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# Demand for Reimbursement Insurance

The theory of demand for insurance has been studied by Arrow (1963, 1973a,b), Mossin (1968), Smith (1968), Pratt (1964), Ehrlich and Becker (1973; hereafter EB), and others; and demand estimates have been published by Fuchs and Kramer (1972), Feldstein (1973), and Phelps (1973). These works have generally specified insurance contracts in which (1) the losses are not affected by the amount of insurance (EB is an exception), or (2) the amount of insurance in each period can be chosen directly. I will investigate here demand for insurance under two more restrictive conditions that normally apply to health insurance: (1) the insurance coverage rate must be equal in all insured states of the world (i.e., constant coinsurance in all insured states) and (2) states of the world are aggregated in some fashion before a loading fee is computed.

I assume that the consumer has a utility function in a composite good ($x$) and health ($H$), where health is produced by medical care ($h$) and own-time in fixed production coefficients, and where the stock of health $H$ is subject to random losses ("illnesses") having a known distribution $f(l)$, where $l$ is the illness amount. Hence, the final level of health is $H_0 - l + g(h)$, where $H_0$ is the endowed level

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of health. To capture the essence of the model I assume that the consumer may choose only some coinsurance rate $C$ (the fraction of medical care bills paid by the consumer) and a maximum payment amount $h^*$ (in units of care). The insurance is a subsidy for purchasing medical care, so calculating the optimal values of $C$ and $h^*$ must take into account that more $h$ is purchased for lower values of $C$ (Phelps, 1973; Phelps and Newhouse, 1974).

The consumer pays for his insurance with a premium, $R$, that reflects both the expected expenses for medical care and administrative costs of the insurance process. Several specifications of the insurance premium will be investigated to examine differential effects of different pricing systems on demand for insurance.

In addition, effects of income, and of income tax subsidy of insurance, will be investigated. It can be shown that estimates of the income elasticity of demand for insurance include (1) a "pure" income effect and (2) a substitution effect that depends on the progressiveness of income taxes. Estimated income elasticities of demand for insurance can be separated into these components.

Finally, I will investigate effects of changes in the price level of medical care. It has been asserted that rising prices for medical care induce additional demand for insurance (Feldstein, 1971). This model shows that that conclusion does not hold except under specific circumstances.

The model to be used assumes that consumers maximize expected utility over all possible states of illness. The distribution of illnesses $f(l)$ is assumed to be continuous and smooth in the range $0 \leqslant l < \infty$. Thus, expected utility can be written as

$$Z = s_0 U(x_0, H_0) + \int_0^{l^*} U[x, H_0 - l + g(h)] f(l) dl + \int_{l^*}^{\infty} U[x, H_0 - l + g(h)] f(l) dl \tag{1}$$

where $l^*$ is the loss that induces the consumer to purchase $h^*$ units of care,[1] and where

$$s_0 + \int_0^{\infty} f(l) dl = 1$$

For $l \leqslant l^*$, the budget constraint is:

$$I = x + C p_h h + R \tag{2}$$

reflecting the insurance subsidy for medical care.

For the second integral, in which insurance no longer pays, the budget constraint is:

$$I = x + p_h h + R - (1 - C) p_h h^* \tag{3}$$

where $(1-C)p_h h^*$ is the maximum amount paid by the insurance plan. Optimal levels of $C$ and $h^*$ are derived by solving the first derivatives when set to zero. When Z is differentiated with respect to $C$ and $h^*$ and set to zero subject to the appropriate budget constraints (2) and (3), the results are:[2]

$$s_0(-R_c)\cdot\lambda(l_0) + \int_0^{l^*}\lambda(l)\left[-R_c - p_h h(l)\right]f(l)dl + \int_{l^*}^{\infty}\lambda(l)(-R_c - p_h h^*)f(l)dl = 0 \tag{4}$$

and

$$s_0(-R_{h^*})\cdot\lambda(l_0) + \int_0^{l^*}(-R_{h^*})\cdot\lambda(l)f(l)dl + \int_{l^*}^{\infty}\left[-R_{h^*} + (1-C)p_h\right]\lambda(l)f(l)dl = 0 \tag{5}$$

where $R_c$ and $R_{h^*}$ are the partial derivatives of $R$ with respect to $C$ and $h^*$, respectively, and $\lambda(l)$ is the marginal utility of income for the health loss $l$.

Expressions (4) and (5) are subject to an interpretation parallel to that used by EB. In (5), if $h^*$ increases, the net income flow out of all states more favorable than $h^*$ is $-R_{h^*}$, the derivative of the premium as $h^*$ changes. For states less favorable than $l^*$ (i.e., larger losses), the net income transfer rate is $-R_{h^*}$ (the premium still changes) plus $(1-C)p_h$, the last expression reflecting the additional rate of income flow into those states with illnesses greater than $l^*$. The real price of insurance, defined in the EB sense, is $R_{h^*}/[-R_{h^*} + (1-C)p_h]$, the negative of the ratio of income flow rates between states below and above $h^*$. Equation (5) can be solved for this price, showing

$$P_{h^*} = \frac{R_{h^*}}{(1-C)p_h - R_{h^*}} = \frac{\int_{l^*}^{\infty}\lambda(l)f(l)dl}{s_0\lambda(l_0) + \int_0^{l^*}\lambda(l)f(l)dl} = \frac{\bar{U}'}{\underline{U}'} \tag{6}$$

The optimal $h^*$ is chosen so that the ratio of *expected* marginal utility of incomes above and below $h^*$ is equal to the real price of insurance—the ratio of income transfers between less favorable and more favorable states. This is analogous to the two-state problem analyzed by EB, except that here the consumer divides all possible states of nature into two categories on the basis of his choice of $h^*$.

