively small, and seemingly less negative than might be expected given the redistributive spirit of the public programs and conventionally held notions concerning the distribution of medical need in the population. It will be interesting to see whether kindred effects surface in the present study, not least because it, unlike those cited above, focuses entirely on in-hospital utilization, and, as well, entails considerable data disaggregation.

Expectations in respect of the INCOME and WELFARE variables are

⁶²Enterline, et al., op.cit., p.1175.

⁶³In respect of this result, Beck argues that the low income elasticity of use (+0.05) is "consistent with the successful operation of the insurance scheme." Beck, op.cit., (1971), p.80.

⁶⁴Manga, op.cit., p.29.

based on several considerations. First, as mentioned above, positive income effects are logically contra-indicated by the "free" care policy: quantities consumed are not directly constrained by inability-to-pay. Second, as between zero and negative income effects, the latter are most plausible since: (a) family income may be reduced in consequence of hospitalization, especially hospitalization of the prime wage-earner;⁶⁵ (b) in the absence of the requisite income maintenance data (sick leave benefits, etc.), it is reasonable to assume that the risk of income loss due to serious illness is greatest among low income families; (c) high income families tend to be better educated, and hence might well be more cognizant of the benefits to be derived from early detection of illness and/or the maintenance of good health;⁶⁶ and (d) notwithstanding major ambiguities in the direction of causation, an inverse relationship between socio-economic status and indicators of illness and utilization behavior has been a common finding in the medical sociological literature.⁶⁷ Third, the explicit distinction between socially assisted and socially unassisted families is indicated, since: (a) a low income

⁶⁵For an analysis of this and other employment-earnings effects based on the U.S. data see Harold S. Luft, "The Impact of Poor Health on Earnings," The Review of Economics and Statistics, Vol.LVII, 1975, pp. 43-57.

⁶⁶Alternatively, it might be argued, following Grossman, that education makes the individual a more efficient producer of good health, and hence less reliant on professional health "services. See Michael Grossman, The Demand for Health: A Theoretical and Empirical Investigation (New York: National Bureau of Economic Research, Occasional Paper 119, 1972), pp. 27-28. For Canadian evidence that income and education are positively correlated with preventive health behavior, see David Coburn and Clyde R. Pope, "Socio-economic Status and Preventive Health Behavior," Journal of Health and Social Behavior, Vol.15, No. 2, June 1974, pp. 67-78.

⁶⁷See McKinlay, op.cit., for a review of the pertinent studies.

family is not definitionally equivalent to a socially assisted family; (b) it identifies a group of seriously disadvantaged families who, it can be fairly assumed, must daily contend with a uniquely broad spectrum of health, housing, economic, and social problems; and whose utilization of inpatient facilities is, not surprisingly, well above the provincial average;⁶⁸ (c) in consequence of (a) and (b), it is reasonable to expect that the income-use relationship will be discontinuous (i.e., non-linear) over the relevant range of low incomes, i.e., there will be a net welfare effect; and (d) socially assisted families were exempted from payment of the utilization fees. Thus it is anticipated that the coefficients of the INCOME and WELFARE variables will be negative and positive respectively in the three basic equations. Needless to say, however, their magnitudes across all "narrow" definitions of use (nonuse) are not ascertainable a priori.

This completes the discussion of the independent variables and their hypothetical relationships to the dependent utilization variables. These relationships, together with the assumptions on which they are based, comprise the conceptual model adopted in this study for analyzing families' use (nonuse) of inpatient facilities before and after the introduction of copayment. Like all models, it has its limitations. It abstracts only certain aspects of the complex reality to which it relates, it entails simplifications and approximations, and it embodies

⁶⁸In particular, "hospital utilization studies based on provincial data indicate that on the average social assistance beneficiaries utilize hospitals 50% more than the average, after adjusting for age/sex differences." See Graham Clarkson, A Study of Health Care Services in Regina with Special Emphasis on Hospital Facilities and Services (Edmonton: J. Graham Clarkson Consultants Ltd., February 1973), p.47.

few if any unconventional wisdoms. On the other, more positive hand, it has elements of "special tailoring" which improve upon previous formulations of the copayment question, and equally compelling, a built-in consistency with the available data which guarantees its operationality. Having apprised the reader of these facts, it remains only to discuss some particulars of the implementation strategy.

(c) The Statistical Model: Regression Equations and Estimation Method

In this sub-section the conceptual model is formulated in terms which directly anticipate the empirical analysis. Presented and discussed are the various estimating forms of each basic equation and the associated method of multivariate regression analysis via which the net effects of copayment and other factors will be simultaneously ascertained. At the outset, however, it is well to restate the three basic equations in their general form, and also, to clarify the analytical status of the earlier noted (page 105) average length of stay equation implicit in the model.

As previously specified the three basic equations are:

$$(1) \text{Pr}(H)_{it} = f(P_{it}, Z_{it})$$

$$(2) \text{DAY}_{it} = g(P_{it}, Z_{it})$$

$$(3) \text{ADM}_{it} = h(P_{it}, Z_{it}).$$

The average length of stay relationship implicit in the model follows straightforwardly from the identity:

$$LS = \frac{\text{DAY}}{\text{ADM}}$$

where LS is the average length of stay, DAY is the number of patient-

days and ADM is the number of admissions. Thus equations (2) and (3) together yield:

$$(4) \quad LS_{it} = j(P_{it}, Z_{it})$$

where LS_{it} is the average duration of stay per admission for the i^{th} family in year t . While this equation will not be directly estimated, its parameters are ascertainable from the parameters of equation (2) and (3). To illustrate, consider the following simple "use-price" relationships:

$$(2') \quad DAY = a + bP$$

$$(3') \quad ADM = c + dP.$$

From the definition of average length of stay, it follows that:

$$LS = \frac{a + bP}{c + dP}$$

Differentiating LS with respect to P yields:

$$\frac{\partial LS}{\partial P} = \frac{b}{c + dP} - \frac{(a + bP)d}{(c + dP)^2} = \frac{b}{ADM} - \frac{DAY}{ADM^2} \cdot d$$

Substituting the estimated values of b and d and the mean values of ADM and DAY yields the desired price parameter in the LS equation. In general, and regardless of specification, the parameters of equations (2) and (3) provide sufficient information to obtain the corresponding parameters in equation (4). For this reason, subsequent discussion is confined to the adopted specifications of the three basic equations.

Equation (1) addresses the probabilistic aspects of hospitalization and is representable as a linear probability function in which the regressand is a dichotomous or binary variable taking on the

values one or zero when the family is a user or nonuser of hospital services respectively.⁶⁹ This function is fitted to all the observations in the sample and its parameters are obtained via the ordinary least squares regression method. The estimated value of the dependent variable may be interpreted as the conditional probability that hospitalization will occur in any year, given family characteristics defined by the vector of socio-demographic and economic variables.

Two econometric problems peculiar to the linear probability function should be noted. The first problem stems from the assumed linearity between the regressand and the regressors. Linearity implies that the predicted value of the regressand can fall outside the zero-one range, i.e., outside the range of admissible probabilities. Following Feldstein, the linear approximation is adopted in preference to other more appropriate non-linear specifications "because of the greater ease of calculation and interpretation."⁷⁰ Moreover, the full extent of the problem can only be assessed on an ex post facto basis. The second problem is heteroscedasticity of the error terms due to the binary nature of the dependent variable. For this reason ordinary least squares parameter estimates are inefficient, although they are not biased. While the generalized least squares regression is appropriate in these circumstances, the costs of carrying out the requisite two-step

⁶⁹At this point the reader might consult Table 3.4, where these binary variables (prefixed "BIN") are specifically named and defined.

⁷⁰M.S. Feldstein, Economic Analysis for Health Service Efficiency (Amsterdam: North Holland Publishing Co., 1967), p.238. The other specifications, and, in particular, the probit and logit models, are discussed in Henri Theil, Principles of Econometrics, (New York: John Wiley and Sons Inc., 1971) pp. 628-635.

estimation procedure are prohibitively great in large sample situations.⁷¹

The statistical analogues of equations (2) and (3) are less problematic since the regressands are continuous variables taking on positive values, defined as the volume and frequency of hospitalization respectively. Restricted to the user sub-samples, these linear utilization functions generate estimates of the expected levels of use for families with characteristics defined by the values of the independent variables. Thus they also yield conditional probabilities, but, in this instance, the condition is the family's presence in the user sub-sample.

Three specifications of the linear probability function and the linear utilization functions will be estimated. The specifications differ in their assumptions about the nature and stability of the structural relations. Since the specifications are the same for each basic equation they will be illustrated with reference to equation (1) only, i.e., the linear probability function. The first specification, henceforth known as "model one," allows for non-linearities in the size-use and age-use relationships. It may be expressed symbolically as:

$$Y_{it} = a_0 + a_1 \text{SINGLE MALE} + a_2 \text{SINGLE FEMALE} + a_3 \text{URBAN} \\ + a_4 \text{WELFARE} + a_5 \text{INCOME} + \sum_{i=1}^4 b_i D_i + \sum_{i=1}^5 c_i C_i \\ + \sum_{i=1}^6 d_i A_i + u_{it}$$

⁷¹Feldstein rejects the GLS procedure on the basis that "the gain in efficiency probably does not justify the additional calculations." Feldstein, *op. cit.*, p.239. His sample contained approximately 3500 observations.

where the independent variables are defined as in Table 3.4; u_{it} is a stochastic error term; and Y_{it} is any one of the various binary dependent variables previously defined. The coefficients relevant to determining the non-linear effects of family size and age of head are the c_1 and d_1 respectively. The price effects are given by the b_1 .

Specification two, henceforth known as "model two," imposes linearity on the age and size relationships, and accordingly inserts these variables in their continuous form. Symbolically:

$$Y_{it} = a_0 + a_1 \text{SINGLE MALE} + a_2 \text{SINGLE FEMALE} + a_3 \text{URBAN} \\ + a_4 \text{WELFARE} + a_5 \text{INCOME} + a_6 \text{SIZE} + a_7 \text{AGE} \\ + \sum b_1 D_1 + u_{it}$$

where it suffices to note that SIZE and AGE are the continuous size and age variates. This specification affords an opportunity to assess the information loss ascribable to the assumption of constant incremental size and age effects given by the coefficients a_6 and a_7 respectively. In addition, it serves as a transition to the third and final specification.

The third specification, "model three," features selected interaction effects. Symbolically:—

$$Y_{it} = a_0 + a_1 \text{SINGLE MALE} + a_2 \text{SINGLE FEMALE} + a_3 \text{URBAN} \\ + a_4 \text{WELFARE} + a_5 \text{INCOME} + a_6 \text{SIZE} + a_7 \text{AGE} \\ + a_8 (\text{WELFARE} \times \text{INCOME}) + \sum b_1 D_1 + \sum e_1 (D_1 \times \text{INCOME}) \\ + \sum f_1 (D_1 \times \text{WELFARE} \times \text{INCOME}) + \sum g_1 (D_1 \times \text{SIZE}) \\ + \sum h_1 (D_1 \times \text{AGE}) + u_{it}$$

where it suffices to note that the interactor variable with coefficients a_8 , e_1 , f_1 , g_1 , and h_1 are straightforward products of the relevant component variables. This specification differs from models 1 and 2 by

allowing price effects to be captured not only by the shift parameters b_1 , but also by the slope coefficients on the age, size and income variables. That is, the cross-sectional relationships between age-use, size-use, and income-use may be altered by the introduction of copayment. Under the hypothesis that "older" families were more affected by copayment than "younger" families, the marginal age effect in the copayment years, given by a_7 , should be larger than the marginal age effect in any one or all of the copayment years, given by $a_7 + h_1$. By similar reasoning, the marginal size effect in the non-copayment years, given by a_6 , should be larger than the marginal size effect in the copayment years, given by $a_6 + g_1$, under the hypothesis that "larger" families were more affected by copayment than "smaller" families.

The investigation of income and welfare effects is also substantially refined in this specification by the inclusion of both first order and second order interaction terms. A two-way interactor between income and welfare allows the slope of the income-use relationship to vary between the socially assisted and socially unassisted sub-samples. A two-way interactor between income and copayment allows the marginal income effect in the non-copayment years, given by a_5 , to differ from its value in the copayment years, given by $a_5 + e_1$; positive coefficients on any of the e_1 would suggest that copayment had less effect on high income families than on low income families. These two-way interactors, coupled with the three-way interactor term between income, welfare, and copayment identify in their various permutations ten unique income-use relationships. The relevant parameters of these relationships are identified in Table 3.5. The fact

TABLE 3.5

Parameters of the Partial Income-Use Relationship

Period	Sub-sample	Intercept	Slope
1966-67	Non-Welfare	a_0	a_5
	Welfare	$a_0 + a_4$	$a_5 + a_8$
1968	Non-Welfare	$a_0 + b_1$	$a_5 + e_1$
	Welfare	$a_0 + a_4 + b_1$	$a_5 + a_8 + e_1 + f_1$
1969	Non-Welfare	$a_0 + b_2$	$a_5 + e_2$
	Welfare	$a_0 + a_4 + b_2$	$a_5 + a_8 + e_2 + f_2$
1970	Non-Welfare	$a_0 + b_3$	$a_5 + e_3$
	Welfare	$a_0 + a_4 + b_3$	$a_5 + a_8 + e_3 + f_3$
1971	Non-Welfare	$a_0 + b_4$	$a_5 + e_4$
	Welfare	$a_0 + a_4 + b_4$	$a_5 + a_8 + e_4 + f_4$

that socially assisted families were exempted from payment of the utilization fees suggests that the income-use relationship should be more stable for the socially assisted sub-sample over the non-copayment and copayment periods.

It must be emphasized, in conclusion, that each of the three basic equations is alternatively specified in the form of model one, model two, and model three. In total therefore, there are nine structural equations to be estimated. Furthermore, each of the nine equations will be estimated using the "narrow" and "broad" definitions of hospital use described in sub-section (a). All equations will be estimated via the ordinary least squares regression method. The results of these statistical experiments are presented in the following two chapters. Chapter four presents the estimated linear probability functions, while chapter five presents the estimated linear utilization functions.

CHAPTER IV

REGRESSION RESULTS: THE LINEAR PROBABILITY EQUATIONS

As mentioned in the previous chapter, the statistical model employed in this study to assess the effects of per diem user charges on the demand for hospital care involves the estimation of two distinct types of regression equations. First, linear probability functions with dichotomous dependent variables distinguishing user from non-user families are fitted to all the observations in the sample. Designed to accommodate the large number of families who do not experience hospitalization during any given year, these equations generate probabilities of use for families having specific demographic and economic characteristics, and, more importantly, allow these probabilities to change with the introduction of the copayment policy. Second, linear utilization functions with continuous dependent variables, alternatively defined as the number of patient-days and separations, are fitted to sub-sets of observations restricted to those families who experienced some hospitalization during the study period; these equations yield estimates of the expected level of utilization for various type of families given that hospitalization occurs, and relate these estimates to years when copayment did and did not apply in order to assess "price effects" among user families. The purpose of this chapter is to present and analyze results based upon the first of these two types of equations.

The repeated estimation of the linear probability function for each definition of contact with the hospital system employed in this study suggests a discussion running from the general to the particular. Accordingly, section one analyzes the likelihood of contact with data unclassified as to diagnosis, admitting physician, or service capability of the host institution. Results specific to each of these sub-classifications of contact are then analyzed in sections two, three, and four respectively. A feature common to all four sections is the presentation of estimates for three specifications (Model 1, Model 2, and Model 3) of the probability equation. Discussed in detail in the previous chapter, these specifications differ in their assumptions about the nature (linear vs. non-linear) and stability (pre-and post-copayment) of the relationship between probability of hospitalization and select demographic and economic variables. The chapter concludes with a summary discussion of the evidence and its contribution to our understanding of the hospitalization process.

Non-Specific Hospitalization Probabilities.

The dependent variable in the following analysis (BINDIS) defines the hospitalization event as one involving admission to and subsequent discharge from an acute care general hospital regardless of the patient's illness or condition, the admitting physician's practice mode or specialty, or the host institution's service capability. Ordinary least squares estimates of the coefficients for the three variants of the probability equation based on this broad dichotomization

of the sample families are presented in Table 4.1.¹

Three general feature of these equations are worthy of explicit mention before discussing their component variables and associated coefficients in detail. One is their low overall ability to explain why in any year some families do and others do not enter the hospital system. The R^2 's of 10-11 per cent reported in the third to last row of the table clearly imply that a variety of undifferentiated, chance factors play a dominant role in this process. This finding is hardly surprising given the highly micro-analytic nature of the equation, plus the fact the event itself has always been classed as insurable precisely because of this essential unpredictability. In any case the R^2 values are statistically significant and compare quite favourably with those reported in other health care utilization studies where the family or other elemental consumption unit constituted the unit of observation. A second feature of the equations is their generation of probabilities within the unit (0,1) interval; no combination of the demographic, economic, and time (copayment) variables results in estimates less than zero or greater than one. When it is recalled that the equations are linear in their parameters and hence do not constrain the dependent variable to logically possible values, this finding suggests that little would have been gained had more costly probit models incorporating this constraint a priori been estimated. Third, it should be borne in mind that while the ordinary least squares estimates are unbiased, they are inefficient on account of the heteroscedastic error terms. The resultant bias in the reported values argues for cautious interpretation

¹Corresponding estimates for a slightly broader definition of the hospitalization event (BINDTX) are presented in Appendix F.

TABLE 4.1

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINDIS
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE (S)		0.05904 (78.72)	0.06227 (56.10)
D68.S			-0.00358* (-1.96)
D69.S			-0.00856 (-4.65)
D70.S			-0.00386 (-2.06)
D71.S			-0.00449 (-2.39)
C1	0.16171 (43.12)		
C2	0.18935 (47.70)		
C3	0.23062 (49.70)		
C4	0.28499 (51.17)		
C5	0.35591 (61.68)		
AGE (A)		0.00081 (11.57)	0.00055 (4.58)
D68.A			-0.00036* (-1.89)
D69.A			0.00068 (3.58)
D70.A			0.00069 (3.63)
D71.A			0.00044 (2.32)
A1	0.05775 (14.15)		
A2	-0.02218 (-5.09)		
A3	-0.03160 (-7.75)		
A4	0.00434* (1.07)		
A5	0.07370 (15.85)		
A6	0.16234 (29.62)		

TABLE 4.1 Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.13260 (-39.58)	-0.10267 (-29.42)	-0.10331 (-29.52)
SINGLE FEMALE	-0.11514 (-31.63)	-0.08031 (-21.25)	-0.08132 (-21.46)
URBAN	-0.06385 (-27.52)	-0.05704 (-24.48)	-0.05723 (-24.56)
WELFARE	0.10488 (17.51)	0.11218 (18.63)	0.11240 (13.40)
INCOME	-0.00204 (-8.50)	-0.00262 (-10.92)	-0.00336 (-7.64)
W.Y			0.00019*
D68.Y			0.00099*
D69.Y			(1.38)
D70.Y			0.00154
D71.Y			(2.17)
			0.00130*
			(1.86)
			0.00063*
			(.94)
D68.W.Y			0.00506*
D69.W.Y			(1.22)
D70.W.Y			-0.00408*
D71.W.Y			(-1.16)
			-0.00075*
			(-.18)
			0.00141*
			(.39)
D68	-0.00075*	-0.00209*	0.01787*
D69	(-.23)	(-.63)	(1.68)
D70	-0.00817 (-2.48)	-0.00743 (-2.24)	-0.02107*
D71	0.00479*	0.00491*	(-1.98)
	(1.44)	(1.47)	-0.02161*
	0.00659 (1.99)	0.00651*	(-1.99)
constant	0.25341	0.15722	-0.00385*
			(-.36)
R ²	0.11265	0.10111	0.16345
S.E.E.	0.42760	0.43036	0.10153
F	956.650	1541.186	0.43028
			608.167

*Insignificant at 95% Confidence Level

of the results.

Consider first the Model 1 results. The constant term in this equation has a special interpretation (without parallel in Models 2 and 3) owing to the nature of the dummy variable schemes defining family characteristics and years when copayment did and did not apply. In particular, 0.253 is a point estimate of the probability that a rural couple with a head aged 18-24 years, no children, and no social assistance receipts experienced hospitalization in either of the non-copayment years 1966-1967. The coefficients on the dummy variables D68-D71 and the dummy variables describing number of children (C1,..., C5), age of head (A1,..., A6), absence of spouse (SINGLE MALE, SINGLE FEMALE), social assistance (WELFARE), and urban residence (URBAN), are the amounts by which this probability must be increased or decreased in order to derive point estimates for subsequent years and/or for families different in one or more respects from those included in this reference group.²

The uniformly positive and significant coefficients on the family size variables are consistent with the "common sense" hypothesis that the probability of hospitalization is directly related to the number of potential patients in the family. However, the form of the relationship exhibits a marked non-linearity as may be seen from the first differences of these coefficients which define the incremental probabilities associated with the presence of additional children. Their pattern -- 16 per cent, 3 per cent, 4 per cent, 6 per cent, and 7

²Though unnecessary for the purpose of illustrating the dummy variable scheme, mention should be made of the fact that the probabilities can be also adjusted for family income differences using the coefficient of the continuously defined income variate (Y).

per cent — indicates that the net effect of the first child is 3 per cent more than the combined net effects of the next three, and only 4 per cent short of the combined net effect of the next four or more. The most plausible explanation for this non-linearity is that it reflects age differences in the children and, specifically, an assumed tendency for children in "large" families to fall within the age range 5-18 years where the usual perils of early childhood have been overcome and the risk of hospitalization concomitantly reduced.³

Further evidence of the important role played by demographic factors in discriminating hospitalization probabilities is revealed by the coefficients on the variables describing the age and marital status of family heads. The age coefficients demonstrate quite clearly certain expected (and well-known) life cycle effects: hospitalization increases (above that observed for the 18-24 age group) during the childbearing years 25-34 (A1), decreases during the middle years 35-44 (A2), and 45-54 (A3) when accidents and illness are less common, and increases sharply during the later years 65-74 (A5) and 75+ (A6) when health problems requiring hospitalization are most likely to occur. The marital status coefficients (SINGLE MALE, SINGLE FEMALE) indicate that absence of either spouse reduces the likelihood of hospitalization. While consistent with the supposition that single, widowed, or divorced male and female heads may be less willing or able to leave the home environment (especially if dependent children are present), the

³Over the period 1966-71, separations per thousand males and females were consistently in excess of 250 for the age group 1-4, and consistently below 150 for the age group 5-18. Saskatchewan Hospital Services Plan, Annual Reports. (Regina: Queen's Printer, various annual issues).

magnitude of the coefficients militates against this interpretation; intuitively, they are more consistent with the scale or "fewer potential patients" effect previously argued in connection with the family size variables.

The coefficient on the location variable (URBAN) indicates that urban families stand a 6.4 per cent lesser chance of being admitted to hospital than their rural counterparts. This is not surprising, particularly in the light of evidence in Chapter 3 showing that the province's urban centres mix a relatively low effective bed-population ratio with a relatively high physician-population ratio.⁴ The implied "supply effect" -- physicians tend to treat more patients on an outpatient than on an inpatient basis when and where beds are scarce -- finds further support in Beck's earlier study where the same urban/rural distinction was drawn and where, with the notable exception of hospital visits, it was found that urban families make substantially greater use of physician's services.⁵

Two variables of particular interest in the equation are those that distinguish families according to income and social assistance status. Before assessing their performance, it is worthwhile recalling that their inclusion in the analysis has a somewhat unconventional rationale. It relates not to standard "ability to pay" considerations, for these are stripped of significance by the policy of "free" hospital care, but rather to "ability to benefit" considerations. On this reasoning, families at different levels of income are posited to have

⁴See page 130-131.

⁵Beck, op.cit., p. 88.

different needs for hospital care due to an assumed relationship between income and health status. The precise nature of this relationship is of course the subject of considerable controversy, largely because no study has convincingly isolated the independent effects of income and education on health status and care-seeking behavior. Suffice it to say, the assumption that medical needs are greatest among the lower income classes is a common one, particularly in social policy discussions. This, coupled with an "equal access to care" assumption, implies that the poor will receive more medical care, or at least more hospital care, than will the non-poor.⁶ The evidence germane to this hypothesis and the more general income-hospitalization relationship may now be examined.

As expected, the welfare coefficient (WELFARE) is positive, large in absolute value, and statistically very significant (being more than seventeen times its standard error). It implies, for example, that the probability of hospitalization for an assisted urban family composed of a female head (age 25-34) and two children (under the age of 18) is .436, fully 33 per cent higher than the probability of .321 for an identical unassisted family. This disparity leaves little doubt that welfare recipients and their dependents have more than their share of medical misfortune, and correspondingly confirms their status as a "high risk" group.

Equally interesting is the fact that families with incomes above

⁶To be sure, the definition of poverty employed in this study, namely, eligibility for social assistance, is a very stringent one, but it has the advantage of identifying a group of families among whom malnutrition and other health threatening problems are, it is generally agreed, most prevalent.

the social assistance, or extreme poverty level, face quite similar admission probabilities. The income coefficient (INCOME) is negative and significant, but very small. It indicates that the probability declines by only 0.002 (basis points) for every \$1,000 in additional family income, or, equivalently, that the mean income elasticity is only -0.05. The latter statistic can be put in keener perspective by noting its implication that the proportional difference in probabilities between a family with an income of \$6,000 and a family with an income of \$12,000 (a 100 per cent difference) is only 5 per cent. The apparent weakness of the income-hospitalization relation (adjusted for the welfare effect) is something of a surprise. A stronger inverse relationship was expected under the hypothesis that income and health status are positively correlated. The observed relationship, together with the above-noted welfare effect, would suggest that this correlation obtains only across two very disparate family types, namely, those who are and those who are not in receipt of social assistance. Among the unassisted lower, middle, and upper income groups it may well be that so-called "diseases of choice" (respiratory and circulatory problems associated with smoking, drinking, and lack of exercise) have compressed health status differentials to the point that risks of hospitalization are approximately equal. This is arguably the most plausible explanation of the results. At any rate, such surprise as it does elicit is nothing in comparison to what it would have been had the income coefficient been positive. Universal insurance programs are expressly tailored to eliminate the "normality" (positive income

elasticities) that apparently characterized hospital care under the old voluntary insurance and prepayment arrangements.⁷ In this broader policy evaluation sense, the observed income coefficient is a measure of S.H.S.P.'s success in engineering a more equitable distribution of care.

The copayment policy in force over the period April, 1968 to August, 1971 is modelled in this equation by the dummy variables D68-D71. They test the null hypothesis that the intercept of the probability function has the same value in each of the copayment years as it does in the non-copayment years 1966-67. As argued in Chapter 3, the expected magnitude of the coefficients on these shift parameters is somewhat ambiguous depending upon the assumed discretionary vs. non-discretionary nature of the admission decision. When technological imperatives dominate the physician's decision to recommend hospitalization, the family is unlikely to be influenced by a direct charge of \$2.50 per day, and the null hypothesis will be sustained. As it turns out, the coefficients largely support this medically oriented view of the decision-making process. Three of them (D68, D70, and D71) clearly deny a price effect by virtue of being insignificant and/or "wrongly" signed, while the fourth (D69) can be challenged on logical grounds. Its negativity and significance strictly suggest that a weak price effect (3.2 per cent) occurred in 1969, but the timing is

⁷Income related differentials in prepayment arrangements and direct payments to hospitals are reported in C. Lloyd Francis, "Expenditure Patterns from the Canadian Sickness Survey, 1950-1951," Canadian Journal of Public Health, Vol. 47, August 1956, pp. 328-331.

implausible without evidence of an equal or greater transitory effect in 1968.⁸ Our tentative conclusion then, pending the results of the more disaggregate analyses to be reported in subsequent sections of this chapter, is that the copayment policy did not reduce or otherwise alter the representative family's chances of being hospitalized.

The key conclusion drawn from the foregoing analyses -- that price and income elasticities of hospitalization are effectively zero -- emerge with equal clarity from Model 2, as the reader may verify by scanning the results reported in the second column of the Table. Comments on these results will therefore be confined to assessing the consequences of employing continuous, as opposed to categorical, measures of family size and age of head, that is, of specifying these characteristics in such a way as to suppress the non-linearities shown to be important in Model 1. Noteworthy among these effects is the 10 per cent reduction in R^2 from 0.113 to 0.101, since it provides one approximation to the information loss attending this decision. An alternative, more direct way of gauging the "price" paid by assuming linearity is to compare the coefficients attached to the size and age variables in Model 2 with their generic counterparts in Model 1. Thus the size coefficient in Model 2 implies that an additional family member (child) raises the likelihood of hospitalization by some 6 per cent irrespective of initial family size, while the first difference of the size dummies in Model 1 indicate this effect ranges all the way from 3 per cent for the second child to 16 per cent for the first child. In

⁸The estimated arc elasticity of hospitalization based on the D69 coefficient is -0.01, where the coinsurance range is zero to approximately seven per cent.

general, the non-linearity is sufficiently pronounced that Model 2 understates the probability of use for small to medium size families but overstates it for large families. Similarly poor approximations are provided by the age coefficient in Model 2. It implies that yearly increments in the age of the family head increase the probability of use by 0.08 per cent; this figure must of course be inflated by a factor of 10 (i.e., to 0.8 per cent) to facilitate comparisons with the Model 1 coefficients, since the latter relate to dummy variables defined for the 10 year age intervals, 25-34, 35-44, etc. Even with this adjustment, however, the coefficient falls well short of capturing the non-linearity evident in the Model 1 results. It obscures not only the inverted U-shaped relationship over the age range 18-54, but also the linearly positive relationship over the 55-64, 65-74 and 75+ age groups; put another way, it misses the "child-bearing effect," the "healthy middle age effect," and the "unhealthy old age effect" so clearly demonstrated by the Model 1 coefficients.

It bears emphasizing that these distortions do not extend beyond the variables in question. Examination of the Table shows the coefficients of the other variables in Model 2 to be quite close to their respective values in Model 1, and hence largely insensitive to the specification of the family size and age of head regressors. Marginal differences are noted for the coefficients of SINGLE MALE and SINGLE FEMALE; but not to the extent that reversals in sign or significance occur. Moreover, the coefficients of the copayment (D68-71) variables exhibit an identical pattern in the two models. From the standpoint then of our main concern with price and income effects, there is little choosing between Model 1 and Model 2.