It is possible to show that some insurance is optimal with the loading fee sufficiently small. To prove this point, notice that where

$$R = \int_0^{l^*} (1-C)p_h hf(l)dl + \int_{l^*}^{\infty} (1-C)p_h h^* f(l)dl$$

is an actuarily fair premium (no load), then

$$\frac{\partial R}{\partial h^*} = R_{h^*} = (1-C)p_h \int_{l^*}^{\infty} f(l)dl = (1-C)p_h Q(l^*)$$

Then to show that some $h^*>0$ is optimal, I evaluate (5) at $h^* = 0$. At that point, (5) becomes

(7a)
$$-s_0\lambda(l_0)(1-C)p_h Q(l^*) + \int_0^{\infty} \lambda(l)\ [(1-C)p_h - (1-C)p_h Q(l^*)]\, f(l)dl$$
$$= \left\{-s_0\lambda(l_0)Q(l^*) + (1-s_0)\bar{\lambda}[1-Q(l^*)]\right\}(1-C)p_h$$

where $(1-s_0)\bar{\lambda} = \int_0^{\infty} \lambda(l)f(l)dl$ is expected marginal utility for all states of the world with losses. Notice that for $l = 0$, $Q(l) = 1-s_0$, so (7a) becomes

(7b)
$$[-\lambda(l_0)(s_0)(1-s_0) + \bar{\lambda}(1-s_0)(s_0)]\,(1-C)p_h$$
$$= [\bar{\lambda} - \lambda(l_0)]\,(1-C)p_h s_0(1-s_0) > 0$$

since $\lambda(l_0)$, the marginal utility of income with no loss, is less than $\lambda(l)$ for any larger $l$. (A loss, $l$, can be translated into an income loss, so that $\partial\lambda/\partial l > 0$.) Thus for "small" $\theta$, *some* purchase of insurance is optimal.

It also follows that some loading fee of sufficient size exists to deter purchase. Let the premium function be $R = (1+\theta)E$(benefits), where $\theta$ is the load. Then $R_{h^*} = (1+\theta)(1-C)p_h Q(l^*)$, and $\partial E(U)/\partial h^*$ at $h^* = 0$ is:

(8)
$$-s_0\lambda(l_0)(1+\theta)(1-C)p_h Q(l^*) + \int_0^{\infty} \lambda(l)[(1-C)p_h$$
$$-(1+\theta)(1-C)p_h Q(l^*)f(l)dl](1-C)p_h$$
$$\left\{-s_0\lambda(l_0)(1+\theta)(1-s_0) + (1-s_0)\bar{\lambda}[1-(1+\theta)(1-s_0)]\right\}$$
$$= (1-C)p_h(1-s_0)\left\{s_0(1+\theta)\ [\bar{\lambda}-\lambda(l_0)]\ - Q\bar{\lambda}\right\}$$

This expression is negative (i.e., $E(U)$ falls as $h^*$ rises from $h^* = 0$) if

(9)
$$\theta > \frac{s_0[(\bar{\lambda} - \lambda(l_0)]}{s_0\lambda(l_0) + (1-s_0)\bar{\lambda}} = \frac{s_0\Delta\lambda}{E(\lambda)}$$

That is, if the loading fee at $h^* = 0$ exceeds the difference between marginal utilities of income in the states with and without loss (relative to their expectation), times the probability of no loss, then no insurance will be purchased. Notice the similarity between the right-hand-side of (9) and the usual risk-aversion measure.

The problem with respect to choice of $C$ is more complex and is in fact the major difference between choice of single-coverage reimbursement insurance and other forms of insurance. The consumer must select one level of coinsurance $C$ that maximizes expected utility over all possible insured states.

As seen from Equation (10) below, $R_c$ is negative. The income flow rate into any state with $l \leq l^*$ is $(p_h h + R_c)$. For small losses ("favorable states") the flow rate is negative. There is some loss $l$ between 0 and $l^*$ such that the net income flow is zero.[3] For larger losses, the net income flow is positive, including those for $l > l^*$, in which case the income flow rate is $(p_h h^* + R_c)$. The "real price of insurance for state $l$" is $(-R_c)/[p_h h(l) + R_c)$, the ratio of income transfers from the least serious state ($h = 0$) to the more serious states. If this "price" is negative, the income flow is away from that state. Notice that the consumer does not choose some income transfer for each state, but rather only one value for $C$. Hence, he is really not directly allocating incomes according to the prices in each state; that allocation follows from the choice of $C$ and from his choice of $h$ given $l$. Only for the trivial case of $h = \bar{h}$ for all $l \leq l^*$ is a direct solution of (4) possible. Thus, it is difficult to speak of "the price of $C$" in the usual sense. I shall show below that where $P_l = -R_c/[p_h h(l) + R_c]$, the price of $C$ for loss $l$, the comparative statics of demand for $C$ will hinge crucially on how $p_l$ changes with $l$. Since changes in income or medical prices change both $R$ (and hence $R_c$) and $h(l)$, an income-generated own-price effect on demand for care may translate into a modified price effect of demand for insurance. I will show that if $P_l$ increases with $l$ (for whatever reason), then demand for $C$ falls commensurately. It is by this method that comparative statics of demand for $C$ may be analyzed. It can be shown also that some insurance is optimal (i.e., $C < 1$) for a sufficiently small loading fee and small $s_0$. For this proof, I make the simplifying assumption that $h^* = \infty$, because this simplifies the calculation of the premium change function $R_c$. In general, the premium function is

$$R = (1 + \theta)\left[ \int_0^{l^*} (1-C)p_h h(l) f(l)dl + \int_{l^*}^{\infty} (1-C)p_h h^* f(l)dl \right]$$

so

$$-R_c = (1 + \theta) \left\{ \int_0^{l^*} \frac{p_h h(l)}{C} \left[C - (1-C)\eta_{hc}\right] f(l)dl + p_h h^* Q(l^*) \right\} \tag{10}$$

Clearly, at $\theta = 0$ and at $C = 1$, this becomes $-R_c = p_h \bar{h} + p_h h^* Q(l^*)$. For $h^* = \infty$, $-R_c = p_h \bar{h}$, the average expense in states with positive illness, where $h$ is a function of $p_h$, $l$, and $C$, but is evaluated at $C = 1$. Expression (4), showing how expected utility changes with $C$, can be evaluated at $C = 1$ to determine if $E(U)$ rises or falls with $C$—if it falls ($dZ/dC < 0$), then some insurance purchase is optimal. Expression (4), using the above substitution for $-R_c$, becomes

$$\text{(11)} \quad s_0 \cdot \lambda(l_0) \cdot (p_h\bar{h}) - \int_0^\infty \lambda(l)\left[-p_h\bar{h} + p_h h(l)\right]f(l)dl$$
$$= s_0 \cdot \lambda(l_0) \cdot (p_h\bar{h}) - p_h \int_0^\infty \lambda(l)\left[h(l) - \bar{h}\right]f(l)dl$$