Pursuant to this concern, Model 3 further refines the probability function to test for the presence of interaction effects between copayment and selected family characteristics. In particular, it explores the possibility that price effects are concentrated within certain family size, age, and income groups, i.e., that the assumption implicit in Models 1 and 2, of equal response to user charges across demographically and economically different household classes is inappropriate. This is facilitated by the so-called "jack-knifing" technique whereby the continuously defined size, age, and income variables are jointly classified with the yearly dummies D68-71 to form the three sets of multiplicative (interactive) variables: D68.S, ..., D71.S; D68.A, ..., D71.A; and D68.Y, ..., D71.Y respectively. The coefficients of these variables show the extent to which the net hospitalization-size, age, and income relations differ in each of the years 1968-71 from their base period (1966-67) values given by the coefficients of the singly-classified S, A, and Y variables. On the assumption that all factors except copayment were constant throughout the six years, they may thus form the basis for arguments about price effects peculiar to certain family types. An additional possibility not previously allowed for is non-linearity in the cross-sectional probability of use-income relationship. This is accomplished by segmenting the income observations for each year according to the welfare status of the family, and specifically, by including in the equation the composite variable W.Y and D68.W.Y, ..., D71.W.Y. The coefficients of these variables provide the basis for inferences about the (in)equality of income effects between welfare ("poor") and non-welfare ("non-poor") families in the pre-copayment period 1966-67, and

in each of the copayment years 1968-71.

As the results make clear, the gains from these analytical extensions are negligible. The goodness of fit of Model 3 is virtually identical to that of Model 2 and inferior to that of Model 1 even though seventeen new variables have been added to the equation. Allowance for temporal variation in the size coefficient yields results formally consistent with the presence of an interaction effect between copayment and family size, but the magnitudes of the $D_1.S$ coefficients are extremely small. They imply that the size coefficient decreased from +0.06227 in 1966-67 to +0.05869 (i.e., $0.06227 - 0.00358$) in 1968, +0.05371 in 1969, +0.05841 in 1970, and +0.05778 in 1971, or, in general, that only minor weakening occurred in the size-use relationship during the copayment years,⁹ certainly not enough to warrant inferences about large families being more affected by user charges than small families. Nor is there any compelling evidence of price effects varying systematically with the age of the family head. Three of the four $D_1.A$ coefficients are statistically significant and large relative to the 1966-67 age coefficient of +0.00055, but they still yield very flat age-use relationships. The implied age coefficients for 1969, 1970, and 1971 are only +0.00123, +0.00124, and +0.00099 respectively. As for the relationship between family income and probability of hospitalization, simultaneous adjustments for welfare status and copayment make no difference. In only one instance does the income coefficient depart significantly ($p < .05$) from its "reference" (i.e., 1966-67, non-welfare

⁹Evaluated at the means of the variables, the estimated elasticity of hospitalization with respect to family size declined marginally from +0.61 (1966-67) to +0.57 (1968), +0.53 (1969), +0.55 (1970), and +0.54 (1971).

sub-sample) value of -0.00336 , and that is in 1969 when it assumes the value of -0.00182 (i.e., $-0.0336 + 0.00154$). But neither this nor the reference value is substantively different from the small, negative income coefficient common to Models 1 and 2, so the hypothesis of temporal instability cum price-income interaction must be informally rejected.

Probabilities of Hospitalization by Diagnostic
and Treatment Category.

Central to the analysis of the preceding section was a "broad" definition of hospitalization that made no reference to patient, provider, or institutional characteristics. Examined in this fashion, the data failed to indicate any marked reduction in the admission probability in consequence of copayment. Before accepting this as a firm conclusion, it is desirable to experiment with "narrow" definitions of hospitalization to check for behavioral responses to copayment that may have netted-out in the aggregation of heterogeneous case types. One dimension along which diverse price effects may in theory occur is that of medical diagnosis. Indications for admission vary considerably across different illnesses and conditions, not only in terms of strict medical criteria (the technical imperatives), but also in terms of the implicit weight attached to extra-medical criteria such as "convenience" to patient and/or physician. A second dimension relates to whether patients have been admitted for medical treatment or surgical treatment of their conditions. Accordingly, the primary purpose of this section is to determine whether the "inconvenience value" of copayment is

detectably different among families whose members experience different medical problems and different treatment of those problems.

To test for this, the data matrix has been refined to permit separate analysis of certain groups of disease entities classified according to diagnostic information routinely written on the admission-discharge form. The diagnostic categories selected for analysis are: diseases of the circulatory system; diseases of the respiratory system; diseases of the digestive system; diseases of the genito-urinary system; deliveries and complications of pregnancy, childbirth, and the puerperium; and accidents, poisonings and violence. The treatment categories are medical and surgical. No presumption is made that all cases within each of these categories are perfectly homogeneous with respect to severity, complexity, or other medical criteria; to do so would be unrealistic given the large number of different diseases that may be identified within each category. However, it is contended that this breakdown serves to eliminate or at least reduce those gross differences in health states that may be relevant to judgments on the copayment question.

Thus we introduce eight new diagnostic and treatment definitions of the hospitalization event, and correspondingly, eight new binary dependent variables, BINCIR, BINRES, BINDIG, BINGUR, BINDEL, BINA⁴PV, BINMED, and BINSUR. Estimates of these dichotomies of the sample families are presented in Tables 4.2 through 4.9 respectively. Many of the same comments may be made about these equations as have been made in the previous section for the all discharges (BINDIS) equation. What follows is not a detailed table by table, model by model reporting of coefficients, but a discussion intended to orient the reader to certain

TABLE 4.2

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINCIR
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

<u>Probability of Hospitalization</u>			
Independent Variables	Model 1	Model 2	Model 3
SIZE (S)		0.00167 (5.96)	0.00166 (3.95)
D68.S			0.00095* (1.38)
D69.S			-0.00018* (-.26)
D70.S			-0.00020* (-.29)
D71.S			-0.00071* (-1.00)
C1	0.00384 (2.70)		
C2	0.00412 (2.75)		
C3	0.01069 (6.07)		
C4	0.01149 (5.45)		
C5	0.01874 (8.56)		
AGE (A)		0.00120 (60.00)	0.00123 (30.75)
D68.A			-0.00015 (-2.14)
D69.A			0.00002* (.29)
D70.A			-0.00003* (-.43)
D71.A			-0.00002* (-.29)
A1	-0.00155* (-1.00)		
A2	0.00689 (4.18)		
A3	0.01566 (10.17)		
A4	0.02994 (19.44)		
A5	0.05396 (30.66)		
A6	0.08209 (39.66)		

TABLE 4.2 - Continued

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3.
SINGLE MALE	-0.00920 (-7.24)	-0.00453 (-3.46)	-0.00464 (-3.52)
SINGLE FEMALE	-0.01500 (-10.87)	-0.00981 (-6.91)	-0.00995 (-6.96)
URBAN	-0.00844 (-9.59)	-0.00752 (-8.55)	-0.00748 (-8.50)
WELFARE	0.01550 (6.83)	0.01649 (7.26)	0.00836 (2.65)
INCOME	-0.00009* (-1.00)	-0.00034 (-3.78)	-0.00050 (-2.94)
W.Y			0.00226 (2.09)
D68.Y			0.00045* (1.67)
D69.Y			-0.00002* (-.07)
D70.Y			0.00032* (1.23)
D71.Y			-0.00001* (-.04)
D68.W.Y			0.00165* (1.05)
D69.W.Y			-0.00198* (-1.50)
D70.W.Y			0.00121* (.77)
D71.W.Y			0.00059* (.43)
D68	-0.00057* (-.46)	-0.00058* (-.46)	0.00047* (.12)
D69	-0.00068* (-.54)	-0.00051* (-.41)	-0.00031* (-.08)
D70	-0.00072* (-.57)	-0.00059* (-.47)	-0.00073* (-.18)
D71	-0.00082* (-.66)	-0.00054* (-.43)	0.00245* (.61)
constant	0.01459	-0.02278	-0.02305
R ²	0.02128	0.01960	0.01983
S.E.E.	0.16186	0.16199	0.16198
F	163.850	273.887	108.878

*Insignificant at 95% Confidence Level

TABLE 4.3

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINRES
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE (S)		0.03020 (70.23)	0.03342 (53.05)
D68.S			-0.00227 (-2.20)
D69.S			-0.00746 (-7.17)
D70.S			-0.00605 (-5.71)
D71.S			-0.00446 (-4.21)
C1	0.04253 (19.97)		
C2	0.08603 (38.24)		
C3	0.11365 (43.21)		
C4	0.13740 (43.48)		
C5	0.17207 (52.46)		
AGE (A)		-0.00036 (-9.00)	-0.00048 (-6.86)
D68.A			0.00002* (.18)
D69.A			0.00019* (1.73)
D70.A			0.00017* (1.55)
D71.A			0.00033 (3.00)
A1	0.02249 (9.69)		
A2	-0.01326 (-5.35)		
A3	-0.02867 (-12.41)		
A4	-0.01723 (-7.49)		
A5	0.00192* (.73)		
A6	0.01849 (5.95)		

TABLE 4.3 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.02104 (-11.07)	-0.00013*	-0.00031*
SINGLE FEMALE	-0.01625 (-7.85)	0.00599 (2.80)	0.00545 (2.55)
URBAN	-0.02511 (-19.02)	-0.02262 (-17.14)	-0.02274 (-17.23)
WELFARE	0.03749 (11.03)	0.03834 (11.28)	0.03797 (8.01)
INCOME	-0.00093 (-6.64)	-0.00112 (-8.00)	-0.00149 (-5.96)
W.Y			-0.00081* (-.50)
D68.Y			0.00011* (.27)
D69.Y			0.00098 (2.45)
D70.Y			0.00079 (1.98)
D71.Y			0.00030* (.79)
D68.W.Y			0.00250* (1.06)
D69.W.Y			0.0021* (.11)
D70.W.Y			0.0114* (.48)
D71.W.Y			0.00247 (1.20)
D68	-0.00244* (-1.30)	-0.00294 (-1.56)	0.00190* (.32)
D69	-0.00945 (-5.05)	-0.00911 (-4.87)	-0.00161* (-.27)
D70	-0.00414 (-2.19)	-0.00406 (-2.15)	0.00108* (.18)
D71	-0.00606 (-3.22)	-0.00613 (-3.26)	-0.00990* (-1.63)
constant	0.05470	0.01631	0.01475
R ²	0.05794	0.05361	0.05414
S.E.E.	0.24267	0.24322	0.24316
F	463.487	776.095	308.098

*Insignificant at 95% Confidence Level

TABLE 4.4

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINDIG
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE (S)		0.01282 (31.27)	0.01323 (22.05)
D68.S			-0.00230 (-2.32)
D69.S			-0.00053* (-.54)
D70.S			0.00123* (1.22)
D71.S			-0.00104* (-1.03)
C1	0.02619 (12.9)		
C2	0.03154 (14.67)		
C3	0.04211 (16.71)		
C4	0.05730 (18.97)		
C5	0.07990 (25.53)		
AGE (A)		0.00038 (9.5)	0.00031 (5.17)
D68.A			0.00003* (.30)
D69.A			0.00025 (2.50)
D70.A			0.00017* (1.70)
D71.A			-0.00003* (-.30)
A1	0.00567 (2.55)		
A2	-0.00049* (-.21)		
A3	0.00236* (1.07)		
A4	0.01016 (4.62)		
A5	0.02365 (9.35)		
A6	0.03043 (10.25)		

TABLE 4.4 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.03191 (-17.53)	-0.02199 (-11.70)	-0.02226 (-11.84)
SINGLE FEMALE	-0.03068 (-15.49)	-0.02018 (-9.89)	-0.02034 (-9.97)
URBAN	-0.02368 (-18.79)	-0.02282 (-18.26)	-0.02276 (-18.06)
WELFARE	0.03397 (10.45)	0.03469 (10.67)	0.02734 (6.04)
INCOME	-0.00064 (-4.92)	-0.00078 (-6.00)	-0.00090 (-3.75)
D68.Y			0.00046* (1.18)
D69.Y			0.00002* (.05)
D70.Y			-0.00012* (-.32)
D71.Y			0.00019* (.53)
D68.W.Y			0.00648 (2.89)
D69.W.Y			-0.00060* (-.32)
D70.W.Y			0.00013* (.06)
D71.W.Y			0.00163* (.83)
D68	-0.00232* (-1.30)	-0.00254* (-1.42)	-0.00056* (-.10)
D69	-0.00010* (-.06)	0.00003* (-.02)	-0.00932* (-1.62)
D70	0.00295* (1.63)	0.00294* (1.62)	-0.00710* (-1.21)
D71	-0.00168* (-.93)	-0.00172* (-.96)	0.00112* (.19)
constant	0.05732	0.02639	0.02931
R ²	0.02141	0.02131	0.02153
S.E.E.	0.23199	0.23200	0.23198
F	164.896	298.405	118.512

*Insignificant at 95% Confidence Level

TABLE 4.5

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINGUR
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.00783 (23.03)	0.00729 (14.58)
D68.S			0.00034* (.41)
D69.S			-0.00050* (-.61)
D70.S			0.00310 (3.73)
D71.S			0.00046* (.55)
C1	0.01410 (8.39)		
C2	0.02126 (11.94)		
C3	0.02410 (11.59)		
C4	0.03001 (12.00)		
C5	0.04268 (16.48)		
AGE (A)		0.00039 (13.00)	0.00039 (7.80)
D68.A			-0.00005* (-.56)
D69.A			0.00007* (.88)
D70.A			0.00008* (.89)
D71.A			-0.00010* (-1.25)
A1	0.00531 (2.90)		
A2	0.00974 (4.97)		
A3	0.01588 (8.68)		
A4	0.01667 (9.16)		
A5	0.02013 (9.63)		
A6	0.02496 (10.15)		

TABLE 4.5 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.02531 (-16.87)	-0.02018 (-12.94)	-0.02023 (-12.97)
SINGLE FEMALE	-0.02005 (-12.23)	-0.01473 (-8.72)	-0.01471 (-8.70)
URBAN	-0.01134 (-10.90)	-0.01125 (-10.82)	-0.01123 (-10.80)
WELFARE	0.01027 (3.82)	0.01011 (3.76)	0.00844 (2.26)
INCOME	-0.00018* (-1.64)	-0.00013* (-1.18)	-0.00009* (-.45)
W.Y			0.00068* (.53)
D68.Y			-0.00020* (-.63)
D69.Y			0.00026* (.81)
D70.Y			-0.00036* (-1.16)
D71.Y			0.00005* (.17)
D68.W.Y			0.00126* (.68)
D69.W.Y			-0.00159* (-1.01)
D70.W.Y			-0.00060* (-.32)
D71.W.Y			-0.00007* (-.04)
D68	0.00255* (1.72)	0.00240* (1.62)	0.00433* (.91)
D69	0.0084* (1.24)	0.00183* (1.24)	-0.00105* (-.22)
D70	0.00479 (3.19)	0.00469 (3.15)	-0.00540* (-1.11)
D71	0.00421 (2.83)	0.00401 (2.69)	0.00667* (1.40)
constant	0.02856	0.00899	0.01034
R ²	0.01425	0.01422	0.01439
S.E.E.	0.01913	0.19193	0.19192
F	108.901	197.608	78.571

*Insignificant at 95% Confidence Level

TABLE 4.6

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINDEL
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.01203 (31.66)	0.01606 (28.68)
D68.S			-0.00506 (-5.50)
D69.S			-0.00443 (-4.82)
D70.S			-0.00758 (-8.06)
D71.S			-0.00884 (-9.31)
C1	0.08341 (44.6)		
C2	0.04782 (24.15)		
C3	0.04729 (20.47)		
C4	0.06198 (22.29)		
C5	0.10220 (35.49)		
AGE (A)		-0.00235 (-78.33)	-0.00243 (-40.50)
D68.A			-0.00016* (-1.60)
D69.A			0.00016* (1.60)
D70.A			0.00033 (3.30)
D71.A			0.00012* (1.33)
A1	0.04144 (20.62)		
A2	-0.05882 (-27.11)		
A3	-0.09985 (-49.19)		
A4	-0.09268 (-45.88)		
A5	-0.07709 (-33.23)		
A6	-0.06679 (-24.47)		

TABLE 4.6 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.05799 (-34.72)	-0.06252 (-35.73)	-0.06334 (-35.99)
SINGLE FEMALE	-0.03741 (-20.67)	-0.03993 (-21.02)	-0.04081 (-21.48)
URBAN	-0.00699 (-6.08)	-0.00301 (-2.57)	-0.00319 (-2.73)
WELFARE	0.02692 (9.03)	0.03370 (11.12)	0.04045 (9.59)
INCOME	-0.00082 (-6.83)	-0.00122 (-10.17)	-0.00181 (-8.23)
W.Y			-0.00061* (-.42)
D68.Y			0.00082 (2.28)
D69.Y			0.00087 (2.42)
D70.Y			0.00107 (3.06)
D71.Y			0.00089 (2.62)
D68.W.Y			-0.00283* (1.35)
D69.W.Y			-0.00153* (-.86)
D70.W.Y			0.00073* (.35)
D71.W.Y			-0.00172* (-.95)
D68	-0.00033* (-.20)	-0.00073* (-.44)	0.01646* (3.08)
D69	-0.00411 (-2.51)	-0.00365 (-2.19)	-0.00257* (-.48)
D70	-0.00744 (-4.48)	-0.00699 (-4.16)	-0.00616* (-1.13)
D71	-0.00517 (-3.13)	-0.00467 (-2.78)	0.00996* (1.85)
constant	0.09521	0.15014	0.14557
R ²	0.09288	0.06401	0.06497
S.E.E.	0.21298	0.21633	0.21623
F	771.542	936.985	373.982

*Insignificant at 95% Confidence Level

TABLE 4.7

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINAPV
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.01541 (45.32)	0.01539 (30.78)
D68.S			0.00087* (1.06)
D69.S			-0.00203 (-2.48)
D70.S			-0.00052* (-.63)
D71.S			0.00212 (2.52)
C1	0.01720 (10.24)		
C2	0.03294 (18.51)		
C3	0.04571 (21.98)		
C4	0.06410 (25.64)		
C5	0.10193 (39.36)		
AGE (A)		0.00017 (5.67)	0.00014 (2.80)
D68.A			-0.00008* (-.89)
D69.A			0.00010* (1.11)
D70.A			0.00013* (1.44)
D71.A			0.00006* (.75)
A1	-0.00270* (-1.48)		
A2	-0.01062 (-5.42)		
A3	-0.00129* (-.70)		
A4	0.00202* (1.11)		
A5	0.00195* (.93)		
A6	0.01672 (6.80)		

TABLE 4.7 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.00299 (-1.99)	0.01257 (8.06)	0.01271 (8.15)
SINGLE FEMALE	-0.00817 (-4.98)	0.00789 (4.67)	0.00787 (4.66)
URBAN	-0.01487 (-14.30)	-0.01454 (-13.98)	-0.01461 (-14.05)
WELFARE	0.01952 (7.26)	0.01960 (7.29)	0.02361 (6.30)
INCOME	-0.00053 (-4.82)	-0.00070 (-6.36)	-0.00099 (-4.95)
W.Y			0.00051* (.40)
D68.Y			0.00025* (.78)
D69.Y			0.00055* (1.72)
D70.Y			0.00059* (1.90)
D71.Y			0.00031* (1.03)
D68.W.Y			-0.00333* (-1.79)
D69.W.Y			-0.00059* (-.38)
D70.W.Y			-0.00174* (-.93)
D71.W.Y			-0.00316* (-1.95)
D68	0.00038* (.26)	0.00012* (.08)	0.00016* (.03)
D69	0.00186* (1.26)	0.00194* (1.31)	-0.00001* (-0.00)
D70	0.00442 (2.95)	0.00435 (2.90)	-0.00293* (-.60)
D71	0.00526 (3.53)	0.00516 (3.46)	-0.00448* (-.94)
constant	0.02993	-0.00719	-0.00423
R ²	0.02044	0.02026	0.02051
S.E.E.	0.19202	0.19203	0.19202
F	157.213	283.312	112.703

*Insignificant at 95% Confidence Level

TABLE 4.8

Parameter Estimates for Three Specifications
of the Linear Probability Function with
Dichotomous Dependent Variable, BINSUR
150,729 Positive Income Families, Pooled 1966-72 Sample Data
(t Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.03750 (64.66)	0.03806 (44.26)
D68.S			-0.00076* (-.54)
D69.S			-0.00393 (-2.75)
D70.S			-0.00048* (-.33)
D71.S			0.00166* (1.14)
C1	0.04891 (16.69)		
C2	0.09455 (30.40)		
C3	0.13500 (37.19)		
C4	0.16083 (36.89)		
C5	0.19839 (43.89)		
AGE (A)		0.00079 (16.8)	0.00071 (7.89)
D68.A			-0.00008* (-.53)
D69.A			0.00039 (2.60)
D70.A			0.00018* (1.20)
D71.A			-0.00002* (-.13)
A1	0.01964 (6.14)		
A2	0.02214 (6.47)		
A3	0.02811 (8.81)		
A4	0.04096 (12.88)		
A5	0.04895 (13.45)		
A6	0.05482 (12.78)		

TABLE 4.8 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.05647 (-21.55)	-0.02987 (-11.02)	-0.03001 (-11.03)
SINGLE FEMALE	-0.04987 (-17.50)	-0.02248 (-7.65)	-0.02277 (-7.74)
URBAN	-0.00171* (-.94)	-0.00092* (-.51)	-0.00092* (-.51)
WELFARE	0.01494 (3.19)	0.01358 (2.90)	0.01052* (1.61)
INCOME (Y)	-0.00004 (-.21)	0.00009 (.47)	0.00001* (.03)
W.Y			0.00153* (.68)
D68.Y			0.00018* (.32)
D69.Y			0.00041* (.75)
D70.Y			0.00023* (.43)
D71.Y			-0.00031* (-.60)
D68.W.Y			0.00321* (.99)
D69.W.Y			-0.00392* (-1.44)
D70.W.Y			-0.00108* (-.33)
D71.W.Y			-0.00066* (-.23)
D68	-0.00074* (-.29)	-0.00139* (-.54)	0.00300* (.36)
D69	-0.00472* (-1.83)	-0.00469* (-1.82)	-0.01258* (-1.52)
D70	0.00294* (1.13)	0.00255* (.98)	-0.00493* (-.58)
D71	0.00207* (.80)	0.00128* (.49)	0.00001* (0.00)
constant	0.08115	0.00676	0.00915
R ²	0.05381	0.05367	0.05385
S.E.E.	0.33474	0.33475	0.33474
F	428.545	777.015	306.331

*Insignificant at 95% Confidence Level

TABLE 4.9

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINMED
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE		0.03597 (58.97)	0.03719 (41.32)
D68.S			-0.00118* (-.80)
D69.S			-0.00404 (-2.73)
D70.S			-0.00105* (-.70)
D71.S			-0.00115* (-.76)
C1	0.07067 (23.25)		
C2	0.10480 (32.55)		
C3	0.13290 (35.35)		
C4	0.16767 (37.10)		
C5	0.23262 (49.71)		
AGE (A)		0.00183 (36.60)	0.00165 (18.33)
D68.A			-0.00010* (-.63)
D69.A			0.00027* (1.80)
D70.A			0.00039 (2.60)
D71.A			0.00047 (3.13)
A1	0.01672 (5.05)		
A2	-0.01050 (-2.97)		
A3	0.00585 (1.77)		
A4	0.03168 (9.63)		
A5	0.08983 (23.83)		
A6	0.17444 (39.29)		

TABLE 4.9 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.04986 (-18.40)	-0.01976 (-7.01)	-0.01979 (-7.02)
SINGLE FEMALE	-0.05393 (-18.28)	-0.02101 (-6.89)	-0.02133 (-6.97)
URBAN	-0.08415 (-44.76)	-0.08003 (-42.57)	-0.08011 (-42.61)
WELFARE (W)	0.10618 (21.85)	0.10975 (22.58)	0.10117 (14.92)
INCOME (Y)	-0.00225 (-11.25)	-0.00292 (-14.60)	-0.00354 (-10.11)
W.Y			0.00322* (1.38)
D68.Y			0.00080* (1.38)
D69.Y			0.00111* (1.95)
D70.Y			0.00125 (2.19)
D71.Y			0.00034* (.63)
D68.W.Y			0.00443* (1.32)
D69.W.Y			-0.00477* (-1.68)
D70.W.Y			-0.00088* (-.26)
D71.W.Y			-0.00003* (-.01)
D68	-0.00151 (-.54)	-0.00220 (-.82)	0.00050* (.06)
D69	-0.00336 (-1.26)	-0.00278 (-1.04)	-0.00880* (-1.02)
D70	0.00620 (2.30)	0.00642 (2.37)	-0.01453* (-1.65)
D71	0.00719 (2.67)	0.00751 (2.79)	-0.01155* (-1.34)
constant	0.13572	0.02519	0.03306*
R ²	0.07135	0.06669	0.06695
S.E.E	0.34665	0.34751	0.34748
F	578.957	979.105	386.183

*Insignificant at 95% Confidence Level

3

general patterns evident in the tables.

Note first that the explanatory power of the equations is very low. The R^2 values range from .014 in the three variants of the BINGUR equation (Table 4.5) to .136 in the Model 1 variant of the BINMED equation, (Table 4.8). Of these magnitudes the former is by far the more representative, since most of the other equations yield R^2 's in the very narrow range, .014 - .021. The reason for such poor fits in the diagnostic equations may be the relatively small number of families actually experiencing hospitalization and the concomitant multitude of zero observations on the dependent variables.¹⁰ At any rate, it is clearly a disaggregation effect since the diagnosis-unadjusted equation, BINDIS, yielded $R^2 = .11$ (see Table 4.1).

Insofar as the effects of the various predictor variables are concerned, the equations contain few surprises. The family size coefficients are uniformly positive and significant as expected under the simple "potential patients" hypothesis. Some interest attaches to the fact the Model 1 size coefficients, C_1 , generally indicate a much less pronounced non-linearity in the size-use relationship than was apparent in the BINDIS equation. Only in the deliveries (BINDEL) equation and to a lesser extent the digestive (BINDIG) equation does the "one child effect" dominate the relationship; in the other diagnostic categories, the C_1 coefficients and their first differences suggest near equality in the net effects of additional children over the observed range of family sizes. It follows that these incremental effects are

¹⁰"User" proportions calculated from the pooled 1966-71 sample data range from a low of .028 for circulatory conditions to a high of 0.067 for respiratory conditions. For further details, see Appendix Table D.7.

reasonably approximated by the coefficient of the linearly defined size variable S in the majority of Model 2 equations.

The distinguishing feature of the age-use relationship is its variability across the diagnostic categories. This is particularly evident in the Model 1 results, where frequent and often substantial cross-tabular differences occur in the sign, significance, and magnitude of the A_1 coefficients. Compare, for example, the circulatory and delivery equations in Tables 4.2 and 4.6 respectively. In the former, the age terms are all positive with the exception of A_1 (25-34) which is negative (and insignificant); in the latter, they are all negative with the exception of A_1 which is positive. The differences are of course understandable. Circulatory conditions encompass a variety of heart, vein, and artery diseases that are markedly age-related, being more likely in the old than the young. Obstetrical conditions, principally routine deliveries, are also age-related, but tend naturally enough to reach a peak in the 25-34 age group and decline rapidly thereafter. With respect to the other four diagnostic categories, there may be special epidemiological significance to the fact the A_1 coefficients tend to be very small and, in consequence, to yield very weak age-use relationships. Suffice it to say, there is evidence of linearity in the genito-urinary and digestive equations where the probability of hospitalization monotonically increases with age of head, and evidence of non-linearity in the respiratory and accidents equations where the likelihood is minimized in the 45-54 and 35-44 age groups respectively.

The marital status coefficients, M and F, are notable primarily for their negativity and significance, and secondarily for their magnitude similarities within and between the diagnostic categories.

They indicate that families headed by a single male or single female are on average 2-3 per cent less likely to make contact with the hospital system than those where both spouses are present. Again, in the absence of more discriminating data on the care-seeking behavior of one-spouse households, it can and should be argued that this is a simple size effect supplemental to and independent of that picked up by the coefficients on the number of children variables.

Two other characteristics shown by these equations to be empirically relevant predictors of hospitalization are families' place of residence and social assistance status. The coefficient of the location variable, U , is uniformly negative and significant (except for the BINSUR equation), indicating that urban families are less likely to be hospitalized for treatment of any of the specified conditions than rural families. This finding was a feature of the analysis in the previous section, and the interpretation there offered -- of it being a supply effect generated by (or at least consistent with) underlying differences in the bed-population and physician-population ratios between the province's rural and urban areas -- is equally appropriate here. For its part, the coefficient of the welfare variable, W , always carries a positive sign, suggesting that members of families receiving social assistance stand a relatively high chance of being hospitalized. This too invites recall of an earlier expressed view, namely, that the differential owes its existence to a complex combination of unfavourable social, economic and health factors unique to the disadvantaged minority.

In contrast to the variables discussed thus far, family income adds little or no explanatory power to the equations. This is not to

say there is no statistical relationship between the economic position of the family and the probability of hospitalization, but rather that the relationship is so weak as to be practically non-existent. To illustrate, Table 4.10 below draws together the Model 2 income coefficients and associated income elasticities for each of the diagnostic and treatment categories.

It is evident from the table and the t values in particular that the coefficients are, with two exceptions, substantially in excess of their standard errors, and hence formally consistent with the probability of hospitalization being a decreasing function of family income. However, their magnitudes make equally clear that the relationship is anything but strong. For example, in the deliveries equation where it is most pronounced, the coefficient implies that, all else equal, the female adult in a family with an income of \$10,000 is only 1.2 per cent less likely to experience admission for obstetrical reasons than her counterpart in a family with an income of \$1,000; evaluated at the means of the variables, the income elasticity of pregnancy-related admissions is only -0.13. Since income effects of a similarly diminutive nature are apparent in the other equations as well, the overall results may be taken as further evidence of the insurance plan's success in diluting the influence of income on access to inpatient care.

Finally, with respect to the key question of price effects, the equations yield results very much in line with those reported in the previous section. The dummy variables D68-D71 introduced to test for downward shifts in the probability functions during the copayment period 1968-70 more often than not feature coefficients that are either

TABLE 4.10

Family Income Coefficients and Elasticities by
Diagnostic & Treatment Category

Diagnostic Category	Income Coefficient (t Value)	Income Elasticity
Circulatory	-0.00034 (-3.78)	-0.07
Respiratory	-0.00112 (-8.00)	-0.09
Digestive	-0.00078 (-6.00)	-0.08
Genito-Urinary	-0.00013* (-1.18)	-0.02
Deliveries	-0.00122 (-10.17)	-0.13
Accidents	-0.00070 (-6.36)	-0.10
Medical	-0.00225 (-11.25)	-0.11
Surgical	-0.00004* (-.210)	0.0

*Insignificant (P<.05)

"wrongly" signed (i.e., positive) or statistically insignificant, or both. Only in the respiratory equations (Table 4.3) and the delivery equations (Table 4.6) are they predominantly negative and significant. The respiratory coefficients are intriguing. They relate to a diagnostic group that (unlike the others) is dominated by conditions -- acute respiratory infections, influenza, and bronchitis -- that, according to one clinical taxonomy, "do not generally require hospitalization".¹¹ The discretionary element necessary for an "extra-medical" factor like copayment to affect the admission decision is arguably greater in these cases than others, and hence the coefficients invite interpretation as a bona fide price effect. The estimated arc elasticity of hospitalization over the observed coinsurance range is -0.07 in 1969, -0.03 in 1970, and -0.04 in 1971, an inter-year pattern that points to a long-run price elasticity of zero. For their part, the delivery coefficients are clearly unrelated to copayment ("theory" or at least common sense excludes these sorts of prices from the conception function!). What they do reflect is the steady decline in the birth rate from 19.9 per 1000 in 1966 to 17.3 per 1000 in 1971¹². Thus, viewed over all diagnostic and treatment categories, the coefficients sustain the first impression that copayment did not substantially alter probabilities of admission.