The expression in the integral is positive; if $\lambda(l)$ were constant over all losses, then the integral would merely be the integral of deviations of $h(l)$ about its mean; but each value is weighted by $\lambda(l)$, which increases as $l$ increases. Thus the larger terms in the integral (recall that $h(l)$ increases with $l$) carry a larger weight, and the integral itself is positive. Therefore, (11) is negative for a sufficiently small value of $s_0$, or if the marginal utility of income rises sufficiently between the state with no loss ($l_0$) and all other states, which means that $E(U)$ will rise if $C$ is reduced from 1 to some smaller value—some insurance is purchased.

When selecting $C$ less than 1, the model shows that the larger the coinsurance elasticity of demand for $h$, the larger will be the optimal coinsurance level. Optimal $C$ is chosen so that for any smaller $C$, expected utility ($Z$) falls; in other words, $dZ/dC$ becomes positive, which can happen only if $-R_c$ grows "large" relative to the average expense $p_h\bar{h}$. From (10), $-R_c$ obviously becomes large as either $\theta$ increases or as $\eta_{hc}$ increases in absolute value. If $\eta_{hc}$ is sufficiently large, the optimal $C$ could be near 1.[4]

One prediction that might be made informally from this model is that the equilibrium coinsurance rate (*ceteris paribus*) will be larger (less insurance), the larger the own-price elasticity of demand for that service.

## 1. COMPARATIVE STATICS OF DEMAND FOR MAXIMUM COVERAGE

First, I will analyze effects of income, medical prices, and various insurance pricing strategies on the demand for $h^*$.

### Income

The comparative statics are developed by fully differentiating the first-order conditions and solving the resultant equation for the appropriate derivative. Let $\partial Z$ represent Equation (5) and $X$ represent any independent variable of interest. Then the full derivative is

$$dz = \frac{-\partial(\partial Z)}{\partial X}\,dX + \frac{\partial(\partial Z)}{\partial h^*}\,dh^* \tag{12}$$

which must be set at zero in order to maintain the optimality of (5). Hence

$$dh^*/dX = \frac{-\partial(\partial Z)}{\partial X}\Big/\frac{\partial(\partial Z)}{\partial h^*} = \frac{-\partial^2 Z}{\partial h^*\,\partial X}\Big/\frac{\partial^2 Z}{\partial h^{*2}} \tag{13}$$

The denominator of this expression is simply the second-order condition for optimality with respect to $h^*$—it must be negative if expected utility is maximized.[5] Thus the sign of $dC/dX$ is identical to the sign of the cross-partial derivative $\partial^2 Z/\partial C\,\partial X$; i.e., the derivative of (5) with respect to $X$. A similar result holds, of course, for the comparative statics with respect to $C$, where the derivative of (4) is taken. That is,

$$\frac{dC}{dX} = \frac{-\partial^2 Z}{\partial C^*\partial X}\Big/\frac{\partial^2 Z}{\partial C^2} \tag{14}$$

The basic income effect on $h^*$ is found by establishing the sign of:[6]

$$\frac{\partial^2 Z}{\partial h^*\partial l} \propto \int_0^{l^*} (-R_{h^*})\,\frac{\partial\lambda(l)}{\partial l} f(l)dl + \int_{l^*}^{\infty} [-R_{h^*} + (1-C)p_h]\,\frac{\partial\lambda(l)}{\partial l} f(l)dl \tag{15}$$

$$= \int_0^{l^*} (-R_{h^*})(-\lambda)(l)r(\mathrm{I},l) + \int_{l^*}^{\infty} [-R_{h^*} + (1-C)p_h](-\lambda)(l)r(\mathrm{I},l)f(l)dl$$

where $r(\mathrm{I},l)$ is the Arrow-Pratt risk-aversion measure, defined as $r(\mathrm{I},l) = -\partial^2U/\partial I^2/\partial U/\partial I = -(\partial\lambda/\partial I)/\lambda$. Equation (15) is obviously just the negative of the first-order condition (5) weighted by the risk-aversion measure. If the person exhibits constant risk aversion over all illness levels (equivalent to constant risk aversion over all incomes), then this "income effect" on demand for maximum coverage is zero. If risk aversion increases with illnesses (decreases with income), then $r(\mathrm{I},l)$ acts as a weighting scheme making (15) negative, so that demand for $h^*$ falls as income rises. If risk aversion is increasing in income (decreasing as illnesses increase), then (15) becomes positive and demand for $h^*$ increases as income increases.

This "basic income effect" appears throughout the comparative statics of demand for reimbursement insurance. Invariably, expressions appear where an income effect is present, based on how the marginal utility of income ($\lambda(l)$) changes as some variable, $X$, changes. The income effect in these derivatives is always determined by how $\partial\lambda/\partial X$ changes over values of $l$. Hence, it is sufficient to determine the behavior of $\partial\lambda/\partial X$ over $l$ in order to determine the "income component" of effects of any variable, $X$, on demand.