¹¹Arnold V. Hurtado and Merwyn R. Greenlick, "A Disease Classification System for Analysis of Medical Care Utilization, with a Note on Symptom Classification," Health Services Research, Fall 1971, pp. 235-250.

¹²Saskatchewan Department of Public Health, Vital Statistics 1971 (Regina: Queen's Printer, 1973), Table 1, page 10.

Probabilities by Type of Practitioner.

In this, as in the previous section, the primary objective is to determine whether or not conclusions regarding the effect of copayment on the probability of hospitalization vary depending upon the definition of the hospitalization event chosen for analysis. Specifically, the question is whether the "no effect" conclusion tentatively drawn in section 1 from an analysis based upon the very general "admission to hospital" definition is also implied by analyses where the event is more narrowly defined in terms of patient, provider or institutional characteristics. Having just determined that adjustment for diagnosis and treatment makes little or no difference to judgments on the matter, it is now appropriate to consider whether the copayment effect is mediated by the practice setting of the admitting physician. The rationale for choosing "practice setting" and not some other attribute of the admitting physician has been given previously in Chapter 3. Recall it is hypothesized that probabilities of admission by solo general practitioners are more likely to decrease with copayment than are probabilities of admission by group general practitioners. The hypothesis clearly assumes a pre-copayment concentration of "unnecessary" or at least "questionable" admissions among solo practitioners, and a concomitant willingness and ability on their part to become more circumspect when advising hospitalization during the copayment period.

To determine if the data support this line of reasoning, two new binary dependent variables, BINS GP and BING GP, have been constructed. As their names suggest, the former defines the occurrence of the

hospitalization event as one involving admission (to, and subsequent discharge from, an acute care general hospital) by a solo general practitioner, while the latter defines it as admission by a group general practitioner.¹³ Estimates of the probability function for these two mutually exclusive dichotomies of the sample are presented in Table 4.11 and Table 4.12 respectively.

First among the general comments invited by these results concerns the overall explanatory power of the equations. It is, as the R^2 values indicate, generally low and in no instance close to the .11 level reached in the "aggregate" BINDIS equations reported in Section 1. The three BINGSP equations register R^2 values only slightly in excess of 0.03, while the three BINGGP equations do somewhat better, but still fail to get beyond the 0.07 mark. Again, we are obviously observing the consequences of a disaggregation strategy according to which the vast majority of families (92 per cent in the SGP dichotomy and 85 per cent in the GGP dichotomy) are allocated to the "not-hospitalized" category.

As for the superiority of the "group" over the "solo" equations, it is not immediately obvious why the various factors hypothesized to affect admission probabilities are collectively more relevant in one than the other. All that can be said with certainty is that the comparatively high "group" R^2 's have more to do with the explanatory

¹³Notice that both variables assume the value of zero for a family in which one or more members were admitted to hospital by a medical or surgical specialist, but not by a solo or group general practitioner. To accommodate this category of hospitalized family, a third "practitioner-specific" binary dependent variable, BINMSP, was constructed which defined the event as one involving admission by a medical or surgical specialist. However, since it relates only to specialty and not to practice setting it is of less analytical interest. The reader with a taste for "completeness" is referred to Appendix F where probability functions based upon it are reported.

TABLE 4.11

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINS GP
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.01821 (40.47)	0.01870 (27.91)
D68.S			-0.00249 (-2.26)
D69.S			-0.00332 (-3.02)
D70.S			0.00129* (1.15)
D71.S			0.00150* (1.33)
C1	0.04437 (19.63)		
C2	0.05396 (22.58)		
C3	0.06543 (23.37)		
C4	0.09019 (26.84)		
C5	0.11530 (33.16)		
AGE (A)		0.00034 (8.50)	0.00021 (7.00)
D68.A			-0.00006* (-.50)
D69.A			0.00020* (1.82)
D70.A			0.00039 (3.55)
D71.A			0.00017* (1.55)
A1	0.01253 (5.09)		
A2	-0.01123 (-4.27)		
A3	-0.00806 (-3.28)		
A4	0.00017* (.07)		
A5	0.02179 (7.75)		
A6	0.05561 (16.85)		

TABLE 4.11 - Continued

Independent Variable	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.03379 (-16.73)	-0.02176 (-10.41)	-0.02184 (-10.40)
SINGLE FEMALE	-0.02899 (-13.18)	-0.01541 (-6.79)	-0.01555 (-6.85)
URBAN	-0.04849 (-34.64)	-0.04659 (-33.28)	-0.04648 (-33.20)
WELFARE	0.02241 (6.21)	0.02454 (6.80)	0.0165 (3.29)
INCOME	-0.00150 (-10.00)	-0.00175 (-11.67)	-0.00152 (-5.85)
W.Y.			0.00238* (1.38)
D68.Y			0.00041* (.95)
D69.Y			0.00009* (.21)
D70.Y			-0.00083 (-1.98)
D71.Y			-0.00093 (-2.33)
D68.W.Y			0.00181* (.73)
D69.W.Y			-0.00204* (-.97)
D70.W.Y			0.00506 (2.01)
D71.W.Y			-0.00269* (-1.23)
D68	-0.00606 (-3.05)	-0.00645 (-3.24)	0.00070* (.11)
D69	-0.00237* (-1.19)	-0.00213* (-1.07)	-0.00211* (-.33)
D70	0.01163 (5.79)	0.01170 (5.82)	-0.00476* (-.73)
D71	0.00846 (4.23)	0.00852 (4.26)	0.00316* (.49)
constant	0.07963	0.04295	0.04577
R ²	0.03557	0.03279	0.03319
S.E.E.	0.25771	0.25807	0.25803
F	277.948	464.518	184.781

*Insignificant at 95% Confidence Level

TABLE 4.12

Parameter Estimates for Three Specifications
of the Linear Probability Function with
Dichotomous Dependent Variable, BINGGP
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.03745 (61.39)	0.04166 (46.29)
D68.S			-0.00361 (-2.44)
D69.S			-0.00613 (-4.14)
D70.S			-0.00938 (-6.21)
D71.S			-0.00770 (-5.07)
C1	0.09285 (30.54)		
C2	0.11268 (34.99)		
C3	0.14128 (37.57)		
C4	0.17710 (39.27)		
C5	0.23781 (50.81)		
AGE (A)		-0.00002* (-.40)	-0.00009* (-1.00)
D68.A			-0.00035 (-2.19)
D69.A			0.00030 (2.00)
D70.A			0.00013* (.87)
D71.A			0.00028* (1.87)
A1	0.02889 (8.73)		
A2	-0.03350 (-9.46)		
A3	-0.04264 (-12.92)		
A4	-0.02145 (-6.52)		
A5	0.01904 (5.05)		
A6	0.06758 (15.22)		

TABLE 4.12 - Continued

Independent Variable	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.07911 (-29.19)	-0.05603 (-19.87)	-0.05677 (-20.13)
SINGLE FEMALE	-0.06044 (-20.49)	-0.03444 (-11.25)	-0.03544 (-11.58)
URBAN	-0.07358 (-39.14)	-0.06920 (-36.81)	-0.06942 (-36.93)
WELFARE	0.07942 (16.38)	0.08396 (17.28)	0.09065 (13.37)
INCOME	-0.00226 (-11.30)	-0.00280 (-14.00)	-0.00356 (-10.17)
W.Y			-0.00331* (-1.42)
D68.Y			0.00039* (.67)
D69.Y			0.00066* (1.16)
D70.Y			0.00192 (3.37)
D71.Y			0.00148 (2.74)
D68.W.Y			0.00399* (1.19)
D69.W.Y			0.00114* (.40)
D70.W.Y			0.00009* (.03)
D71.W.Y			0.00380* (1.30)
D68	-0.00213* (-.79)	-0.00290* (-1.08)	0.02045 (2.38)
D69	-0.01067 (-4.00)	-0.01011 (-3.77)	-0.00950* (-1.10)
D70	-0.01607 (-5.95)	-0.01585 (-5.85)	-0.00591* (-1.67)
D71	-0.01007 (-3.76)	-0.00988 (-3.67)	-0.00992* (-1.15)
constant	0.17211	0.11522	0.11068
R ²	0.07296	0.06624	0.06675
S.E.E.	0.34645	0.34769	0.34762
F	593.089	972.023	384.959

*Insignificant at 95% Confidence Level

power of the sub-set of socio-demographic and economic variables than with the sub-set of copayment variables. Thus, the deletion of the copayment variables is no more consequential in terms of R^2 reductions to the group equations than it is to the solo equations.

Notice also that inter-model differences in goodness-of-fit are extremely small. Only seven-tenths of one per cent separate the "best" from the "worst" group equation, while only three-tenths of one per cent separate the "best" from the "worst" solo equations. In each instance the upper and lower limits to R^2 are associated with Model 1 and Model 2 respectively. This tendency, also noted in earlier sections, for Model 1 to slightly outperform the other two specifications is ascribable of course to its featuring non-linear age-use and size-use relationships. Empirically they are the most relevant of the non-linearities considered, as attested to in both tables by the failure of Model 3 with its lengthy list of interactive variables to measure up to Model 1.

Other points of tangency between these results and those reported in earlier sections are not difficult to find. Thus, the family composition variables play their familiar roles. Size of family, measured either continuously (SIZE) or categorically (C_1), exerts a strong positive influence on the likelihood of hospitalization. An interesting, though not surprising, result in this regard is the similarity of the size effect in both the solo and group equations; for example, both Model 1 variants imply that, ceteris paribus, a family with five or more children present is nearly two and one-half times more

likely to experience hospitalization in any given year than a family with no children present.¹⁴ Age of family head also figures importantly in both equations, particularly in the Model 1 variants where the A_1 coefficients preserve certain non-linearities peculiar to the age-use relationship. The coefficients of the marital status variables, M and F, carry negative signs indicating once more that absence of either spouse brings about a significant reduction in the probability of hospitalization. All in all there are no surprises here.

The same can be said of the coefficients on the variables describing families' place of residence (URBAN), social assistance status (WELFARE), and income position (INCOME). As expected, urban families are substantially less likely to be hospitalized than are rural families. The only new and interesting detail in this connection is the similarity of the rural/urban differential across "practice setting"; both the solo and group equations imply that, ceteris paribus, urban residence reduces the probability of hospitalization by roughly fifty per cent.¹⁵ This further reinforces the pure bed supply effect interpretation of the differential, and concomitantly rules out the possibility of it being a "practice setting effect" due to the

¹⁴For the solo equation, the ratio $P(H/5+ \text{ children}) / P(0 \text{ children}) = (0.080 + 0.115) / 0.080 = 2.4$; similarly, for the group equation the ratio $= (0.172 + 0.238) / 0.172 = 2.4$.

¹⁵This is readily seen in the context of the Model 1 results (though it is also implied by Models 2 and 3 as well) where the constant term defines, inter alia, the rural probability and the URBAN coefficient the urban adjustment factor. Expressing the latter in terms of the former yields 0.060 for the solo equation and 0.43 for the group equation.

relatively high (low) percentage of soloists among rural (urban) general practitioners.¹⁶

The social assistance variable is again consistently positive and significant, confirming that disadvantaged families in contact with social agencies are an identifiable "high risk" population. Another more specific result that emerges here is that such families are more likely to be admitted by group general practitioners than by solo general practitioners. The "welfare effect" in the (Model 1) group equation is proportionately much greater than it is in the (Model 1) solo equation as may be seen by deflating the WELFARE coefficient by the constant term in each equation. It is doubtful this reflects any conscious preference on the part of welfare patients for treatment by group practitioners. What it does suggest is that group practitioners work in closer proximity to referring social agencies than do solo practitioners. For its part, the income coefficient is small, negative, and significant in both the solo and group equations regardless of specification. Evaluated at the means, the implied income elasticities are -0.13 (solo) and -0.10 (group).

Finally, the data offer no support for the hypothesis that the hospitalization decisions of solo GPs are more likely to be affected by per-diem user charges on the patient than those of group GPs. The coefficient of the variables D68-D71 are predominantly small, positive, and insignificant in all three variants of the solo equation. In the group equations their pattern -- predominantly small, negative, and

¹⁶Over the study years 1966-71, soloists accounted for 39.3 per cent of rural general practitioners, but only 30.2 per cent of urban general practitioners. Saskatchewan, Medical Care Insurance Commission, Annual Report (Regina: Queen's Printer, various annual issues).

significant -- has the curious face value interpretation that a weak price effect occurred among group GPs, i.e., in the "wrong" practice setting, one which conventional wisdom associates with relatively conservative (price-inelastic) demands on inpatient facilities. However, further analysis of this result strongly suggests it has little to do with the copayment policy of 1968-71, and more to do with the steady decline which occurred over these years in the relative importance of group GPs in the province's active physician population.¹⁷ The observed decreases in the probability of hospitalization by a group GP (indicated by the negative coefficient on D68-71) are, on this reasoning, due not to any reduction in the propensity of group GPs to hospitalize their patients because of copayment, but to the decreased probability that a care-seeking family would come into contact with a group GP.¹⁸ Put another way, it is not a price effect in the wrong practice setting, but a supply effect in the right practice setting.

¹⁷Specifically, group GP representation fell from 44.2 per cent in 1966 to 37.9 per cent in 1971. Per cent figures for the other study years are 42.6 (1967), 39.1 (1968), 38.3 (1969), and 37.1 (1970). These calculations are based on counts of "resident active" physicians, i.e., physicians resident in the province (but not the Swift Current Health Region) at the end of the year who received \$10,000 or more in Medicare payments during the year. Saskatchewan, Medical Care Insurance Commission, Annual Reports (Regina: Queen's Printer, various issues).

¹⁸Empirical support for this statement is provided by the high correlation ($r=.96$) between the "contact probability" (given by the proportion of physicians in group practice, on the reasonable assumption that patients choose their physicians without reference to the organizational character of the practice) and the probability of hospitalization by a group GP (generated by solving the Model 1 variant of the group equation for the "reference" family, i.e., a rural couple with no children, headed by a male in the 18-24 age group earning an average income and receiving no social assistance).

Probabilities by Type of Hospital.

This section represents a third and final attempt to detect a copayment effect via a disaggregation strategy featuring alternative "narrow" definitions of hospitalization. Thusfar, this strategy has involved differentiating inpatients by medical diagnosis (Section 2), and admitting physician by practice organization (Section 3), and has yielded results generally consistent with those reported for a "broad" definition of hospitalization (Section 1). It remains now to determine whether a similar consistency is evident when the data are broken-down into sub-samples differentiating admissions according to the service capability of the host institution.

The logic of this exercise merits brief restatement prior to presenting and assessing the pertinent facts. Recall first that the concept of "service capability" serves, in the Saskatchewan context, to identify three distinct types of active treatment hospitals. In ascending order of service capability they are the district or community hospitals, the regional hospitals, and the base hospitals. Representative institutions in each category differ not only in terms of the range of diagnostic and therapeutic services offered, but also in terms of rated bed capacity, medical staff organization, and population served. The community hospital tends to be small (10-24 beds), "staffed" by one or two general practitioners, and populated by patients from the local area. In sharp contrast, the regional and base hospitals tend to be much bigger (150-300 beds and 300-550 beds respectively), with formally organized medical staffs consisting of both general

practitioners and specialists; the wide array of services offered by these hospitals results in their serving as referral centers for persons outside the local area.

From these scale and organizational distinctions a number of "performance" propositions can be readily derived, the most important of which for present purposes is the one associating the "low-technology" community hospital with a relatively liberal admissions policy, and "medium-to-high" technology regional and base hospitals with a relatively conservative admissions policy. If this is the case, that is, if indications for admission to a community hospital are more influenced by "extra-medical" criteria, then it follows that copayment may have a greater impact on these institutions than their organizationally complex counterparts. This is the principal hypothesis under test in this section.

The binary dependent variables constructed to effect the test are BINCOM and BINREB. The former distinguishes families who experience admission to a community hospital from those who do not, while the latter distinguishes families who experience admission to a regional or base hospital from those who do not. Equations estimated for these two dichotomies of the data set are presented in Tables 4.13 and 4.14.

Proceeding directly to the rows of these tables where the coefficients of D68-71 are reported, it is evident that the data provide some tantalizing support for the above hypothesis. The contrasting sign patterns -- essentially negative in the community equations, essentially positive in the regional and base equations -- are particularly striking. They indicate that probabilities of admission to small, rural

TABLE 4.13

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINCOM
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t-Values in Parentheses)

Independent Variables	<u>Probability of Hospitalization</u>		
	Model 1	Model 2	Model 3
SIZE (S)		0.03455 (61.70)	0.03718 (44.80)
D68.S			-0.00273 (-2.01)
D69.S			-0.00605 (-4.45)
D70.S			-0.00414 (-2.98)
D71.S			-0.00362 (-2.59)
C1	0.07404 (26.54)		
C2	0.09573 (32.34)		
C3	0.11976 (34.61)		
C4	0.15976 (38.50)		
C5	0.22148 (51.51)		
AGE (A)		0.00035 (7.00)	0.00015* (1.67)
D68.A			-0.00013* (-.93)
D69.A			0.00036 (2.57)
D70.A			0.00041 (2.93)
D71.A			0.00050 (3.57)
A1	0.01731 (5.69)		
A2	-0.02458 (-7.56)		
A3	-0.02238 (-7.36)		
A4	-0.01217 (-4.03)		
A5	0.02624 (7.56)		
A6	0.07889 (19.29)		

TABLE 4.13 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.07531 (-30.24)	-0.04923 (-19.01)	-0.04946 (-19.02)
SINGLE FEMALE	-0.04108 (-15.10)	-0.01264 (-4.50)	-0.01320 (-4.70)
URBAN	-0.21926 (-126.74)	-0.21579 (-124.73)	-0.21591 (-124.80)
WELFARE	0.05450 (12.19)	0.05773 (12.91)	0.06421 (131)
INCOME	-0.00173 (-9.61)	-0.00223 (-12.39)	-0.00244 (-7.39)
W.Y			-0.00449 (-2.10)
D68.Y			0.00055* (1.02)
D69.Y			0.00035* (.66)
D70.Y			0.00053* (1.02)
D71.Y			0.00009* (.18)
D68.W.Y			0.00837 (2.71)
D69.W.Y			0.00477* (1.83)
D70.W.Y			0.00364* (1.17)
D71.W.Y			0.00022* (.08)
D68	-0.00628 (-2.55)	-0.00694 (-2.81)	0.00258* (.33)
D69	-0.00643 (-2.61)	-0.00595 (-2.42)	-0.00715* (-.91)
D70	-0.00144* (-.58)	-0.00129* (-.52)	-0.01096* (-1.36)
D71	-0.00105* (-.43)	-0.00092* (-.37)	-0.01310* (-1.65)
constant	0.21607	0.14757	0.15035
R ²	0.15385	0.14987	0.15024
S.E.E.	0.31875	0.31949	0.31944
F	1370.144	2415.464	951.549

*Insignificant at 95% Confidence Level

TABLE 4.14

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINREG
150,729 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE (S)		0.03173 (49.58)	0.03235 (34.41)
D68.S			-0.00163* (-1.05)
D69.S			-0.00265* (-1.70)
D70.S			0.00039* (.25)
D71.S			-0.00014* (-.09)
C1	0.09690 (30.38)		
C2	0.11227 (33.22)		
C3	0.13317 (33.71)		
C4	0.15541 (32.79)		
C5	0.18236 (37.14)		
A		0.00065 (10.83)	0.00059 (5.90)
D68.A			-0.00032* (-2.00)
D69.A			0.00038 (2.38)
D70.A			0.00034 (2.13)
D71.A			-0.00003* (-.19)
A1	0.04302 (12.36)		
A2	-0.00147* (-.40)		
A3	-0.00948 (-2.73)		
A4	0.02034 (5.90)		
A5	0.05644 (14.25)		
A6	0.10336 (22.18)		

TABLE 4.14 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.07033 (-24.68)	-0.05958 (-20.13)	-0.06005 (-20.22)
SINGLE FEMALE	-0.08506 (-27.44)	-0.07144 (-22.26)	-0.07095 (-22.41)
URBAN	0.12518 (63.54)	0.12904 (65.50)	0.12899 (65.15)
WELFARE	0.06649 (13.04)	0.07086 (13.87)	0.06356 (8.93)
INCOME	-0.00079 (-3.76)	-0.00098 (-4.67)	-0.00151 (-4.08)
W.Y			0.00492 (2.02)
D68.Y			0.00046* (.75)
D69.Y			0.00104* (1.73)
D70.Y			0.00094* (1.59)
D71.Y			0.00050* (.88)
D68.W.Y			0.00039* (.11)
D69.W.Y			-0.00992* (-3.33)
D70.W.Y			-0.00492* (-1.38)
D71.W.Y			0.00031* (.10)
D68	0.00607 (2.16)	0.00529* (1.88)	0.02125 (2.35)
D69	-0.00143* (-.51)	-0.00109* (-.39)	-0.01481* (-1.64)
D70	0.00793 (2.80)	0.00793 (2.79)	-0.01278* (-1.38)
D71	0.00845 (3.00)	0.00829 (2.93)	0.00715* (6.79)
constant	0.06723	0.02038	0.02435
R ²	0.07634	0.06992	0.07020
S.E.E.	0.363	0.036514	0.36510
F	622.835	1030.026	406.345

*Insignificant at 95% Confidence Level

hospitals were indeed lower during each of the copayment years than they were during 1966-67,¹⁹ and also that no similar effect occurred among the larger urban hospitals. However, as in previous instances where statistically significant negative coefficients have appeared, the magnitude of the reductions is quite small. Using the Model 1 results, the implied arc price elasticity of hospitalization is only -0.02 in both 1968 and 1969.²⁰ This calculation, it is well to remember, applies only to the observed coinsurance range of 0-7 per cent,²¹ but even when extrapolated to the wider, and in some (U.S.) contexts, more "realistic" range of 0-25 per cent, the estimate is at -0.07 still demonstrably small. Moreover, knowing (from D70 and D71) that the probabilities had by the last half of the copayment period returned to their 1966-67 levels, the conclusion of a non-zero long-run price elasticity is difficult to sustain.

Other notable results contained in the Tables may now be summarized. First, the community equations are markedly superior to the regional and base equations in "goodness of fit" terms. Indeed, the community R^2 's is 0.15 (cf. the regional base R^2 's of 0.07) are higher

¹⁹The stock of community hospital beds per 1000 rural residents was marginally higher in the copayment years (5.6 - 5.7 beds per 1000) than in 1966-67 (5.5 beds per 1000), so an availability effect cannot be implicated in this finding.

²⁰An alternative, and arguably superior measure of the price effect is given by the ratio of the D69 (or D68) coefficient and the constant term. It indicates that the probability of admission in 1969 (or 1968) was 2.7 per cent below its pre-copayment level.

²¹In 1969 the average public (S.H.S.P.) cost per adult and child patient-day was \$33.78, hence copayment of \$2.50 per day represented a coinsurance rate of 6.9 per cent (i.e., $\$2.50 / \$33.78 + \$2.50$). Saskatchewan Hospital Services Plan, Annual Report 1971 (Regina: Queen's Printer, 1972), Table VII, p. 30.

than any reported in previous sections. This could be seen as indirect support for speculation that "extra-medical" factors are more important in determining the utilization of small rural hospitals than the utilization of large urban hospitals, but only in very blurred circumstances. Close inspection of the stepwise results (not reported in the tables) reveals that over sixty per cent of the explained variance in BINCOM is ascribable to one variable, namely the dummy variable (URBAN) identifying urban families. This immediately suggests that the result is an artifact of the classification scheme, since members of urban families are not normally "cross-hauled" to rural hospitals (indeed, according to Model 1, the probability of this occurring is virtually zero). Thus, the presence of the urban variable in these equations is arguably a specification error, albeit of the minor variety, since the coefficients of the other variables are not detectably affected by its inclusion. It should also be pointed out that the "zero-cell" problem does not arise in the regional and base equations for the simple reason that rural patients are frequently referred to urban hospitals for diagnosis and treatment of serious medical problems. This is the directionally logical "cross-haul" that is captured by the positive constant term in these equations (recall the "omitted" family has rural residence).²²

Second, the family composition variables generally perform as expected. In both the community equations and the regional and base

²²Insofar as urban families are "obviously" more likely than rural families to be admitted to urban hospitals, there is still a certain artifactual quality to the urban coefficient. It indicates the incremental effect to be 12.5 - 13.0 per cent, depending upon specification.

equations, family size is an important predictor of hospitalization. Interestingly, however, the non-linearity that has thusfar characterized this relationship is evident only in the BINREB equations. There the linearity assumption of Model 2 slightly overstates the probability of admission for a couple with no children, and substantially understates it for all others. In contrast, there is no choosing between Models 1 and 2 in the BINCOM equations. They yield virtually identical estimates over the observed range of family sizes, and, as such, provide the first real evidence of a linear size-use relation. These patterns have methodological significance, but for substantive content they do not rival the Model 3 results, especially those for BINCOM where a significant interaction between copayment and family size is detected. Solving this equation for year specific elasticities of hospitalization with respect to family size yields a pattern -- +0.73 (1966-67), +0.70 (1968), +0.64 (1969), +0.66 (1970-71) -- indicating slight temporal weakening in the relationship consistent with the inference that large families were among those somewhat affected (i.e., in terms of admission probabilities) by copayment.²³

The other descriptors of family composition, age, and marital status of head, invite only brief comment. Age is a non-linear correlate of admission (to both small and large general hospitals) as evidenced by the alternating signs on the A1-A6 coefficients in Model 1. This "life-cycle" effect is lost on the continuous age variable in Model

²³To better appreciate the magnitude involved, notice, for example, that prior to copayment, a family of six (two adults, four children) stood a 73 per cent greater chance of hospitalization than a family of three (two adults, one child), whereas in 1969 the proportional difference in probabilities had decreased to 64 per cent.

2 which substantially overstates the probability of hospitalization for members of families with a head in the age ranges 18-24 years and 35-64 years, and substantially understates it for those in the age ranges 25-34 years, 65-74 years, and 75+ years. For their part, the marital status variables, SINGLE MALE and SINGLE FEMALE, carry relatively large negative coefficients, confirming again that single males and single females are, ceterus paribus, less likely to experience admission than married couples. In previous sections this has been interpreted as a "scale effect" (as opposed to a difference in health status of care-seeking behavior) and the view applies equally here.

Finally, there can be no mistaking the differential importance of income and welfare effects in these, as in previous results. The income effects are negligible. Among families not in receipt of social assistance payments, variations in annual income as great as \$10,000 are associated with variations in the admission probability of no more than two per cent. The estimated income elasticities are only -0.07 and -0.03 for the (Model 1) community equations and the regional and base equations respectively. Of much more consequence is the discontinuity in the "income-use" relation picked up by the welfare variable (WELFARE) in each of the equations. It indicates (as it has in previous sections) that families at the extreme low end of the income scale face a very different set of hospitalization probabilities than those in the unassisted middle and upper income ranges. Specifically, the coefficients of +0.05 and +0.07 reported in Tables 13 and 14 put the incremental welfare effect at fully 20-25 per cent of the corresponding non-welfare probability. The nexus between low income and poor health is doubtless a complex one, but the data presented here certainly

reinforce the view held by many health and social scientists that other public initiatives complementary to the existing hospital and medical care insurance plans are necessary to improve the health of the poor.

Summary

Given the quite extensive series of statistical experiments that have been reported and discussed in this chapter, it is well at this juncture to provide a short inventory of the major findings before proceeding to the next phase of the analysis. Thus:

1. There is no compelling evidence that probabilities of "broadly defined" hospitalization are reduced by the imposition of per diem user charges. A few of the "narrowly defined" hospitalization probabilities display some price sensitivity but the effects are generally weak and transitory, and hence, suggestive of a long-run money price elasticity equal to zero. Associated with this key finding is the policy implication that copayment is a tax on, rather than a deterrent to, hospital admission.

2. There is substantial evidence indicating that probabilities of hospitalization are strongly and systematically related to family size and composition. This edifies and further reinforces the traditional view that the demand for in-hospital care varies over the "life cycle" of the family.

3. The likelihood of hospitalization is markedly higher among families resident in the province's rural areas where active treatment beds are relatively abundant and where physicians are relatively scarce. This pattern is consistent with commonly assumed "availability effects,"

but does not constitute a direct test of the "pure" availability effect implied by the so-called "Say's Law" of hospital beds.

4. Families at various income levels above the "poverty" line (stringently defined by social assistance status) face virtually identical probabilities of admission. The income elasticity of hospitalization is effectively zero, and, as such, is consistent with the "theory" of hospital use under conditions of universal, comprehensive health insurance.

5. "Poor" families (i.e., those in receipt of social assistance) are an identifiable high risk group. The likelihood of their experiencing hospitalization (at least once) during any given year is much greater than it is for families located elsewhere in the income distribution. Public universal health insurance is evidently a necessary, but insufficient, means of solving the complex health and social problems presented by the seriously disadvantaged.

CHAPTER V

REGRESSION RESULTS: THE LINEAR UTILIZATION EQUATIONS

The analysis of admission probabilities presented in the previous chapter constitutes the first step in a two-step statistical procedure designed to ascertain the effect of copayment and other factors on families' use of hospital services. The second step, to be reported in this chapter, involves an analysis of the actual quantities of care received by "user" families. Specifically, linear utilization functions with continuous dependent variables, alternatively defined as the annual number of patient-days and separations, are fitted to subsets of observations restricted to those families whose members experienced some hospitalization during the study period. These equations generate estimates of the expected volume of patient-days and separations per year for families with certain socio-demographic and economic characteristics, given that they are users. These conditional utilization estimates are related to years when copayment did and did not apply in order to determine the price elasticity of days and separations respectively. Equally important, they together provide an indirect means of gauging the effect of the per diem user charges on the duration of hospital stays, and, hence, a test of the key hypothesis that discharge decisions are more price sensitive than are admission decisions.

To preserve the essential complementarity of the first and second stage estimations, the results will be presented in an order identical to that of the previous chapter. Thus, in section one,

patient-days, separations, and lengths of stays are "broadly" defined without reference to the diagnostic or treatment status of the patient, to practice organization of the admitting physician, or to size and service capability of the hospital. Results specific to these alternative "narrow" definitions of use are then analyzed in sections two, three, and four respectively. In each section, three forms (Model 1, Model 2, Model 3) of ~~one~~ utilization functions are estimated in order, as before, to test different assumptions about the nature (linear vs. non-linear) and stability (pre- and post-copayment) of the relationship between patient-days, separations, and duration of stay on the one hand, and selected demographic and economic family characteristics on the other. A short summary of the major findings is presented at the end of the Chapter.

Total Use Equations.

The families whose hospitalization experiences are analyzed in this section comprise the largest single group of "users" in the sample data. Included are all families wherein at least one member was admitted to (and discharged from) an acute care general hospital on at least one occasion during one or more of the study years 1966-71. Under this "broad" definition of hospitalization, no allowance is made for inter-patient differences in diagnosis and treatment, inter-physician differences in practice organization, or inter-hospital differences in size and service capability. Ordinary least squares estimates of the

coefficients of the "total use" equations are presented in Table 5.1.¹ The column headings in this table distinguish the three variants (Model 1, Model 2, and Model 3) of the equation employing patient-days per family per year, DAYDIS, as the dependent variable from three variants of the equation employing admissions per family per year, ADMDIS, as the dependent variable.

At the outset, it is well to emphasize that these equations provide only a partial explanation for the fact that families consume different quantities of hospital care. The specified vector of independent variables, comprised mainly of families' socio-demographic and economic characteristics, explains only 8 per cent of inter-family variation in patient-days per year, and only 5 per cent of inter-family variation in admission per year. This result is neither surprising nor disquieting. As mentioned before, large unexplained or residual variances are to be expected in microanalytic circumstances such as these where the family constitutes the elemental unit of observation, and where the behavioral differences being observed are characterized by a fundamental unpredictability well known to actuarial and social scientists alike. This is not to say that "better fits" are impossible to obtain, only that such improvements as might attend further refinements are likely to be marginal.² Moreover, the explanatory power of the equations is, in this analysis, more a matter of procedural than

¹Corresponding estimates for slightly broader definitions of utilization (DAYDTX, ADMDTX) are presented in Appendix F.

²There would, for example, be a substantial return to more complete information on the environmental and life-style factors that affect families' health, and, thereby, the risks of recurrent and prolonged hospitalization.

TABLE 5.1

Parameter Estimates for Three Specifications
of the Linear Utilization Function With Alternate
Continuous Dependent Variables DAYDIS and ADMDIS
43,761 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYDIS per Year Given Hospitalization			ADMDIS per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		0.05395* (1.14)	0.11549* (1.56)		0.09979 (30.42)	0.11220 (22.04)
D68.S			-0.13248* (-1.05)			-0.03853 (-4.44)
D69.S			-0.11759* (-.90)			-0.01090* (-1.21)
D70.S			-0.12930* (-.99)			-0.01249* (-1.39)
D71.S			-0.00782* (-.06)			-0.01311* (-1.46)
C1	-0.06043* (-.23)			0.10067 (5.46)		
C2	-0.63623 (-2.26)			0.20233 (10.42)		
C3	-0.36526* (-1.16)			0.28115 (12.98)		
C4	0.44301* (1.25)			0.40203 (16.47)		
C5	1.97451 (5.60)			0.68117 (28.00)		
AGE (A)		0.26305 (51.08)	0.26311 (28.94)		0.00134 (3.72)	0.00070* (1.11)
D68.A			0.00149* (.10)			-0.00168* (-1.57)
D69.A			0.00653* (.43)			0.00133* (1.25)
D70.A			-0.00904* (-.59)			0.00248 (2.34)
D71.A			-0.00006* (-.00)			0.00179* (1.72)
A1	1.78147 (5.21)			0.07884 (3.34)		
A2	2.55929 (7.09)			-0.06207 (-2.49)		
A3	4.24114 (12.12)			-0.08506 (-3.52)		
A4	7.84145 (22.09)			-0.01865* (-.76)		
A5	11.88346 (30.03)			0.10425 (3.82)		
A6	17.13525 (39.92)			0.23354 (7.88)		

TABLE 5.1 - Continued

Independent Variables	DAYDIS Per Year Given Hospitalization			ADMDIS Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	3.14323 (10.43)	3.84711 (12.6)	3.84250 (12.58)	-0.05392 (-2.59)	0.06498 (3.09)	0.06304 (2.99)
SINGLE FEMALE	-1.20301 (-3.60)	-0.55096* (-1.63)	-0.55052* (-1.62)	-0.23220 (-10.06)	-0.11537 (-4.94)	-0.11998 (-5.13)
URBAN	0.69834 (4.02)	0.91722 (5.28)	0.91426 (5.26)	-0.22313 (-18.59)	-0.20971 (-17.49)	-0.21139 (-17.63)
WELFARE	6.50102 (16.42)	6.64275 (16.71)	6.33163 (11.38)	0.60169 (22.02)	0.59963 (21.92)	0.67115 (17.18)
INCOME	-0.10794 (-6.25)	-0.14879 (-8.70)	-0.06551 (-5.47)	-0.00699 (-5.87)	-0.00907 (-7.69)	-0.01202 (-5.75)
W.Y			0.24889* (1.45)			-0.01197* (-1.01)
D68.Y			0.03244* (.63)			0.00297* (.84)
D69.Y			0.04005* (.81)			0.00798 (2.33)
D70.Y			-0.00208* (-.04)			0.00567* (1.66)
D71.Y			0.01391* (.29)			0.00356* (1.08)
D68.W.Y			-0.63043 (-2.54)			-0.03189 (-1.86)
D69.W.Y			-0.16699* (-.69)			-0.00433* (-.26)
D70.W.Y			-0.09147* (-.36)			-0.01515* (-.87)
D71.W.Y			-0.05018* (-.23)			0.00312* (.21)
D68	-0.45425* (-1.87)	-0.41299* (-1.70)	-0.07894* (-.08)	-0.02967* (-1.77)	-0.02759* (-1.65)	0.17735 (2.66)
D69	-0.68096 (-2.79)	-0.65128 (-2.66)	-0.72850* (-.73)	-0.02596* (-1.54)	-0.02231* (-1.32)	-0.08823* (-1.29)
D70	-0.93984 (-3.86)	-0.93790 (-3.85)	-0.01948* (-.02)	-0.01737* (-1.03)	-0.01478* (-.88)	-0.11210* (-1.61)
D71	-1.43017 (-5.89)	-1.39973 (-5.75)	-1.43457* (-1.45)	-0.03110* (-1.85)	-0.02800* (-1.67)	-0.08225* (-1.21)
constant	8.28034	1.14274	0.98383	1.53564	1.29591	1.29414
R ²	0.08095	0.07661	0.07687	0.05291	0.04960	0.05069
S.E.E.	16.91134	16.94948	16.95043	1.16721	1.16913	1.16868
F	192.638	329.983	130.055	122.187	207.58	83.404

*Insignificant at 95% Confidence Level

substantive concern. It is useful in appraising the merits of alternative specifications and as a quick check on the significance of the overall model, but not much else. Far more important are the coefficients attached to the constituent variables. They test the various principal and secondary hypotheses at issue in the study, and it is to their performance that the discussion now turns.

Consider first the Model 1 results for both the patient-days and admission equations. The constant term in each of these equations refers to the same "standard" family as the constant term in the Model 1 probability functions analyzed in the previous chapter. Thus, 8.28 days and 1.54 admissions are the point estimates of use in either 1966 or 1967 for a rural couple with head aged 18-24 years, no children, and no social assistance benefits. The implied average length of stay for the members of this family is the ratio of these two figures, or 5.4 days. The coefficients on the dummy variables D68-D71, and the dummy variables describing number of children (C1,...,C5), age of head (A1,...,A6), absence of spouse (M,F), social assistance (W), and urban residence (U), are the amounts by which patient-days and admissions must be increased or decreased in order to derive point estimates for subsequent years and/or for families different in one or more respects from those included in the reference group.

The family size and composition variables carry coefficients that are notable for their diverse signs and magnitudes. The C_1 coefficients indicate that the incremental effect of additional children on days and admissions is very weak. Only when five or more children are present does the day total increase, and then by only two days. Over the same range, the case count rises by only .68 to a total of 2.12

admissions (i.e., $1.54 + .68$). Because the latter is proportionately greater than the former, there is an implied decrease in average length of stay of 2.2 days, fully 40 per cent of the above noted 5.4 days average for the reference family. These findings are of course consistent with the assumption that children are relatively healthy, and, except for certain acute respiratory infections, rarely require hospitalization. The age coefficients tell a quite different story. In the days equation they display a pattern that confirms age as an important predictor of use, especially in families where the head is aged 55 years or older. The three "senior" age coefficients, A4, A5, A6 (corresponding to the 55-64, 65-74, and 75+ age groups respectively), indicate they spend from 1 to 2.5 weeks longer in hospital than their counterparts in the 18-24 age group. In the admissions equation, the pattern is quite different in that the coefficients turn negative at the age group 35-44, and remain so until the age group 65-74 when they become positive again. This bears a certain faint resemblance to the "life cycle" effect illustrated in the BINDIS equation of the previous chapter,³ but the actual variation in case frequency is very limited. It follows that the strong positive relationship between age and patient-days is mainly a length of stay effect. This is confirmed by the A₁ coefficients in the implicit average length of stay equation. Their pattern — +.72 (A1), +1.91 (A2), +3.08 (A3), +5.01 (A4), and +7.97 (A5) — clearly demonstrates that duration is much more important than frequency in explaining the patient-day differences between "young" and "old" families. A similar comment applies in respect of the marital

³See page 152 for a discussion.

status coefficients (SINGLE MALE, SINGLE FEMALE). Single, widowed, or divorced males and their dependents spend 3.1 more days in hospital, not because they are admitted more often, but because they stay an average of 2.2 days longer than their counterparts in families where both spouses are present. This finding is of particular interest, since it suggests that discharge is delayed when the patient's domestic resources are limited in this way. Organized home care programs can reduce, if not eliminate, these deficits, facilitate earlier discharge, and thereby contribute significantly to system-wide efficiency.

The coefficients for the location variable (URBAN) indicate that the use patterns vary between rural and urban families. Urban families are admitted less frequently, but they tend to stay an average of 1.6 days longer, so their patient-day total is marginally higher. Recalling that the effective bed-population ratio is lower in the urban areas than in rural areas, these results offer further clarification of the relationship between bed availability and bed use. Under the simplest availability hypothesis, both case volume and average length of stay are expected to be lower where beds are relatively scarce. The present findings confirm the case effect, but not the length of stay effect. Case mix and case severity differences may explain this result. If urban hospitals attract a disproportionate number of patients with heart and cerebrovascular disease, the long stay is understandable. Subsequent analysis controlling for diagnosis will allow an investigation of this phenomenon although the possibility of case severity effects will still remain. In any event, an availability effect could still be present since the shorter stays observed among rural patients could reflect liberal admissions policies induced by the

existence of unfilled beds.

The welfare coefficients indicate that families on social assistance experience (6.5) more days, (.6) more admissions, and (1.1 days) longer average lengths of stay than their unassisted counterparts. The income coefficients confirm the presence of an inverse relationship between family income and utilization, though the implied elasticities are very small. Estimated at the relevant means, the income elasticities of volume, frequency, and duration are only -0.07, -0.03, and -0.04 respectively.

The evidence on price effects among user families is given by the coefficients on the variables D68-D71. It will be noted that in the volume (DAYDIS) equation, three of these coefficients are negative and statistically significant, while in the frequency (ADMDIS) equation they are all uniformly insignificant. This suggests that user families experienced fewer patient-days in 1969, 1970, and 1971 than in the non-copayment years 1966-67 and, moreover, that these volume reductions were largely in consequence of shorter average lengths of stay. This is made apparent when estimates of -0.4, -0.5, and -0.7 are obtained for the parameters of the D69, D70, and D71 variables respectively in the average length of stay equation. These results clearly support the main hypothesis that the discharge decision is more price sensitive than the admission decision.⁴

⁴The estimated arc price elasticities for length of stay are -0.02 in 1969, -0.03 in 1970, and -0.05 in 1971; corresponding elasticities for admission are -0.008, -0.005, and -0.01. It should be noted that since these calculations define the price change in terms of the average of the money prices before (zero) and after (\$2.50) copayment, they are necessarily lower than those which would result if the average were to include some arbitrary time price per day of use. The ambiguity of the arc price elasticity in analytical circumstances

A useful summary measure of the deterrent effect is obtained when the D69 and D70 coefficients in the DAYDIS equation are averaged and related to the mean number of patient-days per family in the pre-copayment period. Thus a decrease in 1969-70 of .81 days from a mean in 1966-67 of 13.05 days yields a deterrent effect of approximately 6.2 per cent. Since aggregate utilization in the pre-copayment years was 1.866 million days, this implies a reduction of 115,700 days for the province as a whole. Pricing this deterrent effect out at 1966-67 costs of \$27.62 per day, implies a resource cost saving of approximately \$3.2 million.

There are however several reasons for not accepting the copayment coefficients at their face value. First, in both the volume and duration equations the D71 coefficient is much larger in magnitude than the D69 and D70 coefficients. This, coupled with the insignificance of the corresponding coefficients in the frequency equation, implies that discharge decisions were more price sensitive in 1971 than in 1969 or 1970. But this is paradoxical since it is reasonable to assume that copayment would have its greatest impact in the years immediately following its introduction due to the "shock-effect" on clinical decision-making. Moreover, since copayment was eliminated in August, 1971, one would expect a proportionately smaller price effect in that year than in 1969 and 1970 when copayment was in force for the full twelve month period. Second, the copayment variables are time related, and will therefore pick up non-price effects peculiar

similar to these is illustrated in Charles E. Phelps and Joseph P. Newhouse, "Effect of Coinsurance: A Multivariate Analysis," op. cit., p. 23.

to any or all of the copayment years. One such effect is the steady decline in utilization over the sample period, a trend which in fact began in 1964.⁵ In 1966 the average user family experienced 13.06 patient days, 1.63 admissions and an average length of stay of 8.02. By 1971 these figures had declined to 12.17, 1.57 and 7.74 respectively. In order to assess the extent to which the copayment policy can be implicated in this trend, it is instructive to compare the utilization of a group of socially assisted families who were exempt from copayment with the utilization of a group of socially unassisted families who were not exempt. Data facilitating these comparisons are presented in Table 5.2 below.

TABLE 5.2

Indices of Inpatient Care Per Family for
Welfare and Non-Welfare Groups, 1966-71 Sample Data

	Patient-Days		Separations		Length of Stay	
	Welfare	Non-Welfare	Welfare	Non-Welfare	Welfare	Non-Welfare
1966	100.0	100.0	100.0	100.0	100.0	100.0
1967	102.8	98.8	114.4	98.8	90.0	100.0
1968	91.6	96.0	109.9	95.7	83.3	100.0
1969	94.9	96.4	110.9	95.7	85.8	100.8
1970	83.6	96.4	105.0	96.3	79.9	100.0
1971	82.8	92.0	104.4	94.4	79.3	96.9

⁵Population data documenting this trend are shown in Appendix Table A.1.

It is evident that both groups of families experienced less hospitalization in the copayment years (1968-71) than they did in the pre-copayment years (1966-67). Equally inconsistent with the operation of a price effect is the fact that the indices of patient-days and length of stay for welfare families decreased more than the corresponding indices for non-welfare families. In the light of this information it is arguable that some factor other than copayment accounts for the negativity and significance of the copayment coefficients in the volume and duration equations, and that accordingly they ought not to be interpreted as bona fide price effects.

This judgment is sustained by the Model 2 estimates, but this is not surprising for, as the reader will recall, Model 2 is similar to Model 1 in all respects save for the substitution of continuous for categorical size and age variables. The Model 2 size coefficient is insignificant, reflecting the non-linear effect of additional children in the Model 1 equation. The linearity approximation is better in the admission equation than in the days equation, as the reader may verify by comparing the relevant equations. For its part, the age-use relationship is more satisfactorily approximated by the linear specification in the days equation than in the admissions equation. However, in no case do the linear specifications affect the coefficients D68-D71.

The Model 3 specification further refines the copayment analysis by explicitly allowing for interaction between price and selected family characteristics. In general, no new effects of any consequence are evident. The size-copayment and age-copayment interaction terms are uniformly insignificant (with the exception of D68*SIZE and D70*AGE in

the admission equation) and small in absolute value. In addition, income-copayment and welfare-copayment interactors fail to capture a differential effect. It may be concluded therefore that copayment provisions did not alter the behavioral relationships between use on the one hand, and income, age, family size, and welfare on the other hand.

Use By Diagnostic and Treatment Category

In this section, the first of several "narrow" definitions of use is considered. The fundamental question here is whether disaggregations based on diagnostic and treatment information reveal a pattern of copayment effects consistent with the above noted decline in broadly defined use, and consistent with expectations concerning the price sensitivity of selected diagnostic groups. The dependent variables in this analysis are the continuous analogues of the binary variables employed in section two of the previous chapter, namely DAYCIR, ADCIR, DAYRES, ADMRES, DAYDIG, ADMDIG, DAYGUR, ADMGUR, DAYDEL, ADMDEL, DAYAPV, ADMAV, DAYSUR, ADMSUR, DAYMED, and ADMMED. Utilization functions estimated with these variables as regressands are reported in Tables 5.3 to 5.10.

Some overall comments on the goodness of fit of these equations are warranted. In general, Model 1 slightly outperforms the other two models in the days equation and slightly underperforms the other two models in the admissions equations. The R^2 values for days range from 0.01 in the DEL equation to 0.14 in the GUR equation while the range for admissions is from 0.01 for DEL to 0.05 for RES. The low explanatory power of the equations is by now a common and recurrent property of the

TABLE 5.3

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYCIR and ADMCIR
4,147 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYCIR Per Year Given Hospitalization			ADMCIR Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.28090* (-1.43)	-0.33016* (-1.05)		-0.00946 (-1.37)	0.01478* (1.33)
D68.S			-0.27009* (.52)			0.00628 (.35)
D69.S			0.17152* (.30)			0.02392* (1.20)
D70.S			-0.27462* (-.50)			0.00229* (.12)
D71.S			0.77298* (1.43)			0.00350* (.18)
C1	-0.68775* (-.71)			0.00961* (.28)		
C2	-0.02743* (-.02)			-0.02545* (-.64)		
C3	-0.68820* (.55)			0.01171 (.27)		
C4	1.19464* (.82)			0.01339* (.26)		
C5	-1.91660* (-14.34)			-0.03544* (-.73)		
AGE (A)		0.23804 (10.71)	0.22648 (5.60)		0.00441 (5.65)	0.00518 (3.62)
D68.A			0.02065* (.31)			-0.00121* (-.51)
D69.A			-0.00329* (-.05)			0.00176* (.72)
D70.A			0.00451* (.07)			-0.00217* (-.93)
D71.A			0.04663* (.72)			-0.00197* (-.86)
A1	1.82873* (.69)			-0.00821* (-.09)		
A2	2.20167* (.87)			-0.03247* (-.36)		
A3	6.46533 (2.64)			0.02327* (.27)		
A4	9.76661 (4.08)			0.08732* (1.03)		
A5	9.90480 (4.13)			0.16308* (1.92)		
A6	13.44618 (5.58)			0.18788 (2.21)		

TABLE 5.3 - Continued

Independent Variables	DAYCIR Per Year Given Hospitalization			ADMCIR Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	3.66153 (4.00)	3.13645 (3.33)	3.20309 (3.39)	0.06486 (2.01)	0.05759* (1.73)	0.05849* (1.75)
SINGLE FEMALE	2.30850 (2.15)	1.76012* (1.60)	1.93477* (1.75)	-0.04060* (-1.07)	-0.05117* (-1.32)	-0.05091* (-1.30)
URBAN	3.26835 (5.13)	3.25806 (5.12)	3.29746 (5.17)	-0.06431 (-2.86)	-0.06374 (-2.84)	-0.06283 (-2.79)
WELFARE	5.08170 (3.95)	4.90458 (3.82)	5.32901 (2.97)	0.24570 (5.41)	0.24827 (5.48)	0.30975 (4.89)
INCOME	-0.09355* (-1.65)	-0.08662* (-1.54)	-0.05011* (-.45)	-0.00032* (-.16)	-0.00064* (-.32)	0.00091* (.23)
W.Y			-0.17033* (-.37)			-0.01776* (-1.10)
D68.Y			-0.02567* (-.16)			-0.00496* (-.90)
D69.Y			-0.21292* (-1.17)			-0.00504* (-.79)
D70.Y			0.03744* (.23)			0.00313* (.53)
D71.Y			-0.06746* (-.37)			0.00010* (.02)
D68.W.Y			-0.76340* (-1.13)			-0.01820* (-.76)
D69.W.Y			0.15576* (.19)			-0.01177* (-.40)
D70.W.Y			0.66342* (1.14)			0.02615* (1.27)
D71.W.Y			0.20739* (.37)			0.01382* (.71)
D68	-1.80394 (-2.06)	-1.82111 (-2.09)	-1.78787 (-.35)	-0.04525* (-1.46)	-0.04596* (-1.49)	0.03613* (.20)
D69	-2.06047 (-2.37)	-2.11495 (-2.43)	-1.35347* (-.25)	-0.03084* (-1.00)	-0.03106* (-1.01)	-0.17889* (-.94)
D70	-3.10002 (-3.56)	-3.09496 (-3.56)	-2.90386* (-.57)	-0.05097* (-1.66)	-0.05015* (-1.63)	0.04504* (.25)
D71	-2.64730 (-3.06)	-2.71655 (-3.15)	-7.29996* (-1.45)	-0.02949* (-.96)	-0.03047* (-1.00)	0.06817* (.38)
constant	8.20217	3.22712	3.81993	1.16443	1.02183	0.98656
R ²	0.07068	0.06987	0.07236	0.02906	0.02805	0.03107
S.E.E.	18.61646	18.60427	18.61769	0.65725	0.65687	0.65720
F	15.690	28.238	11.472	6.175	10.85	4.716

*Insignificant at 95% Confidence Level

TABLE 5.4

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYRES and ADMRES
10,733 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYRES Per Year Given Hospitalization			ADMRES Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.19172 (-4.05)	-0.11786* (-1.65)		0.05015 (13.37)	0.06417 (11.34)
D68.S			-0.13347* (-1.06)			-0.02224 (-2.23)
D69.S			-0.16941* (-1.32)			-0.02502 (-2.46)
D70.S			-0.13707* (-1.06)			-0.02090 (-2.03)
D71.S			-0.01108* (-.09)			-0.02104 (-2.04)
C1	-0.69570 (-2.04)			0.03382* (1.24)		
C2	-0.73860 (-2.17)			0.10774 (3.95)		
C3	-1.05457 (-2.94)			0.15495 (5.40)		
C4	-0.37876* (-.98)			0.22126 (7.19)		
C5	-0.10270* (-.27)			0.34293 (11.26)		
AGE (A)		0.10865 (18.38)	0.11414 (11.38)		-0.00204 (-4.34)	-0.00132* (-1.65)
D68.A			-0.00545* (-.32)			-0.00291 (-2.12)
D69.A			-0.03656* (-2.08)			-0.00062* (-.45)
D70.A			-0.01321* (-.75)			-0.00066* (-.47)
D71.A			0.02207* (1.26)			0.00011* (.08)
A1	0.20237* (.54)			0.00204* (.07)		
A2	0.03162* (.08)			-0.07349 (-2.36)		
A3	0.28400* (.73)			-0.13416 (-4.29)		
A4	2.04884 (4.82)			-0.09804 (-2.94)		
A5	5.01270 (10.41)			-0.06155* (-1.60)		
A6	8.29427 (15.91)			-0.02939* (-.71)		

TABLE 5.4 - Continued

Independent Variables	DAYRES Per Year Given Hospitalization			ADMRES Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	1.97566 (5.75)	2.54914 (7.46)	2.58164 (7.54)	-0.02727*	0.03504*	0.03754*
SINGLE FEMALE	-0.88013 (-2.22)	-0.34189* (-.87)	-0.31310* (-.79)	-0.12491 (-3.95)	-0.06531 (-2.08)	-0.06750 (-2.15)
URBAN	-0.91879 (-5.04)	-0.80494 (-4.40)	-0.79007 (-4.31)	-0.12235 (-8.40)	-0.11764 (-8.09)	-0.11642 (-8.00)
WELFARE	4.01646 (10.65)	4.11996 (10.88)	3.26378 (5.82)	0.23413 (7.77)	0.23173 (7.71)	0.18409 (4.14)
INCOME	-0.04574 (-2.43)	-0.07514 (-4.00)	-0.07141 (-2.34)	-0.00196* (-1.30)	-0.00265* (-1.78)	-0.00505 (-2.09)
W.Y			-0.23294* (-1.38)			-0.01257* (-.94)
D68.Y			0.02165* (.39)			0.00814* (1.83)
D69.Y			-0.02993* (-.55)			0.00067* (.15)
D70.Y			-0.03774* (-.71)			-0.00150 (-.36)
D71.Y			-0.00623* (-.12)			0.00540* (1.31)
D68.W.Y			0.39540* (1.69)			0.00339* (.18)
D69.W.Y			0.81321 (3.38)			0.05362* (2.81)
D70.W.Y			0.60638 (2.39)			-0.00278* (-.14)
D71.W.Y			0.66452 (3.25)			0.06888 (4.25)
D68	0.01753* (.07)	0.04577* (.19)	0.69030* (.66)	-0.00104* (-.05)	-0.00027* (-.01)	0.17141 (2.07)
D69	0.09155* (.36)	0.10449* (.41)	2.42044 (2.24)	-0.01015 (-.51)	-0.01077* (-.54)	0.11178* (1.30)
D70	-0.07594* (-.31)	-0.12916* (-.52)	1.17102* (1.08)	-0.00043* (-.02)	-0.00213* (-.11)	0.13008* (1.51)
D71	-0.41299* (-1.64)	-0.43276* (-1.70)	-1.41745* (-1.29)	-0.03316* (-1.64)	-0.03482* (-1.73)	-0.00054* (-.01)
constant	6.67544	3.40329	2.86939	1.24875	1.18899	1.11264
R ²	0.09152	0.07520	0.07799	0.04703	0.04311	0.04803
S.E.E.	8.25572	8.32581	8.32026	0.66018	0.66124	0.66009
F	50.733	74.525	30.405	24.851	41.29	18.132

*Insignificant at 95% Confidence Level

TABLE 5.5

Parameter Estimates for Three Specifications
of the Linear Utilization Function With Alternate
Continuous Dependent Variables DAYDIG and ADMDIG
8,806 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYDIG Per Year Given Hospitalization			ADMDIG Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.22647 (-3.78)	-0.20906 (-2.26)		.01249 (4.00)	0.01495 (3.10)
D68.S			0.00553* (.03)			-0.00693* (-.84)
D69.S			-0.14930* (-.91)			-0.00263* (-.31)
D70.S			0.07046* (.43)			-0.00527* (-.62)
D71.S			-0.02599* (-.16)			-0.00076* (-.09)
C1	-1.28947 (-3.68)			-0.01045* (-.57)		
C2	-2.10732 (-5.70)			-0.00570* (-.30)		
C3	-2.19535 (-5.38)			-0.01236* (-.58)		
C4	-2.11652 (-4.70)			0.00927* (.39)		
C5	-1.88540 (-4.25)			0.06228 (2.69)		
AGE (A)		0.13699 (20.00)	0.14661 (12.27)		0.00072 (2.00)	0.0017 * (.27)
D68.A			-0.00204* (-.10)			0.00112* (1.06)
D69.A			-0.00948* (-.47)			0.00044* (.42)
D70.A			-0.01342* (-.66)			0.00073* (.69)
D71.A			-0.03353* (-1.63)			0.00088* (.82)
A1	2.68691 (5.64)			0.06255 (2.52)		
A2	3.41896 (7.02)			0.03306* (1.30)		
A3	4.74000 (10.10)			0.05758 (2.36)		
A4	5.17164 (10.92)			0.02339* (.95)		
A5	6.99342 (13.43)			0.02767* (1.02)		
A6	8.16449 (13.95)			0.09532 (3.13)		

TABLE 5.5 - Continued

Independent Variables	DAYDIG Per Year Given Hospitalization			ADMDIG Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	1.74285 (4.37)	1.86550 (4.65)	1.85894 (4.64)	-0.01391* (-.67)	0.00691* (.33)	0.00644* (.31)
SINGLE FEMALE	-0.29574* (-.64)	-0.30480* (-.66)	-0.25973* (-.56)	-0.06655 (-2.78)	-0.04691* (-1.95)	-0.04460* (-1.85)
URBAN	2.38233 (10.35)	2.44719 (10.65)	2.43329 (10.56)	-0.07228 (-6.03)	-0.07010 (-5.86)	-0.07141 (-5.95)
WELFARE	2.79028 (5.83)	2.69855 (5.65)	2.02490 (2.95)	0.13228 (5.31)	0.12533 (5.04)	0.12427 (3.47)
INCOME	-0.03791* (-1.66)	-0.04114* (-1.83)	-0.09173 (-2.33)	-0.00254 (-2.13)	-0.00271 (-2.32)	-0.00355* (-1.73)
W.Y			0.33365* (1.70)			0.02062 (2.02)
D68.Y			0.09259* (1.41)			0.00046* (.13)
D69.Y			0.14444 (2.24)			0.00123* (.37)
D70.Y			-0.00503* (-.08)			0.00126* (.37)
D71.Y			0.03213* (.50)			0.00209* (.62)
D68.W.Y			-0.49305* (-1.86)			-0.04306 (-3.12)
D69.W.Y			-0.48611* (-1.83)			-0.02748* (-1.98)
D70.W.Y			0.27228* (.85)			-0.00129* (-.08)
D71.W.Y			-0.12529* (-.52)			-0.02664 (-2.11)
D68	-0.07959* (-.25)	-0.02861* (-.09)	-0.40512* (-3.31)	-0.01251* (-.76)	-0.01100* (-.67)	-0.03035* (-1.44)
D69	-0.70182 (-2.26)	-0.71446 (-2.30)	-0.47960* (-3.36)	0.00730* (.45)	0.00861* (.53)	-0.00378* (-.05)
D70	-0.68081 (-2.21)	-0.68294 (-2.21)	-0.34024* (-2.25)	-0.01676* (-1.04)	-0.01535* (-.95)	-0.03695* (-1.52)
D71	-0.95515 (-3.01)	-0.91308 (-2.88)	0.57657* (.42)	0.00779* (.47)	0.01018* (.62)	-0.03529* (-1.50)
constant	6.31438	3.69948	3.47160	1.17247	1.13504	1.15336
R ²	0.07614	0.07432	0.07651	0.01472	0.01290	0.01468
S.E.E.	9.74797	9.75260	9.75045	0.50770	0.50791	0.50794
F	36.203	64.184	25.972	6.561	10.45	4.670