## Medical Prices

In addition to the pure income effects, this model can show the effects of a change in $p_h$ on demand for maximum coverage. The income effect of a change in $p_h$ is determined by the behavior of $\partial\lambda/\partial p_h$ as $l$ changes. That derivative is

$$\frac{\partial\lambda(l)}{\partial p_h} = \begin{cases} \left(-Ch - \dfrac{\partial R}{\partial p_h}\right) \cdot r(I) \cdot \lambda + C\lambda \dfrac{\partial h}{\partial I} \text{ for } h \leqslant h^* \\ \left[-(h - h^*) + Ch^* - \dfrac{\partial R}{\partial p_h}\right] r(I) \cdot \lambda + \lambda \dfrac{\partial h}{\partial I} \text{ for } h > h^* \end{cases} \tag{16}$$

which (in elasticity form) becomes the "weights" in the integral similar to how $r(I)$ acts as weights in (15). These weights are

$$\frac{p_h}{\lambda(l)} \cdot \frac{\partial\lambda(l)}{\partial p_h} = \begin{cases} -(\omega_h + \omega_R\eta_{Rp_h}) \cdot r^*(I) + \omega_h\eta_{hI} \text{ for } h \leqslant h^* \\ -(\omega_h + \omega_R\,\eta_{Rp_h}) \cdot r^*(I) + [\omega_h\eta_{hI} + (1-C)p_h h^*]\eta_{hI} \\ \text{for } h > h^* \end{cases} \tag{17}$$

where $\omega_h$ is the *out-of-pocket* budget share—i.e., $Cp_hh/I$ for $h \leqslant h^*$, and $[p_h(h - h^*) + Cp_hh^*]/I$ for $h > h^*$, $\eta_{hI}$ is the income elasticity of demand for $h$, $\eta_{Rp_h}$ is the total elasticity of the premium $R$ with respect to $p_h$, and $\omega_R$ is the budget share of $R$ itself.

Although this income effect appears complex, its sign is basically determined by whether $r^*(I)$—the income elasticity of the marginal utility of income—is larger or smaller than the income elasticity of demand for medical care.[7] The factors are offsetting—when medical prices rise, the income effect $\eta_{hI}$ induces demand for higher coverage, but general purchasing power falls commensurately, tending to reduce demand for $h^*$. Empirical estimates of $r^*(I)$ are in the range of 1.5 to 2.5, whereas estimates of $\eta_{hI}$ are in the range of 0.1 to 0.5, so the general income effect is likely to dominate the income effect of demand for $h$.

Several things are worth remembering about this phenomenon. First, if $C$ is small (nearly zero coinsurance), then the budget share terms $\omega_h$ tend to vanish in (17), but $\omega_R$ will be larger. Thus at one extreme (low $C$, low $\eta_{hI}$), only the terms involving $r^*(I)\omega_R\eta_{R\nu h}$ would remain, which are negative. In this case, the income effect of a change in $p_h$ would be negative—higher $p_h$ would mean lower $h^*$. At another extreme, if $\eta_{R\nu h}$ is near zero—as would occur if demand for medical care were nearly of unitary elasticity—and if $C$ were near zero, then only the final term $(1-C)p_h h^* \eta_{hI}$ would remain in the last integral, which is positive. In this case, demand for $h^*$ would rise as $p_h$ increased, via this income effect. Thus it appears that there are interactions between demand for $C$ and $h^*$, and even interactions between demand for $h^*$ and demand for $h$. The resultant derivatives are, however, ambiguous.

There is no substitution effect on demand for $h^*$ as $p_h$ changes, because the price of medical care is simply a numeraire for that problem. This is most easily shown by noting that the "real price of insurance," defined as $P_{h^*} = R_{h^*}/[-R_{h^*} + (1-C)p_h]$, does not change as $p_h$ changes. This is true because $R_{h^*} = (1+\theta)(1-C)p_h h^* f(l^*)$ and the partial derivative of that with respect to $p_h$ is simply equivalent to dividing through by $p_h$. Since $p_h$ is a common factor in all terms in $P_{h^*}$, there is no change in $P_{h^*}$ as $p_h$ changes.

If the maximum payment is described in physical units (such as $h^*$), then, except for the net income effects, there is no effect from a change in $p_h$. If the limit is in terms of dollars, then the effect is positive, and the measured elasticity should center around unity, not zero. Put differently, people shouldn't change their demand for maximum coverage as the price of care changes, except for effects introduced by income considerations.

## Loading Fees

Finally, I would like to investigate how demand for $h^*$ changes as the loading fee on insurance changes. To do this, I shall introduce several forms of an insurance premium, specifying different types of loading. One common specification is that the loading fee, $\theta_1$, is applied uniformly to all expected benefits, so that

$$R_1 = (1+\theta_1)\left[\int_0^{l^*} (1-C)p_h h(l) f(l)\,dl + (1-C)p_h h^* \int_{l^*}^{\infty} f(l)\,dl\right] \tag{18}$$

so

$$R_{h^*} = (1+\theta)(1-C)p_h Q(l^*)$$

and

$$\frac{\partial^2 R}{\partial h^* \partial \theta} = (1-C)p_h Q(l^*) > 0$$

In this case, it can be shown that demand for maximum coverage behaves much as a usual good—there is a negative substitution effect (demand falls as $\theta_1$ rises), and an income effect that depends on whether demand for $h^*$ rises or falls with income (i.e., whether $h^*$ is a superior or inferior good). If $\partial h^*/\partial I$ is positive, then the income effect (as in the usual case) is negative, and demand for $h^*$ falls as $\theta_1$ rises.

The substitution component of this derivative is simply

$$(19)\qquad \int_0^{l^*} [\lambda(l)] \left(\frac{\partial^2 R}{\partial h^* \partial \theta}\right) f(l)dl + \int_{l^*}^{\infty} [\lambda(l)] \left(\frac{\partial^2 R}{\partial h^* \partial \theta}\right) f(l)dl$$

i.e., the expected value of marginal utility of income times $-\partial^2 R/\partial h^* \partial \theta$, which is positive. Hence, the substitution effect is negative. I compute the income effect by making $\partial \lambda/\partial \theta = (\partial \lambda/\partial I) \cdot R\theta$, where $R_\theta$ is the first partial derivative of $R$ with respect to $\theta$ (holding $C$ and $h^*$ constant) and is positive. Hence, the income effect of a change in $\theta$ is proportional to the pure income effect and therefore depends on whether risk aversion is increasing or decreasing. If $r(I)$ is increasing in income, so that $\partial h^*/\partial I$ is positive, then the entire derivative $\partial h^*/\partial \theta$ is negative, just as in a "standard" consumer demand problem.

Consider now a second possible formulation of the premium function

$$(20)\qquad R_2 = \theta_2 + \int_0^{l^*} (1-C)p_a h(l)f(l)dl + \int_{l^*}^{\infty} (1-C)p_a h^* f(l)dl$$

In this formulation, there is a fixed loading fee, (say, sales costs) but no administrative costs otherwise. Then $\partial R/\partial \theta_2 = 1$, and $\partial^2 R/\partial h^* \partial \theta_2 = 0$. In this case, an increase in $\theta_2$ produces an income effect, but no substitution effect, and will be proportional to the income effect $\partial h^*/\partial I$.

A third possible premium function combines the first two—a fixed loading charge, $\theta_3$, and a declining proportional addition, $\theta_4$—so that

$$(21)\qquad R_3 = \theta_3 - \theta_4 \left[ \int_0^{l^*} (1-C)p_a h(l)f(l)dl + \int_{l^*}^{\infty} (1-C)p_a h^* f(l)dl \right]$$

Under this pricing scheme, there would be a positive fixed charge and a declining marginal cost of coverage, as measured by expected benefits. There is a declining marginal cost of $h^*$, and such a pricing scheme can occur only if there are scale economies in coverage. Clearly, the greater the scale economies (i.e., the higher is $\theta_4$ in absolute value), the more $h^*$ would be demanded.

Now consider a premium with rising marginal costs of coverage:

(21a) $$R_4 = \int_0^{l*} (1-C)p_h h(l)[\theta_5 h(l)] f(l)dl + \int_{l*}^{\infty} (1-C)p_h h^*(\theta_5 h^*) f(l)dl$$

so

(21b) $$R_{h^*} = (1-C)p_h\,(2\theta_5 h^*)\,Q(l^*) > 0$$

and

(21c) $$\frac{\partial^2 R}{\partial h^* \partial \theta_5} = 2(1-C)p_h Q(l^*) > 0$$

Notice that the EB "real price of insurance" can rise with $h^*$, unlike the other premium functions considered. The real price—the rates of income transfer rates between insured and uninsured states—is defined by

(22a) $$P_{h^*} = \frac{R_{h^*}}{(1-C)p_h - R_{h^*}}$$

For this premium function (21a),

(22b) $$P_{h^*} = \frac{(1-C)P_h(2\theta_5 h^*)Q(l^*)}{(1-C)P_h - (1-C)P_h(2\theta_5 h^*)Q(l^*)} = \frac{Z}{1-Z}$$

where $Z = 2\theta_5 h^* Q(l^*)$.
Thus, price changes with $h^*$ according to

(22c) $$\frac{\partial P_{h^*}}{\partial h^*} = \frac{\partial Z/\partial h^*}{(1-Z)^2} = 2\theta_5\left[h^* \frac{\partial Q(l^*)}{\partial h^*} + Q(l^*)\right]$$

$$= 2\theta_5 Q(l^*)\,(1+E)$$

where $E$ is the elasticity of $Q(l^*)$ with respect to $h^*$. Thus, as long as $Q(l^*)$ is inelastic with respect to $h^*$, the real price of insurance (in the EB sense) rises with $h^*$. Hence, $P_{h^*}$ changes with the distribution of $l$, and can fall and then rise as $l^*$ increases. Such a pricing function is consistent with observed finite maximum amounts of insurance coverage in health insurance policies. Notice that $E = -f(l) \cdot l/Q(l)$, since $\partial Q(l)/\partial l = -f(l)$. The ratio of $f(l)/Q(l)$ is known in statistics as the hazard function. Thus, $P_{h^*}$ increases with $h^*$ for illness distributions in which the hazard function decreases faster than $l$ increases, since in that case $(1+E) > 0$.

It is quite clear, since $R_{h^*}$ and $\partial^2 R/\partial h^* \partial \theta$ are of appropriate sign, that the price effect computed for the premium function $R_1$ also holds for this premium function; namely, that increases in $\theta_5$ will reduce demand for $h^*$ as long as the income component of the price effect is normal.

## 2. COMPARATIVE STATICS OF DEMAND FOR COINSURANCE

Selecting the coinsurance level, $C$, is a more complicated analysis (at least for the analyst!), the complication arising primarily because coinsurance subsidizes purchase of a market good (medical care) rather than simply providing an income transfer. The second major complication arises because one and only one level of coinsurance $C$ is selected, which must apply to all states of the world less than $l^*$. Because of the interdependence of the different states, comparative statics effects are complex. First, I identify the basic income effect, found here by establishing the sign of $\partial^2 Z/\partial C\, \partial I$.

$$\frac{\partial^2 Z}{\partial C\, \partial I} = \frac{\partial}{\partial I}\left[\int_0^{l^*} \lambda(l)\ [-R_c - p_h h(l)]\ f(l)dl + \int_{l^*}^{\infty} \lambda(l)\ (-R_c - p_h h^*) f(l)dl\right] \tag{23}$$

The "income-component" is

$$\int_0^{l^*} [-R_c - p_h h(l)]\ \frac{\partial \lambda(l)}{\partial I} f(l)dl + \int_{l^*}^{\infty} (-R_c - p_h h^*)\ \frac{\partial \lambda(l)}{\partial I}\ f(l)dl \tag{24}$$

$$= -\left[\int_0^{l^*} (-R_c - p_h h) r(I)\lambda(l) f(l)dl + \int_{l^*}^{\infty} (-R_c - p_h h^*) r(I)\lambda(l) f(l)dl\right]$$

which is the negative of the first-order condition weighted by the Arrow-Pratt risk-aversion measure. The expression $(-R_c - p_h h)$ becomes more negative as $l$ increases, so increasing risk aversion in income (decreasing risk aversion in $l$) lends less weight to the more negative values, making the expression in brackets positive. Hence, $\partial C/\partial I$ becomes more negative through this component—demand for insurance rises (lower $C$) if risk aversion in income increases. The substitution component produced by a change in income is shown by

$$\int_0^{l^*} \lambda(l) \left[\frac{\partial(-R_c)}{\partial I} - p_h\, \frac{\partial h}{\partial I}\right] f(l)dl + \int_{l^*}^{\infty} \frac{\partial(-R_c)}{\partial I}\ f(l)dl \tag{25}$$

This is most easily understood by treating the case wherein $h^* = \infty$, so that (25) is simply

$$\int_0^{\infty} \lambda(l) \left[\frac{\partial(-R_c)}{\partial I} - p_h\, \frac{\partial h}{\partial I}\right] f(l)dl$$