*Insignificant at 95% Confidence Level

TABLE 5.6

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYGUR and ADMGUR
5,857 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYGUR Per Year Given Hospitalization			ADMGUR Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.04471* (-.64)	-0.22291 (-2.01)		0.00571* (1.74)	0.00942* (1.83)
D68.S			0.1687 (.87)			-0.01455* (-1.61)
D69.S			0.32221* (1.63)			0.00195* (.21)
D70.S			0.37953* (1.95)			-0.00331* (-.36)
D71.S			0.29150* (1.50)			-0.00673* (-.74)
C1	-1.75544 (-4.56)			-0.00173* (-.10)		
C2	-0.96251 (-2.43)			0.00210* (.11)		
C3	-0.64517* (-1.45)			0.01093* (.52)		
C4	-0.24239* (-.48)			0.02031* (.85)		
C5	0.09320* (.19)			0.04632* (1.99)		
AGE (A)		0.20007 (23.54)	0.21875 (14.72)		0.00082 (2.05)	0.00046* (.67)
D68.A			-0.01265* (-.51)			0.00015* (.13)
D69.A			-0.00834* (-.33)			0.00090* (.77)
D70.A			-0.01771* (-.70)			0.00120* (1.02)
D71.A			-0.06393 (-2.58)			-0.00008* (-.07)
A1	0.51473* (.80)			0.00297* (.10)		
A2	1.46542 (2.25)			-0.01691* (-.55)		
A3	2.11080 (3.36)			-0.01345* (-.46)		
A4	4.49743 (7.10)			0.01004* (.34)		
A5	7.72122 (11.25)			0.00920* (.29)		
A6	12.70548 (16.72)			0.06047* (1.69)		

TABLE 5.6 - Continued

Independent Variables	DAYGUR Per Year Given Hospitalization			ADMGUR Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	2.97216 (5.71)	3.91293 (7.47)	3.92589 (7.48)	0.00089* (.04)	0.01650* (.68)	0.01605* (.66)
SINGLE FEMALE	-1.42447 (-2.71)	-0.77112* (-1.45)	-0.75818* (-1.42)	-0.05116 (-2.07)	-0.04014* (-1.61)	-0.04525* (-1.82)
URBAN	1.20118 (4.69)	1.37834 (5.36)	1.36230 (5.29)	-0.05895 (-4.90)	-0.05721 (-4.77)	-0.05827 (-4.85)
WELFARE	2.24188 (3.63)	2.47091 (3.97)	2.42125 (2.65)	0.10014 (3.45)	0.10145 (3.50)	0.12414 (2.91)
INCOME	-0.03379* (-1.40)	-0.06641 (-2.76)	-0.07222* (-1.54)	-0.00174* (-1.53)	-0.00213* (-1.90)	-0.00489 (-2.23)
W.Y			0.23337* (.86)			0.00758* (.60)
D68.Y			-0.03407* (-.44)			0.00222* (.61)
D69.Y			0.01083* (.16)			0.00344* (1.08)
D70.Y			0.02950* (.40)			0.00694 (2.00)
D71.Y			0.01102* (.17)			0.00356* (1.16)
D68.W.Y			-0.22515* (-.66)			-0.03146* (-1.97)
D69.W.Y			-0.54178* (-1.49)			-0.02891* (-1.70)
D70.W.Y			-0.86279 (-2.21)			-0.02571* (-1.41)
D71.W.Y			0.04431* (.13)			0.00781* (.49)
D68	-0.59134* (-1.67)	-0.54701* (-1.53)	-0.37061* (-.23)	-0.00239* (-.14)	-0.00203* (-.12)	0.03655* (.48)
D69	-1.10634 (-3.10)	-1.02924 (-2.87)	-1.84622* (-1.09)	-0.00533* (-.32)	-0.00552* (-.33)	-0.07169* (-.91)
D70	-1.09858 (-3.13)	-1.04651 (-2.96)	-1.72072* (-.98)	-0.00458* (-.28)	-0.00466* (-.28)	-0.08885* (-1.09)
D71	-1.13047 (-3.20)	-1.13345 (-3.19)	0.71047* (.43)	-0.00471* (-.28)	-0.00555* (-.33)	-0.00196* (-.03)
constant	6.35760	-0.46140	-0.63697	1.15552	1.10281	1.12124
R ²	0.14254	0.12811	0.13181	0.01160	0.01058	0.01383
S.E.E.	8.98959	9.05797	9.05190	0.42245	0.42235	0.42227
F	48.509	78.073	31.600	3.426	5.68	2.920

*Insignificant at 95% Confidence Level

TABLE 5.7

Parameter Estimates for Three Specifications
of the Linear Utilization Function With Alternate
Continuous Dependent Variables DAYDEL and ADMDEL
7,956 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYDEL Per Year Given Hospitalization			ADMDEL Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.06058 (-1.59)	-0.05723* (-.98)		0.00100* (.27)	-0.00289* (-.50)
D68.S			-0.07676* (-.72)			-0.00146* (-.14)
D69.S			0.08242* (.74)			0.01376* (1.26)
D70.S			-0.11791* (-1.01)			-0.00273* (-.24)
D71.S			0.09820* (.84)			0.01839* (1.61)
C1	-0.21558* (-1.18)			-0.03299* (-1.85)		
C2	-0.71425 (-3.64)			-0.02994* (-1.55)		
C3	-0.35653* (-1.58)			0.00035* (.02)		
C4	-0.69527 (2.72)			0.00327* (.13)		
C5	-0.39490* (-1.55)			-0.01951* (-.78)		
AGE (A)		0.01108* (1.33)	0.01114* (.88)		-0.00066* (-.80)	-0.00085* (-.69)
D68.A			-0.01308* (-.53)			-0.00292* (-1.21)
D69.A			0.00265* (.11)			0.00247* (1.02)
D70.A			0.02419* (.91)			0.00424* (1.62)
D71.A			-0.01112* (-.40)			-0.00216* (-.84)
A1	-0.12834* (-.79)			-0.01867* (-1.17)		
A2	0.31997* (1.54)			-0.01125* (-.55)		
A3	-0.05953* (-.21)			-0.03041* (-1.10)		
A4	0.23183* (.41)			-0.05947* (1.06)		
A5	-0.78716* (-1.44)			-0.09963* (-.57)		
A6	2132* (.76)			-0.26925* (-.95)		

TABLE 5.7 - Continued

Independent Variables	DAYDEL Per Year Given Hospitalization			ADMDEL Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	0.01257* (.03)	-0.04392* (-.10)	-0.09188* (-.22)	0.01082* (.26)	0.01314* (.32)	0.01373* (.33)
SINGLE FEMALE	-1.32692 (-4.44)	-1.22119 (-4.10)	-1.19767 (-3.99)	-0.08567 (-2.92)	-0.07242 (-2.48)	-0.07432 (-2.53)
URBAN	-0.34797 (-2.97)	-0.33937 (-2.90)	-0.33765 (-2.89)	-0.04807 (-4.18)	-0.04827 (-4.21)	-0.04797 (-4.18)
WELFARE	0.66394 (2.30)	0.67724 (2.34)	0.39796* (.93)	0.13448 (4.73)	0.13173 (4.64)	0.16358 (3.92)
INCOME	-0.02263* (-1.62)	-0.02573* (-1.85)	-0.00354* (-.15)	-0.00432 (-3.13)	-0.00422 (-3.08)	-0.00189* (-.82)
W.Y			-0.04670* (-.35)			-0.00832* (-.63)
D68.Y			-0.02870* (-.69)			-0.00312* (-.77)
D69.Y			0.00402* (.10)			0.00195* (.51)
D70.Y			-0.04633* (-1.16)			-0.00642* (-1.64)
D71.Y			-0.06561* (-1.65)			-0.00499* (-1.28)
D68.W.Y			0.13121* (.60)			-0.00767* (-.36)
D69.W.Y			0.12769* (.72)			0.00139* (.08)
D70.W.Y			0.47203 (2.21)			0.00283* (.14)
D71.W.Y			-0.03903* (-.22)			-0.01652* (-.93)
D68	-0.24415* (-1.50)	-0.21911* (-1.35)	0.65577* (.97)	0.00291* (.18)	0.00393* (.25)	0.11778 (1.78)
D69	-0.21228* (-1.28)	-0.22083* (-1.33)	-0.68023* (-1.01)	-0.00841* (-.52)	-0.00893* (-.55)	-0.15579 (-2.35)
D70	-0.26559* (-1.52)	-0.27251* (-1.56)	-0.33979* (-.47)	-0.01364* (-.79)	-0.01374* (-.80)	-0.09704* (-1.36)
D71	-0.35630 (-2.04)	-0.35028 (-2.00)	0.07635* (.11)	-0.00138* (-.08)	-0.00220* (-.13)	0.02554* (.37)
constant	7.47779	7.01672	6.88050	1.27308	1.25641	1.26554
R ²	0.00814	0.00504	0.00712	0.00915	0.00779	0.01098
S.E.E.	4.97610	4.98105	4.98117	0.48877	0.48883	0.48856
F	3.257	3.659	2.031	3.663	5.68	3.144

*Insignificant at 95% Confidence Level

TABLE 5.8

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYAPV and ADMAPV
5,901 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYAPV Per Year Given Hospitalization			ADMAPV Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.40996 (-4.25)	-0.26576* (-1.78)		0.01797 (7.58)	0.02276 (3.2)
D68.S			-0.14078* (-.56)			-0.01594 (-2.58)
D69.S			-0.68830 (-2.63)			-0.01772 (-2.76)
D70.S			-0.38186* (-1.50)			0.00308* (.49)
D71.S			0.21313* (.86)			0.00068* (.11)
C1	-1.24314* (-1.84)			0.01603* (.97)		
C2	-2.73359 (-4.05)			0.02596* (1.56)		
C3	-2.16792 (-2.98)			0.04274 (2.39)		
C4	-2.20862 (-2.84)			0.04321 (2.26)		
C5	-2.23031 (-3.04)			0.10734 (5.95)		
AGE (A)		0.20559 (17.06)	0.18375 (8.49)		0.00046* (1.53)	0.00099* (1.87)
D68.A			0.04621* (1.24)			-0.00122* (-1.33)
D69.A			0.06616* (1.91)			-0.00133* (-1.56)
D70.A			-0.02570* (-.73)			-0.00023* (-.27)
D71.A			0.03730* (1.06)			-0.00028* (-.33)
A1	1.78167 (2.20)			0.04398 (2.20)		
A2	2.77768 (3.35)			0.03592* (1.76)		
A3	3.80513 (4.73)			0.04268 (2.16)		
A4	5.79445 (7.13)			0.06394 (3.20)		
A5	8.67392 (8.99)			0.03427* (1.44)		
A6	14.27993 (14.43)			0.03475* (1.43)		

TABLE 5.8 - Continued.

Independent Variables	DAYAPV Per Year Given Hospitalization			ADMAPV Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	1.86151 (2.95)	2.50432 (4.02)	2.48690 (3.99)	0.02813* (1.81)	0.03862 (2.52)	0.03712 (2.42)
SINGLE FEMALE	0.66918* (.87)	1.13895* (1.46)	1.06347* (1.37)	-0.01127* (-.59)	0.00093* (.05)	0.00390* (.20)
URBAN	1.91637 (4.67)	2.13691 (5.22)	2.10853 (5.14)	-0.02831 (-2.80)	-0.02959 (-2.94)	-0.03003 (-2.98)
WELFARE	2.00083 (2.43)	2.09367 (2.53)	1.19213* (.98)	0.07955 (3.92)	0.07920 (3.90)	0.05476* (1.84)
INCOME	-0.09108 (-2.25)	-0.12387 (-3.10)	-0.22343 (-2.95)	-0.00029* (-.29)	-0.00018* (-.18)	-0.00378 (-2.03)
W.Y			0.03964* (.13)			0.00602* (.80)
D68.Y			0.09512* (.80)			0.00312* (1.07)
D69.Y			0.19211* (1.56)			0.00815 (2.70)
D70.Y			0.11832* (1.02)			0.00346* (1.22)
D71.Y			0.13568* (1.24)			0.00450* (1.67)
D68.W.Y			-0.85421* (-1.3)			-0.02538* (-1.57)
D69.W.Y			-0.13311* (-.28)			0.02023* (1.76)
D70.W.Y			0.61542* (1.13)			-0.01318* (-.98)
D71.W.Y			1.43083 (2.78)			0.01156* (.92)
D68	-0.35499* (-.63)	-0.35046* (-.62)	-2.17972* (-1.03)	-0.00283* (-.21)	-0.00354* (-.26)	0.10213* (1.97)
D69	-0.06637* (-.12)	-0.02308* (-.04)	-1.37391* (-.64)	-0.00189* (-.14)	-0.00229* (-.17)	0.07575* (1.44)
D70	-0.71681* (-1.30)	-0.78827* (-1.43)	1.16607* (.54)	0.00311* (.23)	0.00317* (.23)	-0.01379* (-.26)
D71	-0.99425* (-1.82)	-1.07712* (-1.97)	-4.61877 (-2.20)	0.00544* (.41)	0.00533* (.40)	-0.01195* (-.23)
constant	5.26244	0.15178	1.12338	1.03179	1.00942	0.98628
R ²	0.07702	0.07123	0.07724	0.01890	0.01745	0.02350
S.E.E.	14.08364	14.11690	14.09151	0.34647	0.34646	0.34589
F	24.533	41.061	17.555	5.665	9.51	5.047

*Insignificant at 95% Confidence Level

TABLE 5.9

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYSUR and ADMSUR
20,680 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYSUR Per Year Given Hospitalization			ADMSUR Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.37646 (-7.44)	-0.43081 (-5.51)		0.03819 (18.72)	0.04998 (15.82)
D68.S			0.15452* (1.16)			-0.02352 (-4.37)
D69.S			-0.14905* (-1.08)			-0.01904 (-3.40)
D70.S			0.11429* (.84)			-0.01001* (-1.83)
D71.S			0.20479* (1.49)			-0.02162 (-3.91)
C1	-1.03524 (-3.42)			0.02535 (2.07)		
C2	-2.21237 (-7.27)			0.05536 (4.50)		
C3	-2.09603 (6.36)			0.10473 (7.86)		
C4	-1.87064 (5.08)			0.13039 (8.76)		
C5	-1.79699 (4.89)			0.22130 (14.90)		
AGE (A)		0.25115 (40.90)	0.26201 (24.65)		0.00064 (2.56)	0.00055* (1.28)
D68.A			0.00423* (.23)			-0.00003* (-.04)
D69.A			-0.00732* (-.40)			-0.00000* (-.00)
D70.A			-0.02628* (-1.46)			0.00034* (.47)
D71.A			-0.03615* (-1.97)			0.00009* (.12)
A1	1.97137 (4.56)			0.03197* (1.83)		
A2	3.43366 (7.89)			0.02084* (1.18)		
A3	5.26228 (12.48)			0.02836* (1.66)		
A4	7.68498 (18.09)			0.02424* (1.41)		
A5	11.91579 (25.10)			0.03286* (1.71)		
A6	15.01370 (27.59)			0.08546 (3.88)		

TABLE 5.9 - Continued

Independent Variables	DAYSUR Per Year Given Hospitalization			ADMSUR Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	3.13681 (8.91)	3.53892 (10.08)	3.53179 (10.04)	-0.04183 (-2.94)	0.00074* (.05)	-0.00098* (-.07)
SINGLE FEMALE	-0.62297* (-1.67)	-0.46647* (-1.23)	-0.44327* (-1.17)	-0.06951 (-4.60)	-0.02709* (-1.78)	-0.02924* (-1.91)
URBAN	0.75742 (4.13)	0.91004 (4.98)	0.91040 (4.98)	-0.02718 (-3.67)	-0.02523 (-3.42)	-0.02538 (-3.44)
WELFARE	3.41399 (7.31)	3.51486 (7.52)	2.79822 (4.27)	0.13442 (7.12)	0.12968 (6.87)	0.13739 (5.19)
INCOME	-0.08070 (-4.68)	-0.10598 (-6.21)	-0.11882 (-3.93)	-0.00188 (-2.69)	-0.00209 (-3.03)	-0.00445 (-3.65)
W.Y			0.32107* (1.78)			0.00255* (.35)
D68.Y			0.03630* (.72)			0.00407* (2.00)
D69.Y			0.01499* (.31)			0.00421 (2.16)
D70.Y			-0.03402* (-.66)			0.00103* (.50)
D71.Y			0.02233* (.47)			0.00404 (2.12)
D68.W.Y			-0.49783 (-1.87)			-0.01160 (1.08)
D69.W.Y			-0.60214 (-2.31)			-0.00209* (-.20)
D70.W.Y			0.12917* (.50)			-0.00170* (-.16)
D71.W.Y			0.05414* (.23)			-0.00395* (-.41)
D68	-0.28621* (-1.10)	-0.25322* (-.97)	-1.21217* (-1.09)	-0.00904* (-.86)	-0.00821* (-.78)	0.06110* (1.35)
D69	-0.38169* (-1.46)	-0.34907* (-1.33)	0.50130* (1.72)	-0.02077* (-1.96)	-0.01999* (-1.89)	0.02811* (.60)
D70	-0.68391 (-2.63)	-0.65522 (-2.52)	0.30918* (.27)	0.00996* (.95)	0.01094* (1.04)	0.03011* (.64)
D71	-1.176 (-4.51)	-1.16924 (-4.48)	-0.46803* (-1.40)	-0.00916* (-.87)	-0.00850* (-.81)	0.04502* (.96)
constant	6.58160	0.47027	0.26583	1.13169	1.04407	1.01540
R ²	0.10975	0.10753	0.10893	0.02719	0.02732	0.02905
S.E.E.	12.45104	12.46383	12.45918	0.50346	0.50332	0.50308
F	127.336	226.374	90.157	28.872	52.77	22.067

* Insignificant at 95% Confidence Level

TABLE 5.10

Parameter Estimates for Three Specifications
of the Linear Utilization Function With Alternate
Continuous Dependent Variables DAYMED and ADMED
23,015 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYMED per Year Given Hospitalization			ADMED per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		0.09377* (1.42)	0.12547* (1.23)		0.06875 (17.95)	0.07007 (11.86)
D68.S			-0.14619* (-.83)			-0.01775* (-1.75)
D69.S			0.03243* (.18)			-0.00290* (-.28)
D70.S			-0.05104* (-.28)			0.00468* (.45)
D71.S			-0.03528* (-.20)			0.01000* (.96)
C1	0.10965* (.28)			0.07887 (3.49)		
C2	-0.43802* (-1.07)			0.13666 (5.74)		
C3	-0.61485* (-1.36)			0.14648 (5.57)		
C4	0.46461* (.92)			0.30398 (10.42)		
C5	1.66384 (3.42)			0.48129 (17.03)		
AGE (A)		0.23337 (32.59)	0.22712 (17.90)		0.00311 (7.40)	0.00276 (3.73)
D68.A			0.00492* (.23)			-0.00163* (-1.29)
D69.A			0.02092* (.98)			0.00061* (.49)
D70.A			-0.01031* (-.48)			0.00169* (1.35)
D71.A			0.02210* (1.06)			0.00181* (1.50)
A1	2.51787 (4.72)			0.11742 (3.80)		
A2	2.99104 (5.47)			0.00288* (.09)		
A3	3.96928 (7.55)			-0.02247* (-.74)		
A4	7.36986 (14.02)			0.06191 (2.03)		
A5	10.26285 (18.28)			0.18452 (5.67)		
A6	15.32622 (26.47)			0.27912 (8.31)		

TABLE 5.10 - Continued

Independent Variables	DAYMED Per Year Given Hospitalization			ADMED Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	3.36811 (8.69)	3.92304 (9.94)	3.90288 (9.88)	0.01980*	0.10156 (4.43)	0.10027 (4.37)
SINGLE FEMALE	0.38198* (.82)	0.98749 (2.10)	0.97292 (2.07)	-0.15404 (-5.72)	-0.06649 (2.44)	-0.06782 (-2.49)
URBAN	2.19286 (8.37)	2.45518 (9.75)	2.44978 (9.35)	-0.19098 (-12.56)	-0.17733 (-11.67)	-0.17908 (-11.78)
WELFARE	5.91038 (11.93)	6.01152 (12.13)	6.34231 (8.93)	0.49185 (17.11)	0.49179 (17.09)	0.52527 (12.76)
INCOME	-0.08341 (-3.28)	-0.11231 (-4.45)	-0.11332 (-2.50)	-0.00432 (-2.92)	-0.00579 (-3.94)	-0.00836 (-3.18)
W.Y			-0.06787* (-.32)			-0.02990 (-2.46)
D68.Y			0.01319* (.18)			0.00257* (.59)
D69.Y			0.0658* (.62)			0.00651* (1.50)
D70.Y			-0.02153* (-.30)			0.00438* (1.05)
D71.Y			-0.00561* (-.08)			0.00367* (-.87)
D68.W.Y			-0.55422* (-1.84)			-0.01274* (-.73)
D69.W.Y			0.47725* (1.51)			0.07433 (4.06)
D70.W.Y			-0.26556* (-.88)			0.00317* (.18)
D71.W.Y			0.18328* (.69)			0.04685 (3.05)
D68	-0.433133* (-1.25)	-0.39985* (-1.16)	-0.03918* (-.03)	-0.03173 (-1.58)	-0.02983* (-1.49)	0.10207* (1.21)
D69	-0.79566 (-2.30)	-0.76963 (-2.22)	-2.23489* (-1.51)	0.00988* (.49)	0.01352 (.67)	-0.05888* (-.69)
D70	-1.12134 (-3.27)	-1.12471 (-3.28)	-0.24872* (-.17)	-0.00530* (-.27)	-0.00430 (-.22)	-0.12715 (-1.46)
D71	-1.31115 (-3.84)	-1.28310 (-3.76)	-2.24575* (-1.54)	-0.00653* (-.33)	-0.00387* (-.20)	-0.15816 (-1.88)
constant	6.25515	0.12405	0.30252	1.32819	1.11739	1.14692
R ²	0.07188	0.06809	0.06875	0.04439	0.03953	0.04145
S.E.E.	17.32619	17.35813	17.35843	1.00517	1.00752	1.00688
F	89.041	152.794	60.602	53.404	86.06	35.503

*Insignificant at 95% Confidence Level

analysis. Nevertheless, the explanatory power of some of the diagnostic categories exceeds that registered by the equations for the broad definition in section one. This can be viewed as a first approximation to the information gains attending the disaggregation strategy.

Consider now the component variables of the equations. The dummy variables (C1-C5) for family size carry coefficients that are generally negative in the volume (DAY) equations and generally positive in the frequency (ADM) equations. This pattern implies that average length of stay tends to be shorter the larger the patient's family, a finding that could reflect either or both of the following: (a) the tendency for short-stay paediatric cases to be concentrated among large families; (b) the possibility that the adult members of large families are inclined toward shorter stays to minimize the economic and social costs of their absence from the home. Notable exceptions to these generalizations are the circulatory and genito-urinary conditions (Tables 5.3 and 5.6 respectively) when the C_i coefficients are uniformly insignificant in both the volume and frequency equations. Also, it is well to draw the reader's attention to the fact that in most of the diagnostic and treatment categories the Model 2 size coefficient (S) provides a poor approximation to the incremental size effect since, as the Model 1 results indicate, that effect is quite non-linear over the relevant range of family sizes.

The dummy variables (A1-A6) for age of head perform much as they did in the total use analysis of the preceding section. Thus, with the exception of the deliveries equation where they are uniformly insignificant, they indicate that volume is more age-related than is frequency, and hence that average length of stay rises markedly with

age. For example, in the CIR and GUR equations, the A6 coefficients imply average length of stays nine to eleven days longer than the reference age group. This result may well reflect the relatively slow recovery period common among elderly patients.

In the total use analysis, the most striking finding with respect to marital status was the relatively long average length of stay among members of families where the female spouse was absent. The same finding emerges from the diagnostic and treatment equations with the exception of the CIR equations where the coefficients of the single male and single female variables in the volume equation are positive and similar in magnitude, i.e., it makes little difference if the single adult is male or female. In general, however, members of families headed by a single male are likely to spend 1.5-2.5 days longer in hospital than members of families where only the female adult is present."

The income variable in the total use analysis displayed negative but weak coefficients, and yielded very low income elasticities. Similarly weak and negative coefficients are evident in the diagnostic and treatment equations.

The welfare variable carries large positive, and statistically significant coefficients in all of these narrowly defined volume and frequency equations. The implicit length of stay differential between socially assisted and socially unassisted families is most pronounced in the respiratory equations where the net welfare effect is 2.2 days. This doubtless reflects a combination of environmental and lifestyle factors which put the disadvantaged at greater risk of such respiratory problems as pneumonia and bronchitis.

Three of the six diagnostic equations reveal significant copayment effects. Length of stay declines are registered in three diagnostic categories, CIR, DIG, and GUR whose component illnesses and conditions tend to involve longer lengths of stay than conditions in the other categories, REG, DEL, and APV where copayment effects are absent. The largest copayment effect is for circulatory conditions where coefficients of -1.8, -2.1, -3.1 and -2.6 are registered in the days equation. Relative to the mean volume of use in the pre-copayment period (17.2 days), these coefficients represent declines ranging from ten to eighteen per cent. The smallest copayment effect is for digestive conditions, where declines in volume range from seven per cent to ten per cent depending upon which of the copayment years is selected. Again, however, the welfare and non-welfare comparisons are instructive since declines in use were common to both groups. For example, in circulatory conditions the volume of care experienced by the non-welfare subsample in 1970 was 14.9 per cent below its pre-copayment level of 16.9 days. Correspondingly, the welfare group who was exempt from copayment experienced an even larger decline of -18.7 per cent. Thus the observed declines in utilization in the copayment period are again open to two interpretations.

Both the medical and surgical treatment equations show significant volume reductions in some of the copayment years. The range of the reductions is approximately the same (6.5 to 11.5 per cent) in both categories although the effects occur in different years. Expectations, therefore, of larger price effects among medical patients are not supported by the data.

Use by Practice Organization of Admitting Physician

The principal hypothesis to be tested in this section is that hospitalization decisions of solo general practitioners are more likely to exhibit a sensitivity to copayment than are the assumedly more conservative decisions of group general practitioners. The dependent variables in the analysis are DAYSGP, ADMSGP, DAYGGP, and ADMGGP. Tables 5.11 and 5.12 contain the ordinary least squares estimates for each of these practice related use definitions.

Many of the patterns previously demonstrated for both the broad and diagnostic definitions of use are present in this set of results. Family size, age of head, and marital status each play some part in explaining inter-family variations in hospitalization experiences. Large families, older families, and families headed by a single adult are in what might be called the high-use categories. There are no noteworthy differences in the incremental effects of size, age, and marital status across practice setting, nor indeed is there any reason to expect such differences. In addition, the rural-urban differential is as pronounced in these equations as in previous results, with the pattern being one of uniformly positive and significant coefficients in the days equation and uniformly negative and significant coefficients in the admissions equation. Together these coefficients imply an incremental length of stay associated with urban residence of two to two and one-half days.

TABLE 5.11

Parameter Estimates for Three Specifications
of the Linear Utilization Function With Alternate
Continuous Dependent Variables DAYSGP and ADMSGP
11,216 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYSGP Per Year Given Hospitalization			ADMSGP Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		0.03354* (.50)	0.17916* (1.70)		0.07047 (13.05)	0.08903 (10.49)
D68.S			-0.33081* (-1.78)			-0.03300 (-2.21)
D69.S			-0.23324* (-1.23)			-0.03212 (-2.11)
D70.S			-0.05839* (-.33)			-0.01401* (-.99)
D71.S			-0.32992* (-1.83)			-0.03842 (-2.64)
C1	-0.02020 (-.05)			0.08460 (2.69)		
C2	-0.71559* (-1.73)			0.11162 (3.35)		
G3	-0.02570* (-.06)			0.22385 (6.07)		
C4	0.90565* (1.80)			0.32262 (7.95)		
C5	1.93236 (3.91)			0.53073 (13.34)		
AGE (A)		0.19806 (26.69)	0.19254 (14.61)		0.00187 (3.12)	0.00120* (1.13)
D68.A			-0.02162* (-.94)			0.00021* (.11)
D69.A			-0.00329* (-.15)			0.00092* (.51)
D70.A			-0.01063* (-.50)			0.00077* (.45)
D71.A			0.05570 (2.61)			0.00177* (1.04)
A1	1.00194* (1.97)			0.02019* (.49)		
A2	0.77545* (1.45)			-0.11886 (-2.77)		
A3	2.04133 (3.96)			-0.14269 (-3.44)		
A4	5.07516 (9.72)			-0.06732* (-1.60)		
A5	8.24698 (14.35)			0.04194* (.91)		
A6	12.96849 (21.52)			0.23762 (4.90)		

TABLE 5.11 - Continued

Independent Variables	DAYSGP Per Year Given Hospitalization			ADMSGP Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	1.75457 (4.17)	2.34378 (5.46)	2.31105 (5.39)	-0.12478 (-3.69)	-0.02698* (-.78)	-0.03001* (-.87)
SINGLE FEMALE	-0.65627* (-1.32)	0.05343* (.11)	-0.01274* (-.03)	-0.19699 (-4.92)	-0.08943 (-2.21)	-0.09595 (-2.37)
URBAN	2.21323 (8.04)	2.48426 (9.01)	2.46832 (8.94)	-0.21110 (-9.54)	-0.19332 (-8.73)	-0.19802 (-8.92)
WELFARE	4.28854 (1.74)	4.42449 (7.69)	4.71949 (5.93)	0.39062 (8.47)	0.39168 (8.47)	0.49044 (7.67)
INCOME	-0.10312 (-3.67)	-0.13916 (-4.96)	-0.18837 (-3.71)	-0.00360* (-1.59)	-0.00571 (-2.54)	-0.00956 (-2.34)
W.Y			0.28103* (1.27)			-0.01065* (-.60)
D68.Y			0.14493* (1.66)			-0.00121* (-.17)
D69.Y			0.02283* (.26)			0.00514* (.74)
D70.Y			0.06541* (.84)			0.00547* (.88)
D71.Y			0.07647* (.97)			0.01621 (2.55)
D68.W.Y			-0.44171* (-1.29)			-0.01916* (-.70)
D69.W.Y			-0.99450 (-3.09)			-0.05760 (-2.21)
D70.W.Y			-0.35311* (-1.21)			-0.02069* (-.89)
D71.W.Y			-0.17717* (-.37)			0.01331* (.52)
D68	-0.70786* (-1.96)	-0.69844* (-1.92)	0.78685* (.53)	-0.00908* (-.31)	-0.00780* (.27)	0.12099* (1.02)
D69	-0.72346 (-2.04)	-0.73903 (-2.07)	0.33011* (.22)	-0.01919* (-.67)	-0.01899* (-.66)	0.04220* (.35)
D70	-0.69481 (-2.05)	-0.74551 (-2.19)	-0.28715* (-.20)	0.02570* (.94)	0.02422* (.89)	0.01880* (.16)
D71	-0.52629* (-1.54)	-0.59262* (-1.72)	-2.38481* (-1.66)	0.00420* (.15)	0.00152* (.05)	-0.03477* (-.30)
constant	6.74910	0.83595	0.75192	1.43817	1.19250	1.17053
R ²	0.09299	0.08254	0.08536	0.04514	0.03671	0.03950
S.E.E.	12.27175	12.33729	12.32770	0.98724	0.99119	0.99050
F	57.389	91.636	37.286	26.459	38.81	16.430