Notice now that the way the premium derivative $R_c$ changes with income is simply the average over all insured states of $(1 + \theta)$ times the average income effect—i.e.,

$$\frac{\partial(-R_c)}{\partial I} = (1+\theta) \int_0^\infty \left[ p_h \frac{\partial h}{\partial I} - (1-C)p_h \frac{\partial^2 h}{\partial C\, \partial I} f(l)dl = (1+\theta)p_h \frac{\partial h}{\partial I} \right]$$

(assuming no income interaction with the price effect, $\partial h/\partial C$). Thus, the substitution effect produced by a change in income is

$$\int_0^\infty \lambda(l) \left[ (1+\theta)p_h \frac{\partial h}{\partial I} - p_h \frac{\partial h(l)}{\partial I} \right] f(l)dl \tag{26}$$

For this to be negative (i.e., making $dC/dI$ negative, so that more insurance is demanded with income), the income effect, $\partial h/\partial I$, must be increasing with the level of illness; if $\partial h/\partial I$ is constant or decreasing as $l$ increases, it is easy to see that (26) is positive, since $\lambda(l)$ increases with $l$. An important point is that since changes in income change the relative prices of insuring different states of the world, a substitution effect on demand for $C$ is introduced when $I$ changes. Thus one cannot infer anything about risk-aversion characteristics of the utility function from income elasticities of demand for coverage; the income elasticity is a composite effect of a "pure" income effect (dependent on the risk-aversion measure) and a substitution effect (dependent on how $\partial h/\partial I$ changes over losses).

One conclusion that may be drawn from statistical studies of demand relates to pricing policies of the health insurance industry. If rates are set independently of income, then $\partial(-R_c)\partial I = 0$, and the only subsidy effect is obviously to increase demand for coverage as $I$ rises. Thus, unless possible decreasing risk aversion outweighs the substitution effect, "community-rated" insurance should lead to a positive income elasticity of demand for coverage. Since the Blue Cross and Blue Shield plans have been traditionally community-rated rather than experience-rated, enrollment shares of the Blues should be higher in high-income areas than in low-income areas, *ceteris paribus*, since the Blues also offer high-coverage insurance.[8]

## Medical Prices

Changes in the price of medical care can also affect demand for $C$, although in complex ways, and the sign of the effect is indeterminate. As with demand for $h^*$, there is an income effect and a substitution effect. The income effect is determined by how the net change in income varies over different $l$. The income effect of a change in $p_h$ is proportional to

$$(27) \quad \int_0^{l^*} (-R_c - p_h h) \frac{\partial \lambda}{\partial p_h} f(l)dl + \int_{l^*}^{\infty} (-R_c - p_h h^*) \frac{\partial \lambda}{\partial p_h} f(l)dl$$

where $\partial\lambda/\partial p_h$ is given in (16) and restated in (17).

Similar conclusions can be drawn with respect to demand for $C$ as are drawn for $h^*$ when $p_h$ changes. First, suppose that $\partial h/\partial I$ is zero, so that the income effect is entirely a function of $r(I)$, since $\partial\lambda/\partial p_h = - S \cdot r^*(I) - \omega_h \eta_{hI}$. The income effect is then simply the first-order condition weighted by $r^*(I)$ and by the expression $S = (\omega_h + \omega_R \eta_{Rph})$. The out-of-pocket budget share, $S$, must increase with $l$, so there are larger and larger weights applied to the more and more negative values in (27), and $\partial C/\partial p_h$ falls (demand for insurance increases as $p_h$ rises). If risk aversion in income is increasing ($r(I)$ is falling with $l$), then that produces an offsetting effect, but the net must be to increase demand for insurance (reduce $C$) unless the risk-aversion measure falls faster than the out-of-pocket budget share rises. I view this as unlikely.

Finally, if the assumption that $\partial h/\partial I = 0$ is removed, then there is an offsetting income effect—there is lower demand for $h$ at all levels of $l$ because of the income effect, so that portion of the income effect reduces demand for coverage (increases $C$) as $p_h$ rises.

The substitution effect as $p_h$ changes is

$$(28) \quad \int_0^{l^*} \lambda(l)\left[- \frac{\partial^2 R}{\partial C\, \partial p_h} - h(1+\eta_{hh})\right] f(l)dl + \int_{l^*}^{\infty} \lambda(l)\left[- \frac{\partial^2 R}{\partial C\, \partial p_h} - h^*\right] f(l)dl$$

where $\eta_{hh}$ is the own-price elasticity for $h$ at any given $l$ and is a function of $l$. The components in brackets in (28) show how the "real price of coverage" changes as $p_h$ changes, the change being a function of how the net income transfer ($p_h h + R_c$) changes. The premium slope ($-R_c$) changes with $p_h$, as does the expenditure $p_h h$, the latter as a function of the own-price elasticity for $h$. The premium slope $-R_c$ changes according to:[9]

$$(29) \quad \frac{-\partial^2 R}{\partial C\, \partial p_h} = \frac{\partial(-R_c)}{\partial p_h} = (1+\theta)\left\{ \int_0^{l^*} \left[h(1+\eta_{hh}) - (1-C)\frac{\partial h(l)}{\partial C}\right] f(l)dl + h^* Q(l^*)\right\}$$

$$= (1+\theta) h \overline{(1+\eta_{hh})} + (1+\theta)h^* Q(l^*) - (1+\theta)(1-C) \int_0^{l^*} \frac{\partial h}{\partial C} f(l)dl$$

Thus the substitution effect in part is a function of the deviation of $h(1+\eta_{hh})$ around its mean (assuming "small" values of $\theta$), so that if the own-price elasticity of demand for $h$ approaches zero as $l$ increases, (28) tends to become more negative, indicating increased demand for insurance (lower $C$). Thus, part of the effect of $p_h$ on demand for $C$ depends on how the own-price elasticity for $h$ interacts with the size of loss $l$. The total effect of a change in $p_h$ is ambiguous on net, however, because of the remaining term in $\partial^2 R/\partial C\, \partial p_h$, as well as additional ambiguity introduced by larger values of $\theta$ that offset increased demand.