*Insignificant at 95% Confidence Level

TABLE 5.12

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYGGP and ADMGGP
23,026 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYGGP Per Year Given Hospitalization			ADMGGP Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		0.01700* (.29)	0.07835* (.88)		0.08153 (23.50)	0.09100 (17.20)
D68.S			-0.14860* (-.97)			-0.02676 (-2.93)
D69.S			-0.16372* (-1.03)			-0.00200* (-.21)
D70.S			-0.13241* (-.80)			-0.01323* (-1.33)
D71.S			0.05016* (.31)			-0.01684* (-1.73)
C1	0.10432 (.30)			0.10707 (5.18)		
C2	-0.34159* (-.95)			0.19087 (8.82)		
C3	-0.34401* (-.86)			0.24555 (10.24)		
C4	0.37213* (.84)			0.34162 (12.82)		
C5	1.70871 (3.90)			0.57351 (21.85)		
AGE (A)		0.22484 (34.54)	0.22056 (19.50)		-0.00013* (-.33)	-0.00080* (-1.18)
D68.A			0.01264* (.66)			-0.00091* (-.79)
D69.A			0.00926* (.48)			0.00128* (1.11)
D70.A			0.00145* (.07)			0.00243 (2.04)
D71.A			0.00430* (.22)			0.00152* (1.32)
A1	1.59682 (3.79)			0.03873* (1.53)		
A2	2.07513 (4.60)			-0.08256 (-3.06)		
A3	3.00690 (6.84)			-0.12661 (-4.81)		
A4	5.84315 (13.04)			-0.07276 (-2.71)		
A5	9.94006 (19.72)			0.03234* (1.07)		
A6	16.53645 (30.26)			0.14731 (4.50)		

TABLE 5.12 - Continued

Independent Variables	DAYGGP Per Year Given Hospitalization			ADMGGP Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	2.88558 (7.37)	3.63266 (9.18)	3.64017 (9.20)	-0.03417* (-1.46)	0.06212 (2.63)	0.06156 (4.23)
SINGLE FEMALE	-0.27* (-.47)	0.57595* (1.31)	0.58418* (1.32)	-0.14937 (-5.72)	-0.05033* (-1.91)	-0.05063* (-1.92)
URBAN	1.25428 (5.49)	1.55778 (6.80)	1.55891 (4.36)	-0.22132 (-16.15)	-0.20840 (-15.23)	-0.20920 (-15.28)
WELFARE	3.86998 (8.07)	3.98472 (8.28)	3.46956 (4.86)	0.41989 (14.61)	0.41892 (14.57)	0.39705 (9.31)
INCOME	-0.08674 (-3.51)	-0.13068 (-5.34)	-0.12388 (-2.97)	-0.00558 (-3.77)	-0.00740 (-5.07)	-0.01113 (-4.47)
W.Y			-0.27711* (-1.24)			-0.01320* (-.99)
D68.Y			-0.02114* (-.29)			0.00730* (1.68)
D69.Y			0.03487* (.48)			0.00825* (1.91)
D70.Y			-0.06867* (-.97)			0.00435* (1.03)
D71.Y			-0.00371* (-.05)			0.00297* (.73)
D68.W.Y			0.14029* (.44)			-0.00220* (-.12)
D69.W.Y			1.08052 (3.58)			0.06912 (3.83)
D70.W.Y			0.16484* (.47)			-0.00151* (-.07)
D71.W.Y			0.73649 (2.62)			0.03752 (2.23)
D68	-0.21953* (-.73)	-0.13477* (-.44)	0.00724* (.01)	-0.03862 (-2.13)	-0.03587* (-1.98)	0.06568* (.92)
D69	-0.38015* (-1.24)	-0.32381* (-1.05)	-0.48600* (-1.39)	-0.03005* (-1.63)	-0.02584* (-1.40)	-0.13108* (-1.75)
D70	-1.03867 (-3.28)	-1.00606 (-3.17)	-0.16790* (-.13)	-0.05688 (-3.00)	-0.05388 (-2.84)	-0.13476* (-1.73)
D71	-1.32581 (-4.24)	-1.20371 (-3.84)	-1.67066* (-1.31)	-0.05038 (-2.69)	-0.04276 (-2.28)	-0.07051* (-.93)
constant	6.70370	0.64135	0.59770	1.43149	1.26705	1.28289
R ²	0.07556	0.06702	0.06801	0.05770	0.05386	0.05581
S.E.E.	15.58279	15.65158	15.64905	0.93366	0.93538	0.93476
F	94.017	150.283	59.931	70.439	119.10	48.545

*Insignificant at 95% Confidence Level

The "economic" variables, which necessarily command greater interest, are income, welfare, and copayment. The first two display a consistency across both the solo and group equations. The implied income elasticities of length of stay are -0.03 (solo) and -0.04 (group). Members of families receiving social assistance spend in excess of one day longer in hospital if admitted by a solo practitioner, but only half a day longer if admitted by a group practitioner. In the analysis of admission probabilities, it was argued that group practitioners may work in closer proximity to referring social agencies than solo practitioners. Here it appears that social agencies interact with the group practitioner in such a way as to facilitate relatively early discharge of the socially-assisted patient.

Finally, a comparison of the copayment coefficients in the two equations provides no support for the hypothesis that discharge decisions by solo practitioners are more price sensitive than those of group practitioners. The coefficients for D69 and D70 are negative and significant in the solo days equation. They imply a reduction in each of these years in utilization of seven per cent relative to the pre-copayment years. This is largely a reflection of the fact that average lengths of stay were a half day shorter in each of these years. Price effects are also evident in the group equations but the effects are specific to 1970 and 1971 rather than 1969 and 1970. The group equation coefficients imply volume reductions of ten to twelve per cent from the pre-copayment level of 10.8 days. When translated into length of stay terms these reductions indicate a one half day reduction, or equivalently, a length of stay seven per cent lower than the pre-copayment level. Refinements of the copayment hypothesis tested in

Model 3 fail to identify any underlying interaction effects between copayment and the families' socio-economic characteristics. That is, there is no evidence that larger and/or older families were more affected by copayment than smaller and/or younger families.

Use By Type of Hospital

Since the size and service capability of active treatment institutions in Saskatchewan varies among regional, base, and community hospitals, it is logical to analyze them separately. The catchment populations also differ for these categories of hospitals. The ability of socio-economic family characteristics to explain variations in the use of such hospitals will vary since case mix and case severity are themselves systematically related to the range of resources and technologies available for diagnostic and therapeutic functions. The underlying assumption is that discharge decisions are more discretionary in respect of patients admitted to small community hospitals and, accordingly, the equations for these hospitals should have greater explanatory power. In addition, the deterrent effect of copayment should be relatively greater among patients of these "low" technology institutions than among patients of the "medium" to "high" technology regional and base hospitals.

Tables 5.13 and 5.14 present the days and admissions equations where the dependent variable is alternatively defined for community and regional-base hospitals respectively. The expectation concerning overall explanatory power is only partly borne out in these results. The R^2 for the community equation is six per cent in the admissions

TABLE 5.13

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYCOM and ADMCOM
21,029 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYCOM Per Year Given Hospitalization			ADMCOM Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		0.28011 (5.33)	0.36057 (4.49)		0.10661 (24.40)	0.11854 (17.75)
D68.S			-0.17427* (-1.25)			-0.03573 (-3.07)
D69.S			-0.14352* (-1.00)			-0.01225* (-1.03)
D70.S			-0.07263* (-.50)			-0.00959* (-.80)
D71.S			-0.09701* (-.67)			-0.01517* (-1.26)
C1	0.54184* (1.69)			0.14273 (5.35)		
C2	0.30164* (.90)			0.23327 (8.34)		
C3	0.64267* (1.74)			0.33669 (10.95)		
C4	1.51848 (3.77)			0.45344 (13.48)		
C5	3.31997 (8.49)			0.75953 (23.29)		
AGE (A)		0.17684 (29.38)	0.17144 (16.34)		0.00100 (2.00)	-0.00045* (-.52)
D68.A			-0.01396* (-.77)			-0.00118* (-.79)
D69.A			-0.01126* (-.63)			0.00216* (1.45)
D70.A			0.00714* (.40)			0.00379 (2.53)
D71.A			0.04892 (2.77)			0.00408 (2.78)
A1	1.54755 (3.68)			0.07224 (2.06)		
A2	1.23702 (2.82)			-0.09999 (-2.73)		
A3	1.61965 (3.83)			-0.15651 (-4.44)		
A4	4.26769 (.86)			-0.07514 (-2.13)		
A5	8.02530 (17.10)			0.07841 (2.00)		
A6	12.99143			0.26370 (6.33)		

TABLE 5.13 - Continued

Independent Variables	DAYCOM Per Year Given Hospitalization			ADMCOM Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	2.26868 (6.63)	3.14895 (9.06)	3.14807 (9.05)	-0.07679 (-2.69)	0.05435* (1.88)	0.05256* (1.82)
SINGLE FEMALE	-0.28131* (-.68)	0.75713* (1.80)	0.72701* (1.73)	-0.25473 (-7.34)	-0.11026 (-3.16)	-0.11786 (-3.37)
URBAN	-2.14618 (-4.71)	-1.73794 (-3.80)	-1.79718 (-3.93)	-0.19846 (-5.22)	-0.17405 (-4.57)	-0.17918 (-4.71)
WELFARE	5.30272 (11.57)	5.41883 (11.77)	5.67494 (8.42)	0.61541 (16.10)	0.61912 (16.17)	0.73520 (13.12)
INCOME	-0.08311 (-3.60)	-0.11369 (-4.93)	-0.16214 (-4.15)	-0.00762 (3.95)	-0.00947 (-4.93)	-0.01303 (-4.01)
W.Y			-0.25501* (-1.19)			-0.05566 (-3.11)
D68.Y			0.06535* (.95)			0.00546* (.95)
D69.Y			0.08353* (1.20)			0.00937* (1.62)
D70.Y			0.05299* (.77)			0.00354* (.62)
D71.Y			0.10055* (1.54)			0.00905* (1.66)
D68.W.Y			-0.01621* (-.05)			-0.01414* (-.58)
D69.W.Y			0.69870 (2.49)			0.07635 (3.28)
D70.W.Y			-0.03344* (-.10)			-0.00989 (-.37)
D71.W.Y.			0.37643* (1.31)			0.06635 (2.78)
D68	-0.55856* (-2.00)	-0.50850* (-1.81)	0.46311* (.40)	-0.03612* (-1.55)	-0.03304* (-1.41)	0.13376* (1.40)
D69	-0.76814 (-2.73)	-0.72341 (-2.55)	-0.19224* (-.16)	-0.02360* (-1.00)	-0.01829* (-.78)	-0.12899* (-1.30)
D70	-1.10043 (-3.91)	-1.10007 (-3.88)	-1.41193* (-1.17)	-0.03622* (-1.54)	-0.03367* (-1.43)	-0.18855* (-1.88)
D71	-0.82885 (-2.96)	-0.77275 (-2.74)	-3.33978 (-2.79)	-0.03240* (-1.39)	-0.02641* (-1.12)	-0.22039 (-2.21)
constant	6.89885	1.46583	1.67895	1.46698	1.21375	1.25412
R ²	0.07142	0.05993	0.06114	0.05790	0.05085	0.05352
S.E.E.	13.52163	13.60213	13.59885	1.12800	1.13198	1.13084
F	80.786	121.795	48.840	64.561	102.35	42.407

*Insignificant at 95% Confidence Level

TABLE 5.14

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYREB and ADMREB
26,141 Positive Income Families, Pooled 1966-71 Sample Data
(t Values in Parentheses)

Independent Variables	DAYREB Per Year Given Hospitalization			ADMREB Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.25801 (-4.07)	-0.24051 (-2.39)		0.05271 (14.52)	0.06061 (10.50)
D68.S			0.02576* (.15)			-0.02028 (-2.09)
D69.S			-0.07182* (-.41)			-0.00953* (-.94)
D70.S			-0.10500* (-.60)			-0.00833* (-.83)
D71.S			0.01786* (.10)			-0.00942* (-.94)
C1	-0.38880* (-1.14)			0.04195 (2.15)		
C2	-1.34994 (-3.78)			0.10129 (4.95)		
C3	-1.17716 (-2.92)			0.14841 (6.43)		
C4	-0.67994* (-1.47)			0.22488 (8.48)		
C5	-0.03285* (-.07)			0.33863 (12.52)		
AGE (A)		0.27966 (41.74)	0.28039 (23.58)		-0.00011* (-.29)	-0.00016 (-.24)
D68.A			0.02532* (1.26)			-0.00093* (-.81)
D69.A			0.01737* (.87)			0.00023* (.20)
D70.A			-0.01470* (-.74)			0.00127* (1.11)
D71.A			-0.03229* (-1.64)			-0.00020* (-.18)
A1	1.62925 (3.77)			0.06922 (1.79)		
A2	2.97462 (6.48)			-0.02335* (-.89)		
A3	5.28525 (11.82)			-0.04037* (-1.58)		
A4	8.92840 (19.54)			-0.00255* (-.10)		
A5	12.62004 (24.69)			0.04806* (1.64)		
A6	17.18431 (30.66)			0.07077 (2.20)		

TABLE 5.14 - Continued

Independent Variables	DAYREB Per Year Given Hospitalization			ADMREB Per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	4.06796 (10.24)	4.34304 (10.81)	4.30405 (10.71)	0.02132* (.94)	0.07721 (3.36)	0.07433 (3.23)
SINGLE FEMALE	-0.84719 (-2.02)	-0.66997* (-1.58)	-0.68489* (-1.61)	-0.15218 (-6.35)	-0.10177 (-4.19)	-0.10437 (-4.29)
URBAN	-0.16453* (-.73)	0.01215* (.05)	0.01295* (.06)	0.12956 (10.07)	0.13719 (10.72)	0.13633 (10.64)
WELFARE	6.08364 (12.06)	6.24698 (12.38)	5.84054 (8.15)	0.41064 (14.21)	0.40566 (14.04)	0.43128 (10.51)
INCOME	-0.10730 (-5.17)	-0.14284 (-6.98)	-0.13954 (-3.74)	-0.00461 (-3.87)	-0.00561 (-4.79)	-0.00814 (-3.82)
W.Y			0.41068 (2.00)			0.01707* (1.45)
D68.Y			0.00244* (.04)			0.00089* (.25)
D69.Y			0.01690* (.28)			0.00803 (2.40)
D70.Y			-0.04819* (-.81)			0.00400* (1.18)
D71.Y			-0.01214* (-.21)			0.00169* (.51)
D68.W.Y			-1.18489 (-3.40)			-0.04780 (-2.78)
D69.W.Y			-0.57127* (-1.82)			-0.04029 (-2.24)
D70.W.Y			-0.00811* (-.03)			-0.01938* (-1.11)
D71.W.Y			-0.21355* (-.84)			-0.02297* (-1.58)
D68	-0.46459* (-1.48)	-0.43669* (-1.39)	-1.48112* (-1.19)	-0.01805* (-1.00)	-0.01766* (-.98)	0.09896* (1.39)
D69	-0.66032 (-2.09)	-0.63910 (-2.02)	-1.18535* (-.92)	-0.02792* (-1.54)	-0.02627* (-1.45)	-0.04701* (-.64)
D70	-0.87094 (-2.77)	-0.85692 (-2.72)	0.49658* (.38)	-0.00232* (-.13)	-0.00126* (-.07)	-0.05095* (-.69)
D71	-1.82204 (-5.77)	-1.79152 (-5.68)	-0.28988* (-.23)	-0.02255* (-1.25)	-0.02204* (-1.22)	0.01519* (.20)
constant	9.32744	2.49846	2.34223	1.21576	1.12906	1.11495
R ²	0.09299	0.09125	0.09216	0.02580	0.02457	0.02546
S.E.E.	16.93112	16.94446	16.94146	0.96982	0.97026	0.97013
F.	133.900	238.514	94.673	34.589	59.84	24.367

*Insignificant at 95% Confidence Level

equations, as compared with three per cent for the corresponding regional-base equation. In contrast, and contrary to expectation, the volume equation for the regional-base hospitals exhibits slightly greater explanatory power than the corresponding equation for the community hospitals.

The hypothesis of a relatively strong deterrent effect among users of community hospitals finds no support in the data. In sign, significance, and magnitude, the coefficients of D68 to D71 are similar across the two hospital settings. The length of stay coefficients implicit in these results indicate a reduction of approximately a half day or less in both community and regional-base hospitals during the copayment period. A perplexing exception, however, is the decline of 1.2 days in average stay of patients admitted to regional-base hospitals during 1971. Since 1971 was not a full copayment year, it is very doubtful that this result should be labelled a price effect. The introduction of interaction variables in Model 3 does not modify any of these conclusions.

Other notable results contained in the tables may now be summarized. The family composition variables continue to perform as they have in the analysis of previous sections. The C_i coefficients in the days equation are uniformly positive for community hospital equation but uniformly negative in the regional-base hospital equation. Short-stay paediatric admissions are of course more likely in large families than in small families, and this undoubtedly explains why the average length of stay implicit in these coefficients is negative. The age coefficients all imply a positive relation between length of stay and age of head which is consistent with use patterns of the geriatric

population. Similarly, the marital status coefficients imply average stays are higher among members of families where the spouse is absent.

The absence of cross-haul of urban patients to rural hospitals reduces the utility of the urban coefficients in the community equations. In the regional and base equations, however, the urban coefficients are more interesting. They indicate that urban patients in urban hospitals stay an average of one day less than rural patients in urban hospitals. Thus, an extra day of hospitalization is involved when the rural patient travels to a hospital other than the one in the immediate vicinity. This suggests that earlier discharge to alternative facilities may be indicated for system-wide efficiency when not contra-indicated by clinical considerations.

Income effects in these equations are weak, negative, and virtually identical in the community and regional-base equations. The welfare variable displays an interesting reversal in the implied length of stay estimates. That is, patients from families receiving social assistance stay nearly one day less in hospital if the admitting hospital is a community hospital, while stays are one and one-half days longer if the same type of patient is admitted to a regional-base hospital. It may well be the case that the "social admission" with its implied shorter length of stay is a more common feature of the case load in the community hospital.

Summary

In this chapter, hypotheses specific to "user" families were examined. These families were studied separately because they represent

a small minority in both the sample and the parent population. By distinguishing this group from the dominant group of "nonusers", and by subjecting observed variations in use within this sub-sample to detailed scrutiny, it has been possible to conduct more refined tests of the various principal and secondary hypotheses at issue in this study. The outcome of these "stage-two" experiments will now be summarized.

1. Taken together, family characteristics explain more of the variance in patient-days than the variance in admissions. This finding obtains across "broad" and "narrow" definitions of use, and indirectly confirms the key role played by average length of stay in the empirical definition of the "high use" family.

2. The volume of inpatient care is only weakly related to the number of dependent children in the family. Large families tend to have more admissions than do small families, but their patient-day totals are not proportionately higher because paediatric cases involve atypically short stays. The magnitude of these effects varies widely across the different definitions of use, especially those that are diagnostic specific.

3. Age and marital status of the family head play a vital role in predicting both broadly and narrowly defined use. In general, admissions and patient-days are greater, and average lengths of stay longer, the older the head, as health status and other epidemiological considerations would suggest. One-spouse families, especially those headed by a single-widowed or divorced male, also tend to have longer lengths of stay. This finding illustrates the importance of domestic resources in the recovery process, and thereby reinforces the economic case for organized home care programs.

4. Rural families are observed to have markedly shorter lengths of stay than urban families. This could be due to the technology of care in rural hospitals where complicated conditions tend not to be treated, but its consistency across use definitions strongly suggests that a "bed availability" effect is present as well. Occupancy rates of 50-60 per cent are common in the province's small rural hospitals, and the presence of unfilled beds could well result in admission criteria being liberalized to the point where conditions not normally requiring hospitalization comprise a significant portion of the case load.

5. Families at various income levels above the "poverty" line (stringently defined by social assistance status) consume virtually identical quantities of inpatient care. Averaged over all use definitions, income elasticities of admissions, patient-days, and lengths of stay are only -0.03, -0.07, and -0.04 respectively. Negative income effects of this magnitude provide only weak support for the conventional "low-income/poor-health/high-use" hypothesis.

6. More compelling support for the above hypothesis is provided by the data on inpatient care received by families in receipt of social assistance. Patient-days, admissions, and lengths of stay are substantially greater for these families than they are for families located elsewhere in the income distribution. This has the important policy implication that public programs more ambitious than universal health insurance are necessary to solve the pervasive health and social problems of the seriously disadvantaged.

7. Finally, there is some evidence that both volume of care and duration of stay were lower during the copayment years, especially 1969 and 1970, than during the pre-copayment years 1966 and 1967. The

copayment coefficients in the broadly defined volume equation indicate 6.2 per cent reduction in utilization over the pre-copayment years, largely due to reductions in length of stay. These reductions were distributed unevenly across "narrowly" defined case types with effects ranging from 0 to 18%. Curiously, the largest declines are not confined to categories where discharge decisions are a priori more discretionary. There are two ways to resolve this apparent conflict between fact and theory. One is to reject the "price effect" interpretation and argue that the coefficients are capturing some unspecified factor other than copayment. The other is to accept the interpretation and argue that the distribution of clinical discretion across use categories is other than has been assumed.

CHAPTER VI

SUMMARY AND CONCLUSIONS

The principal purpose of this study was to develop empirical estimates of the effects of Saskatchewan's copayment policy on families' use of inpatient hospital services. Saskatchewan was selected for study since it was the ~~only~~ Canadian province where these effects could be isolated from the more general price effects associated with the introduction of public universal hospital insurance. The Saskatchewan Hospital Services Plan, the first program of its kind in North America, provided full first dollar coverage of hospital bills to virtually all residents of the province for twenty-one years prior to the introduction of user charges in 1968.

In Chapter 1, the key provisions of the insurance program prior to and during the copayment policy were discussed, as were other essential features of the Saskatchewan setting. In addition, the performance of the program in the pre-copayment period 1947-67 was briefly reviewed to provide a perspective on the public debate which attended the government's decision to implement per diem utilization fees as a "cost-containment" device. The flavour of this debate was illustrated with reference to statements made by the principal contestants, statements which revealed their political differences on the "user-pay" principle, their divergent assumptions on the nature and extent of "abuse" in a system of "free" care, their opposing judgments on the probable impact of copayment across different income groups and, above all, their unavoidable resort to speculation and assertion in the

absence of the relevant facts. Satisfactory answers to the key question, "are there 'price effects' in hospital use?" were not available at that time.

Previous attempts to quantify price effects in hospital use were critically reviewed in Chapter 2. The evidence from these studies required very cautious interpretation. It was noted that "most investigators described their results in terms of demand elasticities as if the observed price-use relationships neatly and necessarily coincided with "true" demand curves. The absence in all studies of convincing argument or evidence that care quantities are determined either by sovereign consumers or by physicians whose sole objective is to maximize the welfare of non-sovereign consumers militated strongly against claims of this sort. It was also noted that methodological problems and data deficiencies undermined the estimated price effects, even when they were neutrally interpreted as price elasticities of use. Most vulnerable to this criticism were those studies where the search for price effects was conducted in macro-data sets. The case for further interrogation of these data was seriously questioned. Research based on micro-data sets was judged marginally superior, but investigators who opted for this strategy also encountered problems, particularly in operationally defining the relevant net price variable, i.e., the "out-of-pocket" price per unit confronting the consumer-patient. The endemic difficulty of identifying these prices in the U.S. private hospital insurance data forced most investigators to adopt some "pseudo-price" measure which compromised both the logic of their exercise and the quality of their resultant elasticity estimates. These estimates were further impaired by the widely assumed tendency of U.S. families to purchase health

insurance policies providing benefits (net prices) consistent with their expectations of illness and attendant demands for care. Known in the insurance literature as "advance selection," this phenomenon doubtless produced artifactually large negative price elasticities in the majority of studies reviewed. However, based on the results presented in one of the more carefully executed studies, it was tentatively concluded that utilization decisions are only weakly sensitive to variations in the net price, certainly less sensitive than some (U.S.) health economists apparently believe. It was also made abundantly clear throughout the review that further research, and in particular, disaggregate analyses of quality micro-data ran no risk of being redundant.

In Chapter 3 the micro-analytic framework was presented and discussed. The elements of this framework were set out in terms of the available data, the methods by which these data were sampled, the representativeness of the sample, the techniques of data assembly and organization, and the model via which the data were analyzed. The available data were discussed in the context of the information system maintained by the Saskatchewan Hospital Services Plan. The two basic data files in this system, one a master registry of program beneficiaries, and the other a claims file of the inpatient experiences of these beneficiaries were described. These files were accessed on a sampling basis with the individual family as the unit of selection. Ten per cent random samples of the insured population in each of the years 1966-71 were drawn from the Master Registration File. These six annual cross-sections, each of which contained approximately 38,000 families, were examined in several dimensions and found to be representative of their respective populations. Subsequently, the inpatient data for

these samples were obtained and a family summary record drawing together the pertinent data from the registration and hospitalization records was constructed for each family. Included was information on all the variables judged relevant to modelling the effect of copayment on families' access to and use of inpatient facilities.

A model was developed to exploit the essential richness of the available data. Three basic equations formed the heart of the model. Each equation involved a unique concept and measure of hospital use, and thereby a unique implicit phrasing of the central price-theoretic question. The first equation focussed upon the probability of hospitalization and asked whether this probability was reduced when the user price was increased. The second and third equations focussed upon the volume (patient-days) and frequency (admissions) of hospitalization respectively and asked whether these indicators of utilization were reduced when the user price was increased. In addition, these two equations permitted an examination of the relationship between user charges and the average duration of stay per admission. Each of the basic equations was specified to include not only the user price of inpatient care, but also a vector of socio-demographic variables. These latter variables were inserted to control for some obvious non-price effects and, as well, to test a number of secondary hypotheses concerning inter-family variation in hospitalization experiences.

Consistent with the argument developed in Chapter 2, data disaggregation was an integral part of the modelling strategy. The rationale for disaggregation was to avoid the assumption, implicit or explicit in many previous studies, that all admission and discharge decisions are equally (non)discretionary and, hence, equally

(in)sensitive to point-of-service charges on inpatients. Difficulties associated with objective determination of the degree of discretion in utilization decisions dictated the adoption of a pragmatic scheme featuring alternative "narrow" definitions of the hospitalization event. These definitions preserved selected factual aspects of the event, including the diagnostic and treatment status of the patient, the practice setting and specialty of the admitting physician, and the size, location, and service capability of the host active-treatment institution. Various measures of utilization within each of these definitions were generated and they became the vehicles via which the search for differential price effects was conducted.

The statistical analogue of the basic three equation model required the estimation of two distinct types of regression equations. First, linear probability functions with dichotomous dependent variables distinguishing user from non-user families, were fitted to all the observations in the pooled annual samples. Designed to accommodate the large number of families whose members did not experience hospitalization during any given study year, these regressions generated probabilities of admission before and after copayment for families having specified socio-demographic and economic characteristics. Second, linear utilization functions with continuous dependent variables, alternatively defined in terms of patient-days and admissions, were fitted to sub-sets of pooled observations restricted to those families whose members experienced some hospitalization during any given year. These equations generated estimates of the volume and frequency of hospitalization before and after copayment of "user" families with specified characteristics. Corresponding estimates of the

duration of hospitalization (i.e., average length of stay) were obtained indirectly from the coefficients of these equations.

The linear probability and linear utilization functions were estimated for each "broad" and "narrow" definition of the hospitalization event. Probabilities of admission for each definition of the event were obtained by inserting the appropriate binary dependent variable in the linear probability function. Similarly, estimates of the volume and frequency of hospitalization for each definition of the event were obtained by inserting the appropriate continuous dependent variable in the linear utilization function.

The vector of independent variables was also varied to accommodate different assumptions concerning the nature and stability of the structural relationships. In the first specification, known as "model one," allowance was made for non-linearities in the relationships between the probability, volume, frequency and duration of hospitalization on the one hand, and the age of the family head and family size on the other. This was accomplished by inserting categorical age and size variates in the various estimating equations. The second specification, known as "model two," assumed linear (i.e., constant incremental) age and size effects and accordingly inserted continuous age and size variates in the various estimating equations. The third specification, known as "model three," featured interaction effects between copayment and selected family characteristics. In particular, dummy variables distinguishing the copayment from non-copayment years were multiplicatively combined with the age, size, income, and social assistance variates to form a number of first and second order interaction terms. This specification permitted tests of

the hypotheses that "old" families, "large" families, and "poor" families were those whose utilization was most affected by the introduction of the copayment policy. These three forms of the linear probability and linear utilization functions were estimated for each "broad" and "narrow" definition of the hospitalization event.

All the equations were estimated via the ordinary least squares regression method. The results of these experiments were presented and analyzed in Chapters 4 and 5. A summary of the major findings is presented in the following section.

Major Findings

The major findings of this study are as follows:

1. Copayment in the form of per diem direct charges did not significantly reduce the probability of "broadly" defined hospitalization. The probability was lower in one copayment year (1969) than in any of the other copayment years, but the proportionate difference was only 3.2 per cent.

2. Some apparent copayment effects were identified in the analysis of "narrowly defined" hospitalization probabilities. The likelihood of pregnancy related admission declined during the copayment years, but this was attributed to contemporaneous declines in the birth rate. Respiratory conditions exhibited the largest declines of any of the diagnostic categories, ranging from 17.3 per cent in 1969 to 3.7 per cent in 1968. Evidence that probabilities of admission by group general practitioners were marginally lower (by .01 basis points) in three copayment years (1969-71) initially indicated that a price effect

occurred in the "wrong" practice setting but further analysis suggested that this was a "supply effect." Finally, small declines of 0.5 to 3.0 per cent were registered in the probabilities of admission to community hospitals.

3. Declines of approximately 6.2 per cent in the volume and duration of "broadly defined" hospitalization occurred during the copayment years. However, when these effects were further investigated by examining the experience of a control group of socially assisted families who were exempted from the copayment charges, the evidence strongly suggested that some factor(s) other than copayment was responsible for the observed declines in use.

4. Only three of the "narrowly defined" diagnostic categories of use exhibited declines in the volume and duration of hospitalization. Circulatory, digestive and genito-urinary conditions registered declines ranging from 7 to 18 per cent, but again an analysis of a control group raised serious doubts that these declines were induced by copayment charges. Expectations that utilization would be more price-sensitive in small, "low" technology hospitals located in rural areas than in large, "medium-to-high" technology hospitals located in urban areas were not substantiated. In addition, decisions by solo and group general practitioners in respect of "user" families were essentially similar in their insensitivity to copayment charges.

5. The effect of copayment on utilization was not systematically related to the socio-demographic and economic characteristics of the family. For all measures and definitions of hospitalization there was no evidence to support the hypothesis that "deterrent effects" were concentrated among "old" families, "large"

families, or "poor" families. The coefficients of the interaction variables employed to test this hypothesis were uniformly small and, in the majority of cases, statistically insignificant.

6. The probability of hospitalization, and hence the probability of "illness taxation," varied systematically with the socio-demographic characteristics of the family. In general, the older the family head and the larger the family size, the greater were these probabilities. Family residence was also highly significant: rural families faced an estimated twenty-five per cent higher probability of "broadly" defined hospitalization than urban families.