## Loading Fees

Finally, I will analyze the effects of changes in the loading fee on demand for $C$. As before, several alternative structures of the premium function $R$ can be hypothesized and their effects detailed. First, consider the function

$$R_1 = (1+\theta)\left[\int_0^{l^*} (1-C)p_h h f(l)dl + (1-C)p_h Q(l^*)\right]$$

Here, a change in $\theta$ produces an income effect on demand for $C$ proportional to

$$(30)\quad \frac{-\partial R}{\partial \theta}\left[\int_0^{l^*} (-R_c - p_h h)\frac{\partial \lambda(l)}{\partial I} f(l)dl + \int_{l^*}^{\infty} (-R_c - p_h h^*)\frac{\partial \lambda(l)}{\partial I} f(l)dl\right]$$

which has the usual income effect properties. Since $R_\theta$ is of necessity positive for this functional form of $R$, the income effect depends on whether there is increasing or decreasing risk aversion in income, and will have the "usual" income effect of a price change if risk aversion is such that insurance is a normal good.

The substitution effect produced by a change in $\theta$ is of the form

$$(31)\quad \int_0^{l^*} \lambda(l)\left[\left(\frac{-\partial^2 R}{\partial C\,\partial \theta}\right) - p_h \frac{\partial h}{\partial \theta}\right] f(l)dl + \int_{l^*}^{\infty} \lambda(l)\left(\frac{-\partial^2 R}{\partial C\,\partial \theta}\right) f(l)dl$$

where

$$\frac{\partial h}{\partial \theta} = -R_\theta \cdot \frac{\partial h}{\partial l}$$

and

$$\frac{-\partial^{2R}}{\partial C\,\partial\theta} = \frac{\partial}{\partial\theta}(-R_c) = \int_0^{l^*}\left[p_h h - (1-C)p_h\frac{\partial h}{\partial C} + (1+\theta)p_h\frac{\partial h}{\partial\theta}\right]f(l)dl + (1-C)p_h Q(l^*)$$

Here, a very unusual result appears. There is a standard substitution effect that can be shown to be approximately

$$\int_0^{\infty}\lambda(l)\cdot\frac{(-R_c)}{(1+\theta)}f(l)dl = E(\lambda)\cdot\frac{(-R_c)}{(1+\theta)} > 0 \tag{32}$$

which is positive (thus reducing demand for coverage as $\theta$ increases). But there is also an effect of changes in $\theta$ on the relative prices of insuring different states of the world, because $R_\theta$ changes available income, which changes demand for care in each state. This substitution term is of the form

$$R_\theta\cdot\int_0^{l^*}\lambda(l)p_h\left[\frac{\partial h}{\partial I} - (1+\theta)\frac{\overline{\partial h}}{\partial I}\,f(l)dl\right] \tag{33}$$

If $\partial h/\partial I$ increases sufficiently over losses, this term is positive, and demand for $C$ behaves as a standard good as $\theta$ increases. But if $\partial h/\partial I$ falls as $l$ increases, then there is a general income effect on demand for $h(l)$, and the real price of insuring states of the world tends to fall as $\theta$ increases. If this effect were sufficiently strong, demand for coverage might actually rise as $\theta$ increased, although I view this as a remote possibility. Empirically, as I shall show in the next section, demand for any measure of insurance I have used behaves "normally" with respect to changes in the loading fee, indicating that this condition does not occur in observed data with sufficient strength to produce an unusual price effect.

The second premium function I shall consider is

$$R_2 = \theta_2 + \int_0^{l^*}(1-C)p_h h(l)dl + (1-C)p_h h^* Q(l^*)$$

i.e., one with a lump sum loading fee and no marginal cost of different levels of $C$. Here, $R_\theta = 1$, and the usual income effects hold—if $C$ is a "normal" good, demand for coverage falls as $\theta_2$ rises. The only substitution effect is this unusual one just discussed with $R_1$, where

$$\frac{\partial(-R_c)}{\partial\theta} = -p_h\frac{\overline{\partial h}}{\partial I}$$

so this substitution effect is

$$\int_0^{l^*} \lambda(l) p_h \left( \frac{\partial h}{\partial I} - \frac{\overline{\partial h}}{\partial I} \right) f(l) dl \tag{34}$$

which can be negative (more demand for insurance with higher $\theta$) if $\partial h/\partial I$ falls as $l$ increases. Thus this reversed "price effect" is also possible under this premium formulation.

The third possible premium function,

$$R_3 = \theta_3 - \theta_4 \int_0^{l^*} (1-C) p_h f(l) dl$$

is a combination of the first two, except that there is declining marginal cost of coverage. Except for the "unusual" price effect here, changes in $\theta_3$ are equivalent to those in the function $R_2$, and increases in $\theta_4$ (the marginal cost reduction factor) will tend to increase demand for coverage (lower $C$).

To summarize these results, recall the specific features of the insurance policy under discussion:

1. There is restricted choice of insurance parameters—only $C$ and $h^*$ may be chosen, rather than an actual net income transfer for each state of nature.
2. States of nature are aggregated when the premium function is written, so that effects of changes in $I$, $p_h$, and $\theta$ are linked across states of the world.

The basic findings are these:

1. The effects of income on demand for insurance depend not only on how risk aversion changes over income (as in the usual case of insurance analyzed in the literature) but on how the income elasticity of demand for medical care changes over states of the world. Hence, income elasticities of demand for health insurance cannot be used to infer behavior of the Arrow-Pratt risk-aversion measure over income changes.
2. Changes in the price of medical care produce ambiguous effects on demand for $C$. As $p_h$ changes, demand for $h^*$ changes only according to income effects. Thus if there were a compensated change in $p_h$, no change in demand for $h^*$ would be expected. If the price change is uncompensated, no prediction of the effect is possible.
3. Although usual loading fees are not the "true" price of insurance, the effects of an increase in $\theta$ are generally analogous to an increase in the price of any commodity or good. A possible theoretical exception arises when considering demand for $C$, which may be more predominant under some premium-writing schemes than others.