7. The volume, frequency, and duration of hospitalization, and hence the magnitude of "illness taxation," varied systematically with the socio-demographic characteristics of "user" families. Utilization was especially sensitive to the age of the family head. Thus, it was estimated that the volume of "broadly" defined hospitalization for a family headed by an individual in the age group 65-74 years was twice that for a family headed by an individual in the age group 25-34 years. Assuming both families were otherwise identical, the former experienced approximately twenty days of hospitalization and incurred annual user charges of fifty dollars, while the latter experienced approximately ten days of hospitalization and incurred annual user charges of twenty-five dollars.

8. Families in receipt of social assistance were an identifiable "high risk," "high use" group. The likelihood of their experiencing hospitalization (at least once) during any given year was estimated to be forty per cent higher than that for socially unassisted families of identical size and composition. The "net welfare effect"

was also highly significant among hospitalized families. For "broadly" defined use, socially assisted families experienced (.6) more admissions and (1.1 days) longer average lengths of stay than their unassisted counterparts; in combination, these frequency and duration differentials resulted in a volume differential of approximately six and one-half patient-days, a proportional difference of nearly eighty per cent. It follows that had socially assisted families not been exempted from payment of the utilization fees, their "out-of-pocket" costs would also have been eighty per cent higher than those incurred by socially unassisted families.

9. The probability, volume, frequency, and duration of hospitalization were only weakly sensitive to variations in the income of socially unassisted families. Income elasticities for probabilities of admission were generally negative, statistically significant, and very close to zero. For "broadly" defined hospitalization, the elasticity estimate (-0.05) implied that the proportional difference in the admission probability between a family with an income of six thousand dollars and a family with an income of twelve thousand dollars was only five per cent. Corresponding differences in the volume, frequency, and duration of hospitalization were only seven per cent, three per cent, and four per cent respectively. Stronger inverse relationships (larger negative income elasticities) were expected on the assumptions that income and health status are positively correlated, and that income is reduced by hospitalization. Nonetheless, these results have two extremely important and quite straightforward interpretations. First, families at widely disparate income levels face nearly equal probabilities of hospitalization and, in the event of hospitalization,

receive nearly equal quantities of publicly insured inpatient care. Second, "illness taxation" is markedly regressive: the "representative user family" (urban residence, no children, male head aged 55 years) with an income of six thousand dollars in 1970 incurred annual copayment charges of approximately \$32.25 (\$2.50 x 12.9 days), while the equivalent family with an income of twelve thousand dollars incurred annual copayment charges of \$30.25 (\$2.50 x 12.1 days). That the user charges of these families differed by only 6.6 per cent while their incomes differed by 100 per cent clearly indicates that the principle of vertical equity (unequal taxation of unequals) was very poorly served by copayment.

Policy Implications

The findings of this study provide no compelling support for the hypothesis that direct charges on inpatients, equivalent to a coinsurance rate of 5 to 7 per cent, reduce hospital utilization under a public universal hospital insurance program. The analysis revealed that declines on the order of 6.2 per cent in broadly defined utilization did occur in the copayment years, declines traced to reduced lengths of stay in several of the long stay diagnostic categories, but it was not at all clear that these declines were triggered by the copayment policy. Indeed examination of time trends in utilization, together with analysis of a control group, strongly suggested that the declines were unrelated to copayment.

Arguments for point-of-service pricing of inpatient care as a "cost-containment" policy should be re-examined. In Saskatchewan, the

user charges of \$2.50 per day "saved" the government, by its own estimates, \$4.5 million dollars annually. However, these sums were merely re-routed from public to private budgets. Moreover, the analysis in this study revealed that the charges were distributed across the user population in a highly regressive manner.

If the observed declines in use over the period 1968-71 are interpreted as price effects, the experiment may be judged to have realized a further saving of \$3.2 million annually in the form of reduced resource costs. Of course, on this reasoning it is a matter of considerable speculation whether the reductions in use were medically appropriate. The analysis revealed that such reductions as did occur were confined to diagnostic groups which may reasonably be considered to involve the least patient discretion.

Voices continue to be heard in favour of user charges as a cost-containment strategy. In 1975, the President of the Canadian Medical Association called for the introduction of deterrent fees to restrain the utilization and costs of publicly insured hospital and medical care in Ontario.¹ More recently, the Ontario Economic Council made public its belief that "some portion of the costs should be borne by patients and these should be related to the patient's use of health care services."² Mindful of the regressiveness of a flat copayment charge, it outlined a scheme featuring, inter alia, deductibles and coinsurance charges of varying magnitude depending on family income. Hypothetical user taxes payable by families would, under various

¹ Toronto Star, February 19, 1975.

² Ontario Economic Council, Issues and Alternatives 1976, (Toronto: The Council, 1976), p. 15.

assumptions, range from \$50 to \$1575.³ Since these sums greatly exceed those levied during the Saskatchewan experiment, and since they may be in the range where utilization decisions are hypothetically price-sensitive, they might precipitate significant deterrent effects and sizeable reductions in total health expenditure. However, the appropriateness of such reductions in medical terms is far from obvious. Indeed, in respect of the existing "free" care system, the Council itself asks the key rhetorical question "Where are the statistical data that show people do, in fact, use health care services unnecessarily?"⁴

Meanwhile, current provincial governments may be well-advised to continue to pursue cost-containment by alternative strategies. Such strategies now include a combination of: budgetary constraints on hospital operating costs; a moratorium on the construction of active treatment facilities; and a promotion of alternative modes of care delivery featuring less reliance on the acute care institution and, concomitantly, greater reliance on relatively low cost extended care, home care, and other community based programs. All this supply-side activity seems eminently reasonable, and a sounder social and economic alternative to direct demands upon the pocket of the consumer-patient.

³ Ibid., Tables 1-3, pp. 45-48.

⁴ Ibid., p. 13.

APPENDICES

APPENDIX A

SELECTED HOSPITAL INSURANCE AND
OTHER PROVINCIAL STATISTICS

TABLE A.1

GENERAL HOSPITAL CARE OF

MALE, FEMALE AND TOTAL SHSP POPULATIONS: 1947-1976

Year	SEPARATIONS			PATIENT-DAYS			AVERAGE		
	1,000 Insured			per 1,000 Insured			DAYS OF STAY		
	Male	Female	Total	Male	Female	Total	Male	Female	Total
1947	113	203 (132)	156	1210	1953	1565	10.7	9.6	10.0
1948	135	225 (156)	178	1522	2261	1875	11.3	10.0	10.5
1949	156	247 (175)	200	1717	2404	2048	11.0	9.7	10.3
1950	159	252 (180)	203	1839	2584	2197	11.6	10.3	10.8
1951	153	247 (174)	199	1870	2557	2201	12.2	10.3	11.1
1952	159	255 (180)	205	1832	2544	2175	11.5	10.0	10.6
1953	160	255 (178)	206	1794	2512	2139	11.2	9.8	10.4
1954	156	255 (176)	204	1781	2409	2084	11.4	9.4	10.2
1955	154	250 (173)	201	1733	2387	2049	11.2	9.5	10.2
1956	156	252 (176)	202	1779	2441	2099	11.4	9.7	10.4
1957	165	259 (184)	211	1830	2429	2120	11.1	9.4	10.1
1958	163	251 (177)	205	1792	2311	2043	11.0	9.2	10.0
1959	166	257 (181)	210	1846	2352 (1873)	2091	11.1	9.2 (10.3)	10.0
1960	171	259 (184)	214	1845	2359 (1900)	2094	10.8	9.1 (10.3)	9.8
1961	169	255 (182)	211	1821	2288 (1848)	2048	10.8	9.0 (10.2)	9.7

TABLE A.1 - Continued

Year	SEPARATIONS			PATIENT-DAYS			AVERAGE		
	per 1,000 Insured			per 1,000 Insured			DAYS OF STAY		
	Male	Female*	Total	Male	Female*	Total	Male	Female*	Total
1962	168	251 (180)	208	1801	2232 (1808)	2010	10.7	8.9 (10.0)	9.7
1963	176	261 (191)	218	1867	2326 (1920)	2090	10.6	8.9 (10.0)	9.6
1964	181	265 (196)	222	1908	2353 (1957)	2125	10.5	8.9 (10.0)	9.6
1965	180	260 (199)	219	1879	2319 (1963)	2094	10.4	8.9 (9.9)	9.6
1966	179	251 (194)	214	1851	2232 (1907)	2038	10.3	8.9 (9.8)	9.5
1967	177	242 (189)	209	1837	2151 (1850)	1992	10.4	8.9 (9.8)	9.5
1968	173	245 (191)	208	1777	2143 (1849)	1957	10.3	8.7 (9.7)	9.4
1969	176	247 (195)	211	1821	2161 (1879)	1989	10.3	8.7 (9.6)	9.4
1970	186	257 (207)	221	1869	2233 (1960)	2049	10.0	8.7 (9.5)	9.3
1971	195	266 (215)	230	1909	2278 (2006)	2092	9.8	8.6 (9.3)	9.1
1972	198	264 (215)	231	1927	2259 (1999)	2092	9.7	8.6 (9.3)	9.1
1973	190	256 (207)	223	1818	2156 (1905)	1985	9.6	8.4 (9.2)	8.9
1974	186	247 (198)	216	1804	2062 (1811)	1932	9.7	8.4 (9.1)	8.9
1975	184	245 (197)	214	1763	2072 (1828)	1916	9.6	8.5 (9.3)	8.9
1976	178	238 (187)	208	1680	1969 (1713)	1823	9.4	8.3 (9.2)	8.8

*Figures in brackets exclude pregnancy related care.

Source: Saskatchewan Hospital Service Plan, Annual Reports.
(Regina: Queen's Printer, various issues).

TABLE A.2

SASKATCHEWAN HOSPITAL SERVICES PLAN, SELECTED
OPERATING STATISTICS FOR VARIOUS YEARS, 1947-74

YEAR	PATIENT-DAYS PER 1000 POP.*	SEPARATIONS PER 1000 POP.*	AVERAGE PAYMENT PER PATIENT-DAY*	AVERAGE PAYMENT PER SEPARATION*
1947	1565	156	\$ 4.97	\$ 49.75
1951	2201	199	7.64	84.61
1956	2099	202	12.16	126.20
1961	2048	211	17.74	172.38
1962	2010	208	19.15	184.98
1963	2090	218	19.33	185.56
1964	2125	222	20.98	200.94
1965	2094	219	23.15	221.14
1966	2038	214	26.25	249.68
1967	1992	209	28.69	273.74
1968	1957	208	31.00	291.21
1969	1989	211	33.68	317.49
1970	2049	221	36.19	334.91
1971	2092	230	39.45	358.33
1972	2092	231	44.71	404.96
1973	1985	223	49.99	445.90
1974	1932	216	61.54	549.24

*Excludes newborn care and services provided by Geriatric, Rehabilitation and Extended Care Hospitals; includes out-of-province hospitalization.

Source: Saskatchewan Hospital Services Plan, Annual Reports, 1947-1974 (Regina: Queen's Printer).

TABLE A.3

PERCENTAGE DISTRIBUTION OF POPULATION BY URBAN
SIZE GROUPS, RURAL NON-FARM AND FARM, CANADA
AND SASKATCHEWAN, 1966, 1971

	1966		1971	
	% of Total		% of Total	
	Saskatchewan	Canada	Saskatchewan	Canada
<u>Urban</u>				
Over 100,000	25.8	47.3	28.7	47.5
30,000-99,000	3.4	8.9	3.4	9.0
10,000-29,999	7.2	5.8	8.1	8.1
5,000-9,999	2.8	3.5	2.9	3.9
2,500-4,999	4.1	4.1	4.2	3.9
1,000-2,499	5.4	4.0	5.6	3.7
Total	49.0	73.6	53.0	76.1
<u>Rural</u>				
Farm	29.2	9.6	25.2	6.6
Non-Farm	21.7	16.9	21.8	17.3
Total	50.9	26.4	47.0	23.9
Total Number	955,344	20,104,880	926,245	21,568,310

Source: Canada, Dominion Bureau of Statistics, 1966 Census of Canada, Population Age Groups, I (1-10), March 1968.

Canada, Statistics Canada, 1971 Census of Canada Population Age Groups, I (1.2-3), April 1973.

TABLE A.4

PERCENTAGE DISTRIBUTION - POPULATION BY FIVE-YEAR AGE GROUPS AND SEX, SASKATCHEWAN AND CANADA, 1966

Age Group	Saskatchewan			Canada		
	% of Total Population	Total	Male	% of Total Population	Total	Female
0-4	11.3	3.8	5.8	11.0	5.6	5.3
5-9	11.5	5.9	5.7	11.5	5.9	5.6
10-14	10.8	5.6	5.3	10.5	5.4	5.1
15-19	9.3	4.7	4.6	9.2	4.6	4.5
20-24	6.5	3.3	3.2	7.3	3.6	3.7
25-29	5.5	2.8	2.7	6.2	3.1	3.1
30-34	5.5	2.8	2.7	6.2	3.2	3.1
35-39	5.8	3.0	2.8	6.4	3.2	3.2
40-44	5.8	2.9	2.9	6.3	3.1	3.2
45-49	5.6	2.8	2.8	5.4	2.7	2.7
50-54	5.2	2.7	2.5	4.9	2.5	2.4
55-59	4.4	2.3	2.1	4.1	2.1	2.0
60-64	3.6	1.9	1.7	3.3	1.6	1.7
65-69	2.9	1.5	1.4	2.7	1.3	1.4
70-74	2.5	1.3	1.2	2.1	1.0	1.1
75-79	2.0	1.1	.9	1.5	.7	.8
80-84	1.2	.7	.6	.9	.4	.5
85-89	.5	.3	.3	.4	.2	.2
90-94	.1	.1	.1	.1	.04	.06
95+	.0	.0	.0	.02	.00	.01
Total						
Number	955,344	489,040	466,304	20,014,880	10,054,344	9,960,536

Source: Canada, Dominion Bureau of Statistics, 1966 Census of Canada Population Age Groups, I (1-10), March 1968, Table 19.

TABLE A.5 .

PERCENTAGE DISTRIBUTION - POPULATION BY FIVE-YEAR AGE
GROUPS AND SEX, SASKATCHEWAN AND CANADA, 1971

Age Group	Saskatchewan			Canada		
	% of Total Population	Total	Male	% of Total Population	Total	Male
			Female			Female
0-4	8.5	4.4	4.2	8.4	4.3	4.1
5-9	10.7	5.4	5.2	10.5	5.3	5.1
10-14	11.0	5.6	5.4	10.7	5.5	5.2
15-19	10.3	5.3	5.1	9.8	5.0	4.8
20-24	7.4	3.8	3.6	8.8	4.4	4.4
25-29	5.7	2.9	2.8	7.3	3.7	3.6
30-34	5.1	2.6	2.5	6.1	3.1	3.0
35-39	5.2	2.7	2.6	5.9	3.0	2.9
40-44	5.5	2.8	2.7	5.9	3.0	2.9
45-49	5.6	2.8	2.8	5.7	2.8	2.9
50-54	5.4	2.7	2.7	4.9	2.4	2.5
55-59	5.0	2.5	2.5	4.4	2.2	2.2
60-64	4.2	2.1	2.0	3.6	1.8	1.8
65-69	3.4	1.7	1.6	2.9	1.4	1.5
70-74	2.5	1.2	1.3	2.1	1.0	1.2
75-79	2.0	1.0	1.0	1.5	0.6	0.9
80-84	1.4	0.7	0.7	1.0	0.4	0.6
85-89	0.7	0.4	0.4	0.5	0.2	0.3
90-94	0.2	0.1	0.1	0.1	0.1	0.1
95+	0.05	0.02	0.03	0.03	0.01	0.02
Total Number (000)	926.2	470.7	455.5	21,568.3	10,795.4	10,772.9

Source: Canada, Statistics Canada, 1971 Census of Canada
Population Age Group, I (1.2-3), April 1973.

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APPENDIX B

SELECTED CHARACTERISTICS OF
SASKATCHEWAN GENERAL HOSPITALS

TABLE B.1

RECOMMENDED DISTRIBUTION OF CLINICAL AND OTHER
SERVICES BY CLASS OF GENERAL HOSPITAL

<u>BASE HOSPITALS</u>	<u>REGIONAL HOSPITALS</u>	<u>COMMUNITY HOSPITALS</u>
Research	Major Surgery	Minor & Limited
Teaching	Internal Medicine	major surgery
Major Surgery	Obstetrics	General Medicine
Internal Medicine	Paediatrics	Obstetrics
Obstetrics	Orthopaedic Surgery	Eye, Ear, Nose, Throat
Orthopaedic Surgery	Eye, Ear, Nose, Throat	Dentistry
Eye, Ear, Nose, Throat	Psychiatric Service	X-Ray
Psychiatric Service	Communicable Disease	Laboratory
Geriatrics	Dentistry	Public Health
Cancer Clinic	Radiology	
Communicable Disease	Laboratory	
Dentistry	Physiotherapy	
Radiology	Teaching	
Laboratory	Dietetics	
Physical Medicine	Medical Social Service	
Physiotherapy	Public Health	
Rehabilitation		
Dietetics		
Medical Social Service		
Public Health		

Source: Saskatchewan Hospital Survey and Master Plan 1961, Part I
(Regina, Department of Public Health, 1962), p. 151.

TABLE B.2

DISTRIBUTION OF GENERAL HOSPITAL BY
TYPE, SIZE, AND POPULATION SERVED

HOSPITALS	Rated Bed Capacity		Population Served	
	1966	1971	1966	1971
COMMUNITY				
(a) 1-24 Beds				
Arborfield Union	8	8	1,165	1,066
Brock Union - Arcola	20	15	3,276	3,515
Beechy Union	5	8	1,188	783
Bengough Union	16	15	2,290	2,191
Bienfait - Coalfields Union	14	12	1,970	897
Big River Union	24	25	2,315	3,104
Birch Hills Memorial Union	15	15	2,807	3,353
Borden Union	10	10	2,548	1,031
Broadview Union	17	22	2,719	2,824
Cabri Union	20	20	2,351	1,830
Carrot River Union	20	18	2,729	2,162
Climax - Border Union	11	11	1,198	1,237
Coronach Union	12	12	1,215	1,101
Craik Community	6	8	1,116	1,004
Cudworth - St. Michael's	20	20	3,717	3,263
Cupar Union	13	13	2,736	2,692
Cut Knife Union	7	7	1,197	1,184
Davidson Union	18	17	2,803	2,320
Delisle Union	12	12	2,031	1,059
Dinsmore Union	10	10	1,36	1,310
Doddsland Union	10	10	1,209	966
The Eastend Union	13	13	1,565	1,921
Eatonia Union	15	12	1,152	1,174
Edam - Lady Minto Union	8	8	1,231	908
Elrose Union	14	10	1,123	783
Eston Union	24	24	4,050	3,291
Fillmore Union	18	15	2,344	2,484
Foam Lake Union	21	21	3,649	3,092
Frontier Union ^a	7	—	598	—
Gainsborough Union	15	14	2,524	2,037
Goodsoil Union	11	11	1,274	1,256
Grenfell Union	18	18	3,500	2,725
Gull Lake Union	16	16	1,739	2,104
Hafford Union	16	16	3,521	2,867
Herbert - More Union	22	19	2,981	1,923
Hodgeville Union ^b	9	—	168	—

TABLE B.2 - CONTINUED

HOSPITALS	Rated Bed Capacity		Population Served	
	1966	1971	1966	1971
Hudson Bay Union	14	18	1,606	2,097
Imperial Union	13	13	2,468	1,813
Invermay - Canora Union	11	11	1,523	1,366
Ituna Union	15	14	2,595	2,389
Kelvington Union	19	22	3,043	2,666
Kincaid Union	10	9	1,557	407
Kinistino Union	16	16	2,649	1,829
Kyle - White Bear Union	8	8	738	1,094
Lafleche Union	8	8	1,581	1,551
La Loche	6	15	780	852
Lampman Union	12	12	1,446	1,018
Lashburn & District Union ^c	11	—	1,689	—
Leoville Union	12	11	1,157	1,156
Leroy Union ^d	17	—	1,191	—
Lestock - St. Joseph's	23	23	4,105	4,130
Loon Lake Union	16	19	1,592	1,417
Lucky Lake Union	13	9	1,520	886
Maidstone Union	16	19	2,139	2,911
Mankota Union	13	10	1,361	1,062
Maryfield Memorial Union ^e	7	—	862	—
Midale Union	8	8	986	873
Milden Union	10	10	1,062	633
Montmartre Union	12	11	2,514	2,186
Mossbank-Sutton Lake Johnson Union ^f	16	—	1,557	—
Neilburg and District Union	12	12	1,577	1,677
Neudorf Union ^g	11	—	1,355	—
Nokomis Union	13	15	3,035	1,909
Norquay - Canora Union	15	15	2,122	1,872
Oxbow Union	21	21	3,377	2,936
Pangman Union	10	10	1,317	863
Paradise Hill Union	13	18	1,623	1,993
Ponteix Union	6	10	1,736	1,717
Porcupine Carragana Union	18	20	3,749	2,954
Prelate Union ^h	8	—	1,156	—
Qu'Appelle Union ⁱ	14	—	2,844	—
Quill Lake Union ^j	12	—	1,810	—
Rabbit Lake Union	10	10	1,058	N.A.
Radville Community	20	18	3,554	2,979
Redvers Union	24	21	3,392	3,058
Rockglen Union	7	7	1,273	1,071
Rose Valley Union	19	17	3,664	2,464
St. Walburg Union	21	20	2,359	2,160

TABLE B.2 - CONTINUED

HOSPITALS	Rated Bed Capacity		Population Served	
	1966	1971	1966	1971
Smeaton Union	11	11	1,062	359
Spalding Union	14	12	2,545	1,861
Spiritwood Union	19	20	2,651	2,533
Theodore Union	16	13	1,639	1,257
Turtleford - Riverside Memorial Union	20	26	2,808	3,113
Vanguard Union	17	11	1,750	762
Wakaw Union	24	26	4,868	4,551
Watson Union	16	14	1,315	1,864
Wawota Union	11	13	1,852	2,024
Whitewood	11	10	1,788	1,203
Willow Bunch Union ^k	9	—	1,152	—
Wolseley Memorial Union	21	19	2,610	2,278
Wynyard Union	21	21	3,531	3,804
Zenon Park - Notre Dame de l'Assomption	13	10	1,158	818

(b) 25-49 Beds

Assiniboia Union	32	42	6,086	6,764
Balcarres Union	34	31	5,016	4,661
Biggar Union	36	34	7,169	5,861
Central Butte Union	25	30	4,617	3,851
Esterhazy - St. Anthony's	30	30	5,210	5,313
Gravelbourg - St. Joseph's	32	32	4,906	5,032
Ile à la Crosse	36	46	2,487	3,209
Indian Head Union	28	23	3,843	4,613
Kamsack Union	49	49	6,636	6,404
Kerrobert Union	30	30	4,687	3,362
Kipling Memorial Union	27	27	4,020	4,124
Langenburg Union	31	29	4,492	2,610
Lanigan Union	11	30	2,313	5,156
Leader Union	34	25	3,755	3,492
La Ronge Hospital	27	28	2,969	3,290
Macklin - St. Joseph's	25	25	2,274	2,715
Maple Creek Union	44	44	5,992	4,86
Meadow Lake Union	46	46	6,284	6,793
Moosomin Union	48	50	7,231	7,233
Outlook Union	30	22	4,771	4,349
Preeceville Union	27	29	5,067	4,697
Rosetown Union	36	34	4,733	6,196
Rosethorn Union	36	36	6,930	6,328
Shaunavon Union	33	37	5,862	5,249

TABLE B.2 - CONTINUED

HOSPITALS	Rated Bed Capacity		Population Served	
	1966	1971	1966	1971
Shellbrook Union	30	30	4,563	4,546
Unity Union	28	28	4,473	4,479
Uranium City Municipal	26	32	2,321	2,566
Wadena Union	46	42	7,063	6,660
Watrous Union	36	32	4,751	3,385
Wilkie Union	31	31	3,421	2,739

(c) 50-99 Beds

Canora Union	66	50	8,375	6,977
Humboldt - St. Elizabeth's	75	81	12,982	11,811
Kindersley Union	35	55	7,264	6,573
Lloydminster	92	92	5,536	6,441
Melfort	85	85	11,836	11,271
Melville - St. Peter's	81	81	10,185	10,436
Nipawin Union	62	62	9,560	10,501
Visdale - St. Therese	75	75	9,393	7,673
Estevan	106	106	11,541	12,496
Weyburn	96	101	14,378	15,589

REGIONAL

Moose Jaw Union	254	241	26,749	27,778
Moose Jaw Providence	159	140	19,925	16,722
North Battleford - Battlefords Union	136	145	23,156	22,538
Prince Albert - Holy Family	150	152	19,179	20,619
Prince Albert - Victoria Union	180	228	20,594	20,907
Swift Current Union	157	148	25,332	25,156
Yorkton Union	193	291	25,213	28,094

BASE

Regina General	681	689	93,656	104,937
Regina - Grey Nuns'	421	430	69,175	68,006
Saskatoon City	309	376	53,798	62,158
Saskatoon - St. Paul's	333	336	53,593	63,057
Saskatoon University	550	546	56,955	56,304

Source: Saskatchewan Hospital Services Plan, Annual Reports
1966, 1971.

TABLE B.2 - CONTINUED

Notes: ^aClosed on July 1, 1969.
^bClosed on July 1, 1969.
^cClosed on November 1, 1969.
^dClosed on July 1, 1969.
^eClosed on August 1, 1969.
^fClosed on November 1, 1969.
^gClosed on July 1, 1969.
^hClosed on April 1, 1968.
ⁱClosed on March 1, 1969.
^jClosed on November 1, 1969.
^kClosed on June 1, 1969.

TABLE B.3

SUMMARY OF HOSPITALS BY BED SIZE
AS OF DECEMBER 31, 1971

GROUP	NUMBER OF HOSPITALS	NUMBER OF RATED BEDS	PERCENT OF TOTAL HOSPITALS	PERCENT OF TOTAL RATED BEDS
Group I (Under 10 beds)	11	88	8.15%	1.30%
Group II (10-24 beds)	69	1051	51.11%	15.48%
ACCUMULATED TOTALS - 24 BEDS AND UNDER	80	1139	59.26%	16.78%
Group III (25-49 beds)	31	1035	22.96%	15.25%
ACCUMULATED TOTALS - 49 BEDS AND UNDER	111	2174	82.22%	32.03%
Group IV (50-99 beds)	10	684	7.40%	10.08%
ACCUMULATED TOTALS - 99 BEDS AND UNDER	121	2858	89.62%	42.11%
Group V (100-199 beds)	6	792	4.44%	11.67%
ACCUMULATED TOTALS 199 BEDS AND UNDER	127	3650	94.06%	53.78%
Group VI (200 + beds)	8	3137	5.94%	46.22%
TOTAL	135	6787	100 00%	100 00%

Source: Directory of General Facilities in Saskatchewan,
(Saskatchewan Hospital Services Plan, December 31, 1971).

APPENDIX C

A COMPARISON OF RESULTS YIELDED BY THE TWIN LINEAR²
REGRESSION MODEL AND THE CLASSICAL
LINEAR REGRESSION MODEL

TABLE C.1

Parameter Estimates and Other Statistics^a for the Twin Linear Probability Model and for the Classical Regression Model:
Positive Income Families, Pooled 1966-71 Sample Data

Regressand	Twin Linear Probability Model			Classical Regression Model	
	BINDIS	DAYDIS	ADMDIS	DAYDIS	ADMDIS
Regressors:					
S	0.05904 (78.72)	0.05395* (1.14)	0.09979 (30.42)	0.68037 (35.71)	0.13525 (84.01)
A	0.00081 (11.57)	0.26305 (51.08)	0.00134 (3.72)	0.08287 (49.92)	0.00133 (9.50)
M	-0.10267 (29.42)	3.84711 (12.60)	0.06498 (3.09)	0.28106 (3.18)	-0.10601 (14.15)
F	-0.08031 (21.25)	-0.55096* (1.63)	-0.11537 (4.94)	-0.97429 (10.17)	-0.09533 (11.75)
U	-0.05704 (24.48)	0.91722 (5.28)	-0.20971 (17.49)	-0.47422 (8.04)	-0.14880 (29.82)
W	0.11218 (18.63)	6.64275 (16.75)	0.59963 (21.92)	3.88788 (25.48)	0.38247 (29.60)
Y	-0.00262 (10.92)	-0.14879 (8.70)	-0.00907 (7.69)	-0.08686 (14.15)	-0.00773 (14.87)
D68	-0.00209* (.63)	-0.41299* (1.70)	-0.02759* (1.65)	-0.19150 (2.27)	-0.01267* (1.78)
D69	-0.00743 (2.24)	-0.65128 (2.66)	-0.02231* (1.32)	0.24304 (2.89)	-0.02009 (3.10)
D70	0.00491* (1.47)	-0.93790 (3.85)	-0.01478* (0.88)	0.17437 (2.05)	0.00265* (0.37)
D71	0.00651* (1.95)	-1.39973 (5.75)	-0.02800* (1.67)	0.30521 (3.61)	0.00102* (0.14)
Constant	0.15722	1.14274	1.29591	-0.94676	0.16250
R ²	0.10111	0.07661	0.04960	0.03800	0.09500
S.E.E.	0.43036	16.950	1.16913	10.904	0.923
n	150,729	43,761	43,761	150,729	150,729

^aThe t statistics of the parameter estimate are given in parentheses.

*Insignificant at the 95 Confidence level.

APPENDIX D

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTIONS

TABLE D.1

Means and Standard Deviations of Selected Variables
in the Linear Probability Function: 24,218 Positive Income
Families, 1966 Sample Data

Variable	Mean	Standard Deviation
BINDIX	.3083	.4618
BINDIS	.3051	.4605
BINCIR	.0270	.1620
BINRES	.0791	.2699
BINDIG	.0590	.2357
BINGUR	.0397	.1953
BINDEL	.0636	.2440
BINAPV	.0404	.1969
BINSGP	.0790	.2698
BINGGP	.1664	.3724
BINCOM	.1504	.3575
BINREB	.1787	.3831
S	3.0141	2.0048
C1	.1335	.3402
C2	.1434	.3505
C3	.1002	.3002
C4	.0645	.2456
C5	.0667	.2496
A	43.5875	17.0028
A1	.1777	.3822
A2	.1877	.3905
A3	.1897	.3921
A4	.1468	.3539
A5	.0788	.2694
A6	.0477	.2131
M	.1798	.3840
F	.1330	.3396
U	.4368	.4960
W	.0247	.1551
Y (\$thousand)	5.3928	4.6196

TABLE D.2

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTION: 25,084 POSITIVE INCOME
FAMILIES, 1967 SAMPLE DATA

Variable	Mean	Standard Deviation
BINDTX	.2956	.4563
BINDIS	.2920	.4547
BINCIR	.0284	.1662
BINRES	.0703	.2556
BINDIG	.0618	.2408
BINGUR	.0359	.1861
BINDEL	.0539	.2258
BINAPV	.0376	.1903
BINSGP	.0718	.2582
BINGGP	.1635	.3699
BINCOM	.1459	.3530
BINREB	.1684	.3742
S	2.8138	1.9038
C1	.1277	.3337
C2	.1296	.3359
C3	.0890	.2848
C4	.0542	.2264
C5	.0529	.2239
A	43.6503	17.2325
A1	.1722	.3775
A2	.1836	.3872
A3	.1846	.3880
A4	.1491	.3562
A5	.0824	.2749
A6	.0486	.2151
M	.1953	.3965
F	.1352	.3419
U	.4348	.4957
W	.0321	.1761
Y (\$Thousand)	5.5926	4.7389

TABLE D.3

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTION: 25,423 POSITIVE INCOME
FAMILIES, 1948 SAMPLE DATA

Variable	Mean	Standard Deviation
BINDTX	.2901	.4538
BINDIS	.2864	.4521
BINCIR	.0269	.1618
BINRES	.0675	.2509
BINDIG	.0558	.2296
BINGUR	.0387	.1930
BINDEL	.0550	.2280
BINAPV	.0372	.1893
BINSGP	.0659	.2482
BINGGP	.1559	.3628
BINCOM	.1358	.3426
BINREB	.1731	.3784
S	2.7678	1.8732
C1	.1263	.3322
C2	.1279	.3340
C3	.0859	.2802
C4	.0534	.2248
C5	.0478	.2132
A	43.6419	17.4380
A1	.1726	.3779
A2	.1779	.3825
A3	.1804	.3846
A4	.1490	.3561
A5	.0848	.2785
A6	.0503	.2185
M	.2066	.4049
F	.1349	.3416
U	.4340	.4956
W	.0349	.1836
Y (\$Thousand)	5.5939	4.8817

TABLE D.4

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTION: 25,644 POSITIVE INCOME
FAMILIES, 1969 SAMPLE DATA

Variable	Mean	Standard Deviation
BINDTX	.2829	.4504
BINDIS	.2797	.4489
BINCIR	.0270	.1622
BINRES	.0603	.2380
BINDIG	.0579	.2336
BINGUR	.0380	.1912
BINDEL	.0520	.2221
BINAPV	.0383	.1920
BINSGP	.0692	.2538
BINGGP	.1473	.3544
BINCOM	.1325	.3390
BINREB	.1690	.3748
S	2.7406	1.8447
C1	.1290	.3352
C2	.1306	.3370
C3	.0848	.2786
C4	.0514	.2207
C5	.0462	.2100
A	43.8637	17.7070
A1	.1727	.3780
A2	.1711	.3766
A3	.1774	.3820
A4	.1472	.3544
A5	.0901	.2863
A6	.0545	.2270
M	.1842	.3876
F	.1549	.3618
U	.4555	.4980
W	.0389	.1933
Y (\$Thousand)	5.5286	4.9992

TABLE D.5

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTION: 24,850 POSITIVE INCOME
FAMILIES, 1970 SAMPLE DATA

Variable	Mean	Standard Deviation
BINDTX	.2961	.4565
BINDIS	.2912	.4543
BINCIR	.0278	.1645
BINRES	.0644	.2455
BINDIG	.0608	.2390
BINGUR	.0410	.1983
BINDEL	.0462	.2100
BINAPV	.0406	.1974
BINSGP	.0828	.2756
BINGGP	.1405	.3475
BINCOM	.1369	.3437
BINREB	.1775	.3821
S	2.7239	1.8257
C1	.1270	.3329
C2	.1309	.3373
C3	.0821	.2746
C4	.0497	.2173
C5	.0446	.2064
A	44.6601	17.8802
A1	.1666	.3726
A2	.1660	.3720
A3	.1799	.3841
A4	.1534	.3603
A5	.0971	.2962
A6	.0590	.2355
M	.1891	.3916
F	.1521	.3591
U	.4531	.4978
W	.0387	.1928
Y. (\$Thousand)	5.6649	5.1450

TABLE D.6

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTION: 25,510 POSITIVE INCOME
FAMILIES, 1971 SAMPLE DATA

Variable	Mean	Standard Deviation
BINDTX	.2924	.4549
BINDIS	.2884	.4530
BINCIR	.0279	.1647
BINRES	.0608	.2390
BINDIG	.0553	.2286
BINGUR	.0398	.1955
BINDEL	.0463	.2102
BINAPV	.0408	.1978
BINSGP	.0781	.2683
BINGGP	.1436	.3507
BINCOM	.1363	.3431
BINREB	.1742	.3793
S	2.6742	1.8006
C1	.1242	.3297
C2	.1295	.3358
C3	.0786	.2692
C4	.0477	.2131
C5	.0411	.1985
A	44.7641	18.0470
A1	.1641	.3704
A2	.1599	.3665
A3	.1786	.3830
A4	.1558	.3627
A5	.0996	.2995
A6	.0604	.2382
M	.2055	.4041
F	.1527	.3597
U	.4426	.4967
W	.0496	.2171
Y(\$Thousand)	6.2741	5.5151

TABLE D.7

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR PROBABILITY FUNCTION: 150,729 POSITIVE INCOME
FAMILIES, POOLED 1966-71 SAMPLE DATA

Variable	Mean	Standard Deviation
BINDTX	0.2941	0.4556
BINDLS	0.2903	0.4539
BINCIR	0.0275	0.1636
BINRES	0.0670	0.2500
BINDIG	0.0584	0.2345
BINGUR	0.0389	0.1933
BINDEL	0.0528	0.2236
BINAPV	0.0391	0.1940
BINSGP	0.0744	0.2624
BINGGP	0.1528	0.3598
BINCOM	0.1395	0.3465
BINREB	0.1734	0.3786
S	2.7873	1.8787
C1	0.1279	0.3340
C2	0.1319	0.3384
C3	0.0867	0.2814
C4	0.0534	0.2248
C5	0.0498	0.2174
A	44.0301	17.5067
A1	0.1709	0.3765
A2	0.1742	0.3793
A3	0.1817	0.3856
A4	0.1502	0.3573
A5	0.0889	0.2845
A6	0.0535	0.2249
M	0.1935	0.3951
F	0.1439	0.3510
U	0.4428	0.4967
W	0.0366	0.1877
Y (\$Thousand)	5.6714	5.0050

APPENDIX E

MEANS AND STANDARD DEVIATIONS OF SELECTED VARIABLES
IN THE LINEAR UTILIZATION FUNCTIONS

TABLE E.1

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYDIS and ADMDIS: 43,761 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYDIS	12.7077	17.6364
ADMDIS	1.5958	1.1991
SIZE (S)	3.6533	1.9943
C1	.1635	.3698
C2	.1812	.3852
C3	.1300	.3364
C4	.0897	.2857
C5	.0959	.2945
AGE (A)	45.1415	16.5115
A1	.2194	.4139
A2	.2191	.4136
A3	.1915	.3935
A4	.1329	.3395
A5	.0835	.2767
A6	.0613	.2399
SINGLE MALE	.0904	.2867
SINGLE FEMALE	.0764	.2656
URBAN	.3796	.4853
WELFARE (W)	.0473	.2123
INCOME (Y)	6.1091	5.0506

TABLE E.2

Means and Standard Deviations of Selected Variables in the Linear
Utilization Function with Alternate Regressands DAYCIR and ADMCIR:
4,147 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYCIR	15.8845	19.2648
ADMCIR	1.2141	.6654
SIZE (S)	3.0422	1.9540
C1	.1278	.3339
C2	.1080	.3105
C3	.0912	.2879
C4	.0591	.2358
C5	.0721	.2587
AGE (A)	57.7921	15.8301
A1	.0622	.2416
A2	.1452	.3523
A3	.1924	.3943
A4	.2170	.4123
A5	.1980	.3985
A6	.1693	.3750
SINGLE MALE	.1271	.3331
SINGLE FEMALE	.0909	.2875
URBAN	.3207	.4668
WELFARE (W)	.0555	.2289
INCOME (Y)	5.5852	5.4305

TABLE E.3

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYRES and ADMRES: 10,093 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYRES	6.7414	8.6530
ADMRES	1.2695	.6756
SI(S)	4.3432	2.0347
C1	.1295	.3358
C2	.2217	.4154
C3	.1797	.3840
C4	.1287	.3349
C5	.1460	.3532
AGE (A)	42.0075	14.6608
A1	.2763	.4472
A2	.2837	.4508
A3	.1799	.3841
A4	.0877	.2828
A5	.0531	.2243
A6	.0406	.1974
SINGLE MALE	.0748	.2631
SINGLE FEMALE	.0581	.2339
URBAN	.3408	.4740
WELFARE (W)	.0580	.2337
INCOME (Y)	6.1979	4.7198

TABLE E.4

Means and Standard Deviations of Selected Variables in the Linear
Utilization Function with Alternate Regressands DAYDIG and ADMDIG:
8,806 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYDIG	9.6715	10.1302
ADMDIG	1.1793	.5109
SIZE (S)	3.7244	2.0500
C1	.1561	.3630
C2	.1705	.3761
C3	.1275	.3336
C4	.0930	.2905
C5	.1081	.3105
AGE (A)	46.4945	16.1030
A1	.1874	.3902
A2	.2225	.4159
A3	.2093	.4068
A4	.1513	.3583
A5	.0948	.2930
A6	.0575	.2327
SINGLE MALE	.0854	.2795
SINGLE FEMALE	.0638	.2444
URBAN	.3250	.4684
WELFARE (W)	.0538	.2257
INCOME (Y)	5.9702	4.8920

TABLE E.5

Means and Standard Deviations of Selected Variables in the Linear Utilization Function With Alternate Regressands DAYGUR and ADMGUR: 5,857 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYGUR	8.6184	9.6915
AMDGUR	1.1289	.4242
SIZE (S)	3.7226	1.9528
C1	.1566	.3634
C2	.1858	.3889
C3	.1294	.3357
C4	.0890	.2847
C5	.1007	.3010
AGE (A)	47.6739	15.1519
A1	.1668	.3728
A2	.2307	.4213
A3	.2426	.4287
A4	.1661	.3722
A5	.0917	.2886
A6	.0553	.2286
SINGLE MALE	.0594	.2364
SINGLE FEMALE	.0637	.2442
URBAN	.3524	.4778
WELFARE (W)	.0411	.1986
INCOME (Y)	6.3329	5.1883

TABLE E.6

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regresssands DAYDEL and ADMDEL: 7,956 Positive Income Families, Pooled 1966-71 Sample Data

Variables	Mean	Standard Deviation
DAYDEL	6.6032	4.9902
ADMDEL	1.1913	.4904
SIZE (S)	4.1313	1.8335
C1	.2768	.4474
C2	.2231	.4164
C3	.1369	.3437
C4	.0933	.2908
C5	.1171	.3216
AGE (A)	31.3172	8.1775
A1	.4969	.5000
A2	.2231	.4164
A3	.0632	.2434
A4	.0104	.1016
A5	.0010	.0317
A6	.0004	.0194
SINGLE MALE	.0187	.1356
SINGLE FEMALE	.0480	.2138
URBAN	.4570	.4982
WELFARE (W)	.0500	.2180
INCOME (Y)	6.3143	4.2764

TABLE E.7

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYAPV and ADMAPV: 5,901 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYAPV	7.8387	14.6346
ADMAPV	1.1013	.3492
SIZE (S)	3.9931	2.2610
C1	.1224	.3277
C2	.1693	.3750
C3	.1376	.3445
C4	.1096	.3125
C5	.1525	.3596
AGE (A)	44.8009	15.8441
A1	.1825	.3863
A2	.2371	.4253
A3	.2276	.4193
A4	.1332	.3398
A5	.0620	.2412
A6	.0549	.2278
SINGLE MALE	.1307	.3371
SINGLE FEMALE	.0749	.2633
URBAN	.3225	.4675
WELFARE (W)	.0581	.2340
INCOME (Y)	5.8566	4.9294

TABLE E.8

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYSUR and ADMSUR: 20,680 Positive Income Families, Pooled 1966-71 Sample Data

Variables	Mean	Standard Deviation
DAYSUR	10.1394	13.1898
ADMSUR	1.1948	.5102
SIZE (S)	3.8467	2.0177
C1	.1394	.3463
C2	.1886	.3912
C3	.1494	.3565
C4	.1022	.3030
C5	.1089	.3115
AGE (A)	45.6098	14.9303
A1	.1908	.3929
A2	.2521	.4342
A3	.2250	.4176
A4	.1474	.3545
A5	.0766	.2660
A6	.0431	.2032
SINGLE MALE	.0764	.2656
SINGLE FEMALE	.0691	.2535
URBAN	.4242	.4942
WELFARE (W)	.0383	.1920
INCOME (Y)	6.5523	5.3873

TABLE E.9

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYMED and ADMED: 23,015 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYMED	12.0193	17.9768
ADMED	1.4660	1.0278
SIZE (S)	3.6414	2.0925
C1	.1425	.3495
C2	.1675	.3734
C3	.1253	.3310
C4	.0898	.2859
C5	.1079	.3103
AGE (A)	45.9987	17.2077
A1	.1791	.3835
A2	.2070	.4052
A3	.2005	.4004
A4	.1452	.3523
A5	.1046	.3060
A6	.0886	.2842
SINGLE MALE	.1131	.3168
SINGLE FEMALE	.0788	.2694
URBAN	.2841	.4510
WELFARE (W)	.0611	.2395
INCOME (Y)	5.6357	4.8083

TABLE E.10

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYSGP and ADMSGP: 11,216 Positive Income Families, Pooled 1966-91 Sample Data

Variable	Mean	Standard Deviation
DAYSGP	9.9142	12.8740
ADMSGP	1.4686	1.0094
SIZE (S)	3.7405	2.0505
C1	.1588	.3655
C2	.1764	.3812
C3	.1281	.3342
C4	.0968	.2957
C5	.1089	.3115
AGE (A)	45.9819	16.7530
A1	.2093	.4068
A2	.2141	.4102
A3	.1988	.3991
A4	.1335	.3401
A5	.0874	.2824
A6	.0720	.2586
SINGLE MALE	.0966	.2954
SINGLE FEMALE	.0698	.2548
URBAN	.2642	.4409
WELFARE (W)	.0465	.2105
INCOME (Y)	5.6291	4.4218

TABLE E.11

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYGGP and ADMGGP: 23,026 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYGGP	10.5629	16.2001
ADMGGP	1.4601	.9614
SIZE (S)	3.7881	2.0436
C1	.1618	.3683
C2	.1831	.3867
C3	.1351	.3418
C4	.0954	.2937
C5	.1098	.3126
AGE (A)	44.4050	16.5303
A1	.2327	.4226
A2	.2219	.4155
A3	.1851	.3884
A4	.1239	.3294
A5	.0786	.2691
A6	.0590	.2356
SINGLE MALE	.0858	.2801
SINGLE FEMALE	.0728	.2598
URBAN	.3147	.4644
WELFARE (W)	.0532	.2245
INCOME (Y)	5.8344	4.4753

TABLE E.12

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYCOM and ADMCOM: 21,029 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYCOM	10.1302	14.0253
AMDCOM	1.6161	1.1616
SIZE (S)	3.8100	2.1026
C1	.1520	.3590
C2	.1736	.3788
C3	.1312	.3376
C4	.0985	.2980
C5	.1186	.3233
AGE (A)	46.3064	16.5171
A1	.1979	.3984
A2	.2180	.4129
A3	.2089	.4065
A4	.1400	.3470
A5	.0874	.2824
A6	.0680	.2517
SINGLE MALE	.0966	.2955
SINGLE FEMALE	.0639	.2445
URBAN	.0456	.2086
WELFARE (W)	.0465	.2106
INCOME (Y)	5.3748	4.2402

TABLE E.13

Means and Standard Deviations of Selected Variables in the Linear Utilization Function with Alternate Regressands DAYREB and ADMREB: 26,141 Positive Income Families, Pooled 1966-71 Sample Data

Variable	Mean	Standard Deviation
DAYREB	13.0926	17.7711
ADMREB	1.3659	.9822
SIZE (S)	3.5735	1.9193
C1	.1688	.3746
C2	.1876	.3904
C3	.1298	.3360
C4	.0844	.2779
C5	.0835	.2766
AGE (A)	44.5509	16.4565
A1	.2322	.4222
A2	.2201	.4143
A3	.1812	.3852
A4	.1300	.3363
A5	.0815	.2736
A6	.0579	.2335
SINGLE MALE	.0847	.2784
SINGLE FEMALE	.0826	.2753
URBAN	.6102	.4877
WELFARE (W)	.0495	.2169
INCOME (Y)	6.6066	5.4795

APPENDIX F

PARAMETER ESTIMATES FOR THREE SPECIFICATIONS
OF THE LINEAR PROBABILITY FUNCTION

TABLE F.1

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINDTX: 150,729
Positive Income Families; Pooled 1966-71 Sample Data

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE (s)		0.05909 (77.75)	0.06247 (55.78)
D68.S			-0.00370 (-2.01)
D69.S			-0.00874 (-4.72)
D70.S			-0.00429 (-2.28)
D71.S			-0.00465 (-2.46)
C1	0.16660 (44.31)		
C2	0.19184 (48.08)		
C3	0.23395 (50.20)		
C4	0.28822 (51.56)		
C5	0.35822 (61.87)		
AGE (A)		0.00111 (15.86)	0.00080 (6.67)
D68.A			-0.00032* (-1.68)
D69.A			0.00070 (3.68)
D70.A			0.00084 (4.42)
D71.A			0.00054 (2.84)
A1	0.05788 (14.12)		
A2	-0.02172 (-4.96)		
A3	-0.02967 (-7.24)		
A4	0.00915 (2.25)		
A5	0.08436 (18.06)		
A6	0.19038 (34.61)		

TABLE F.1 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.12765 (-37.99)	-0.09741 (-27.75)	-0.09802 (-27.93)
SINGLE FEMALE	-0.11445 (-31.27)	-0.07904 (-20.80)	-0.08008 (-21.07)
URBAN	-0.06227 (-26.73)	-0.05497 (-23.49)	-0.05515 (-23.57)
WELFARE (W)	0.10874 (18.09)	0.11659 (19.27)	0.11605 (13.77)
INCOME (Y)	-0.00219 (-8.76)	-0.00284 (-11.83)	-0.00360 (-8.18)
W.Y			0.00097* (.34)
D68.Y			0.00109* (1.49)
D69.Y			0.00158 (2.23)
D70.Y			0.00125* (1.79)
D71.Y			0.00065* (.97)
D68.W.Y			0.00432* (1.03)
D69.W.Y			-0.00497* (-1.41)
D70.W.Y			-0.00201* (-.48)
D71.W.Y			0.00137* (.38)
D68	-0.00054* (-.16)	-0.00190* (-.57)	0.01639* (1.53)
D69	-0.00858 (-2.59)	-0.00780 (-2.34)	-0.02188 (-2.04)
D70	0.00589* (1.76)	0.00603* (1.79)	-0.02552 (-2.34)
D71	0.00669 (2.01)	0.00667 (1.99)	-0.00782* (-.73)
constant	0.25083	0.14630	0.15449
R ²	0.11235	0.09932	0.09979
S.E.E.	42927	0.43240	0.43231
F	953.73	1510.90	596.60

*Insignificant at 95% Confidence Level

TABLE F.2

Parameter Estimates for Three Specifications
of the Linear Probability Function With
Dichotomous Dependent Variable, BINMSP: 150,729
Positive Income Families, Pooled 1966-71 Sample Data

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SIZE (S)		0.01756 (34.43)	0.01642 (21.61)
D68.S			0.00054* (.43)
D69.S			0.00047* (.37)
D70.S			0.00344 (2.69)
D71.S			0.00291 (2.26)
C1	0.04256 (16.50)		
C2	0.05660 (20.73)		
C3	0.06940 (21.76)		
C4	0.07699 (20.10)		
C5	0.09710 (24.46)		
AGE (A)		0.00071 (17.75)	0.00064 (8.00)
D68.A			-0.00015* (-1.15)
D69.A			0.00025* (1.92)
D70.A			0.00030 (2.31)
D71.A			0.00002* (.15)
A1	0.02120 (7.54)		
A2	0.01502 (5.01)		
A3	0.01027 (3.65)		
A4	0.02758 (9.89)		
A5	0.04650 (14.53)		
A6	0.06571 (17.43)		

TABLE F.2 - Continued

Independent Variables	Probability of Hospitalization		
	Model 1	Model 2	Model 3
SINGLE MALE	-0.03955 (-17.20)	-0.03131 (-13.10)	-0.03137 (-13.13)
SINGLE FEMALE	-0.04694 (-18.70)	-0.03774 (-14.57)	-0.03775 (-14.58)
URBAN	0.03861 (24.28)	0.04023 (25.30)	0.04012 (25.23)
WELFARE (W)	0.02431 (5.90)	0.02567 (6.23)	0.02745 (4.78)
INCOME (Y)	0.00087 (5.12)	0.00084 (4.94)	0.00058* (1.93)
W.Y			0.00006 (.01)
D68.Y			0.00012* (.24)
D69.Y			0.00106 (2.16)
D70.Y			0.00061* (1.27)
D71.Y			-0.00013* (-.28)
D68.W.Y			0.00136* (.48)
D69.W.Y			-0.00320* (-1.33)
D70.W.Y			-0.00240* (-.83)
D71.W.Y			0.00088* (.35)
D68	0.00463	0.00425*	0.00827* (1.13)
D69	0.00732	0.00744	-0.01056* (-1.45)
D70	0.01246	0.01237	-0.01682* (-1.86)
D71	0.01124	0.01101	0.00307* (.42)
constant	0.03686	0.00106	0.00884
R ²	0.03042	0.02899	0.02923
S.E.E.	0.29424	0.29445	0.20443
F	236.45	409.07	162.04

*Insignificant at 95% Confidence Level

TABLE F.3

Parameter Estimates for Three Specifications
of the Linear Utilization Function with Alternate
Continuous Dependent Variables DAYDTX and ADMDTX: 44,328
Positive Income Families, Pooled 1966-71 Sample Data

Independent Variables	DAYDTX per Year Given Hospitalization			ADMDTX per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		0.05634* (.83)	0.09563* (.91)		0.10008 (29.52)	0.11301 (21.44)
D68.S			-0.16549* (-.92)			-0.04066 (-4.52)
D69.S			-0.13358* (-.72)			-0.01131* (-1.21)
D70.S			-0.15244* (-.82)			-0.01136* (-1.22)
D71.S			0.21664* (1.17)			-0.01479* (-1.59)
C1	0.29511* (.78)			0.11432 (6.03)		
C2	-0.64911* (-1.63)			0.20696 (10.32)		
C3	-0.16117* (-.36)			0.28435 (12.69)		
C4	0.85074* (1.69)			0.40664 (16.09)		
C5	2.02421 (4.04)			0.69447 (27.56)		
AGE (A)		0.31057 (42.90)	0.30218 (23.61)		0.00251 (6.97)	0.00161 (2.52)
D68.A			-0.00255* (-.12)			-0.00139* (-1.28)
D69.A			0.01665* (.77)			0.00170* (1.56)
D70.A			0.03130* (1.45)			0.00259 (2.38)
D71.A			0.00111* (.05)			0.00255 (2.38)
A1	1.81703 (3.73)			0.08584 (3.50)		
A2	3.09127 (6.01)			-0.05135* (-1.99)		
A3	4.71822 (9.48)			-0.06890 (-2.75)		
A4	8.70172 (17.27)			0.01640* (.65)		
A5	13.49980 (24.16)			0.15384 (5.48)		
A6	20.81712 (34.83)			0.30785 (10.25)		

TABLE F.3 - Continued

Independent Variables	DAYDTX per Year Given Hospitalization			ADMDTX per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	3.95736 (9.41)	4.81396 (11.28)	4.81887 (11.28)	-0.02109* (-1.00)	0.09985 (4.66)	0.09791 (4.56)
SINGLE FEMALE	-1.64204 (-3.48)	-0.81040* (-1.70)	-0.82562* (-1.73)	-0.23629 (-9.98)	-0.11737 (-4.90)	-0.12249 (-5.11)
URBAN	0.84765 (3.44)	1.14628 (4.66)	1.15097 (4.67)	-0.24748 (-19.99)	-0.23258 (-18.82)	-0.23445 (-18.95)
WELFARE (W)	6.97808 (12.51)	7.10740 (12.73)	7.15461 (8.98)	0.61524 (21.94)	0.61446 (21.90)	0.69468 (17.36)
INCOME (Y)	-0.14075 (-5.73)	-0.18856 (-7.76)	-0.17098 (-3.97)	-0.00753 (-6.12)	-0.00979 (-8.02)	-0.01285 (-5.95)
W.Y			0.20910* (.87)			-0.01211* (-1.00)
D68.Y			0.02601* (.36)			0.00315* (.86)
D69.Y			0.01612* (.23)			0.00867 (2.45)
D70.Y			-0.05250* (-.75)			0.00601* (1.70)
D71.Y			-0.07576* (-1.12)			0.00356* (1.04)
D68.W.Y			-0.43817* (-1.25)			-0.03819 (-2.16)
D69.W.Y			-0.25795* (-.76)			0.00387* (.23)
D70.W.Y			-0.31367* (-.88)			-0.01861* (-1.04)
D71.W.Y			-0.24076* (-.81)			0.00139* (.09)
D68	-0.50263* (-1.46)	-0.45870* (-1.33)	0.16841* (.12)	-0.03009* (-1.74)	-0.02813* (-1.62)	0.17159 (2.49)
D69	-0.61199* (-1.77)	-0.57451* (-1.66)	-0.90302* (-.64)	-0.01903* (-1.09)	-0.01523* (-.87)	-0.10057* (-1.42)
D70	-0.61819* (-1.79)	-0.61564* (-1.78)	-1.15061* (-.70)	-0.01574* (-.91)	-0.01302* (-.75)	-0.12148* (-1.69)
D71	-0.97065 (-2.82)	-0.93285 (-2.70)	-1.17750* (-.71)	-0.01808* (-1.04)	-0.01478* (-.85)	-0.09831* (-1.40)
constant	8.18846	-0.29508	-0.20472	1.55056	1.28304	1.29184
R ²	0.05837	0.05503	0.05532	0.05209	0.04844	0.045965
S.E.E.	24.14669	24.18693	24.18791	1.21329	1.21550	1.21496
F	137.32	234.63	92.65	121.74	205.10	82.65

*Insignificant at 95% Confidence Level

TABLE F.4

Parameter Estimates for Three Specifications
of the Linear Utilization Function With Alternate
Continuous Dependent Variables DAYMSP and ADMMSP: 14,935
Positive Income Families, Pooled 1966-71 Sample Data

Independent Variables	DAYMSP per Year Given Hospitalization			ADMMSP per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SIZE (S)		-0.34070 (-4.88)	-0.37419 (-3.29)		0.01787 (5.01)	0.01781 (3.05)
D68.S			0.24783* (1.31)			-0.00482* (-.50)
D69.S			-0.00946* (-.05)			-0.00796* (-.81)
D70.S			-0.04255* (-.22)			0.00941* (.97)
D71.S			0.01566* (.08)			0.00473* (.49)
C1	-0.60884* (-1.60)			0.01037* (.53)		
C2	-1.48412 (-3.77)			0.03246* (1.61)		
C3	-1.30922 (-2.95)			0.06087 (2.69)		
C4	-1.29375 (-2.54)			0.09119 (3.50)		
C5	-1.10948 (-1.95)			0.10295 (3.92)		
AGE (A)		0.22191 (29.01)	0.23091 (16.90)		-0.00014* (-.36)	-0.00006* (-.09)
D68.A			0.01763* (.77)			-0.00088* (-.75)
D69.A			0.00214* (.10)			0.00001* (.01)
D70.A			-0.01618* (-.72)			0.00066* (.57)
D71.A			-0.05561 (-2.49)			-0.00035* (-.31)
A1	1.37513 (2.62)			0.07001 (2.61)		
A2	2.66886 (4.92)			0.01437* (.52)		
A3	4.57768 (8.70)			0.01466* (.54)		
A4	7.69926 (14.51)			0.01456* (.54)		
A5	10.26391 (17.58)			0.06208 (2.08)		
A6	12.33652 (19.12)			0.03546* (1.08)		

TABLE F.4 - Continued

Independent Variables	DAYMSP per Year Given Hospitalization			ADMMSMSP per Year Given Hospitalization		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
SINGLE MALE	3.27176 (7.49)	3.25877 (7.39)	3.21750 (7.29)	-0.00592* (-.27)	0.00785* (.35)	0.00748* (.33)
SINGLE FEMALE	-0.57536* (-1.20)	-0.63235* (-1.30)	-0.66406* (-1.37)	-0.10288 (-4.20)	-0.09384 (-3.79)	-0.09387 (-3.78)
URBAN	-1.48031 (-6.09)	-1.41136 (-5.84)	-1.40038 (-5.79)	0.08555 (6.88)	0.08873 (7.18)	0.08933 (7.22)
WELFARE (W)	5.32957 (9.04)	5.45723 (9.26)	4.84259 (1.21)	0.19797 (6.57)	0.19516 (6.48)	0.17225 (4.08)
INCOME (Y)	-0.05199 (-2.49)	-0.06970 (-3.39)	-0.07978 (-2.11)	-0.00060* (-.56)	-0.00081* (-.77)	-0.00129* (-.67)
W.Y			0.61718 (2.54)			0.01325* (1.06)
D68.Y			0.02764* (.44)			0.00118* (.37)
D69.Y			0.02570* (.45)			0.00306* (1.04)
D70.Y			-0.03035* (-.49)			-0.00042* (-.13)
D71.Y			-0.00312* (-.05)			-0.00216* (-.73)
D68.W.Y			-1.06966 (-3.03)			-0.00862* (-.48)
D69.W.Y			-0.78630 (-2.13)			-0.01748* (-.92)
D70.W.Y			.06789* (.21)			0.01024* (.60)
D71.W.Y			0.65997 (2.23)			-0.01808* (-1.19)
D68	-0.38973* (-1.11)	-0.35805* (-1.02)	-2.07626* (-1.42)	-0.02192* (-1.22)	-0.02208* (-1.23)	0.02822* (.38)
D69	-0.90166 (-2.60)	-0.87899 (-2.54)	-1.02146* (-.70)	-0.01670* (-.94)	-0.01640* (-.93)	-0.00713* (-.10)
D70	-0.77596 (-2.25)	-0.75534 (-2.20)	0.32548* (.22)	0.01167* (.66)	0.01200* (.68)	-0.05068* (-1.67)
D71	-1.73883 (-5.04)	-1.72043 (-4.99)	0.92217* (.64)	-0.00894* (-.51)	-0.00959* (-.54)	0.01032* (.14)
constant	9.80547	4.70259	4.43069	1.13035	1.13338	1.13200
R ²	0.09039	0.09003	0.09187	0.01175	0.01056	0.01130
S.E.E.	14.06846	14.06699	14.06071	0.71922	0.71944	0.71958
F	74.10	134.22	53.86	8.86	14.48	6.08

*Insignificant at 95% Confidence Level

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