## 3. DEMAND ESTIMATION

Demand curves for medical care have been estimated from a household survey of 2,367 families drawn in 1963 by the Center for Health Administration Studies (CHAS) of the University of Chicago. The families responded to questions about their insurance companies during the household interview, and direct verification of the parameters of the policy was made from the insurance companies. This is the only source of data I am aware of (except the 1970 survey by CHAS, not yet completed) that provides explicit data on parameters of insurance policies purchased by individual families; hence its uniqueness. Unfortunately, verification was not complete because some insurers did not respond and response was nonrandom. Thus, only families with all policies verified have been used for this study, and the data have been weighted to reproduce the actual sample means in terms of socioeconomic parameters known to affect the response rate. There remain 1,579 families in the study, 970 of which have some health insurance. An earlier study using these data (Phelps, 1973) provides more complete descriptions of the data source. Those studies established in general that there was a positive income elasticity of demand for health insurance, varying considerably by type of insurance. It was also established that health insurance was highly sensitive to the loading fee on the insurance, measured in this case by the size of the group through which insurance was obtained. A rough estimate from these data is that the elasticity of insurance with respect to changes in $\theta$ was about $-0.3$ to $-0.5$. The effects of changes in the price of medical care were ambiguous, and the results varied considerably depending on the method of estimation used. Here, I have attempted some additional analyses of these data, using new measures of the price of medical care facing each family, in an attempt to improve those estimates. I have also added measures of the wage rate facing the head of each household derived from OLS estimates of the wage rate, using fitted values for heads not employed.

The medical care price measures are obtained through instrumental variable estimation of prices for those with positive utilization. The estimating equations are in the appendix, Table 8. The dependent variables in these equations are expense/unit of service, and the explanatory variables include dummies for eight geographic regions, measures of physician and hospital supply, wages of the family head, nonwage income, and education of the head. The last variables reflect possible demand for higher-quality service with higher income, and education is a measure of efficiency in search for lower prices. Separate price equations were estimated for hospital services, for surgical care, and for total medical expense. The equations explained from 17 per cent of the variance (hospital price) to 5 per cent (total care).[10]

## 4. DEMAND FOR COVERAGE

From the available data in the 1963 CHAS-NORC survey, one parameter closely matches the theoretical variable $C$—the maximum payment per hospital day chosen by the consumer. This variable can vary from $0 (no coverage) to any positive amount. Some policies specify full coverage of semi-private rooms (or hospital wards), and for these purposes I assigned a dollar figure equal to the 95th percentile of the observed distribution of actual maximum payments. Thus the dependent variable ranges from 0 to $40, the latter figure being the largest maximum in these data.

Demand functions have been estimated using both ordinary least squares (OLS) and the limited dependent variable technique proposed by Tobin (1958), known informally as "tobit" analysis. Because of the nature of the dependent variable in these regressions, tobit analysis is preferable *a priori*. Actually, the tobit regressions produced (in three of four dependent variables) slightly higher (0.3 to 2.8 per cent higher) mean square error than the OLS estimates, the exception being demand for major medical insurance, which had the largest fraction of observations with the dependent variable at the limit. Both sets of regressions are reported here. The results are shown in Table 1, with quadratic terms for the continuous explanatory variables. I show elasticities calculated at mean values of all the $x$'s and of $y$, with $t$ ratios in parentheses. The quadratic specification is desirable on grounds of significance of the quadratic terms and will be discussed exclusively.

In the OLS equation, income is positively and significantly related to insurance demand, the elasticity being 0.22. As mentioned previously, one cannot infer from this that there is increasing risk aversion in income because of complex changes in the real price of insurance that are associated with an income change. One can, however, impute some fraction of that income elasticity to another subsidy; namely, that occurring through the income tax deductibility of insurance premiums on federal and state taxes. The

**TABLE 1 Demand for Maximum Payment per Hospital Day (HOSMAX) ($n = 1522$)**

| | OLS | | TOBIT | |
|---|---|---|---|---|
| Variable | Coefficient (*t* statistic) | Elasticity (*t* statistic) | Coefficient (*z* statistic) | Elasticity (*z* statistic) |
| Income | .644 E-03 | | .141 E-02 | |
| | (1.61) | .22 | (1.97) | .23 |
| Income squared | −.153 E-07 | (2.16) | −.577 E-07 | (2.03) |
| | (.57) | | (1.25) | |
| Wage | .0973 | | .133 | |
| | (2.09) | .25 | (1.66) | .25 |
| Wage squared | −.304 E-03 | (1.91) | −.404 E-03 | (1.72) |
| | (1.65) | | (1.28) | |
| Work-group size | 6.756 | | 10.438 | |
| | (11.73) | .67 | (11.46) | .86 |
| Group size squared | −.531 | (23.53) | −.838 | (17.56) |
| | (8.08) | | (8.17) | |
| Hospital price index | 49.132 | | 89.707 | |
| | (4.61) | .08 | (4.71) | .06 |
| Hospital price squared | −28.768 | (.69) | −53.147 | (.45) |
| | (4.72) | | (4.93) | |
| Estimated illness | −5.563 | | −9.326 | |
| | (2.27) | .04 | (2.16) | .02 |
| Illness squared | .841 | (.24) | 1.415 | (.114) |
| | (2.37) | | (2.29) | |
| Education of head | 1.885 | | 4.164 | |
| | (2.26) | .11 | (2.78) | .22 |
| Education squared | −.153 | (.89) | −.323 | (1.76) |
| | (1.77) | | (2.15) | |
| Age of head | −1.593 | | −2.294 | |
| | (1.89) | .16 | (1.62) | .24 |
| Age squared | .220 | (2.82) | .343 | (3.46) |
| | (2.49) | | (2.30) | |
| Race (1=nonwhite) | −1.933 | | −5.588 | |
| | (2.29) | | (3.68) | |
| Sex of head (1=female) | 2.913 | | 4.095 | |
| | (2.83) | | (2.39) | |

TABLE 1 (concluded)

| Variable | OLS Coefficient (t statistic) | OLS Elasticity (t statistic) | TOBIT Coefficient (z statistic) | TOBIT Elasticity (z statistic) |
|---|---|---|---|---|
| Family size | −.248 | | −.483 | |
| | (1.13) | −.06 | (1.28) | −.08 |
| Urban/Rural (1=rural) | .547 | | .880 | |
| | (.90) | | (.88) | |
| Welfare care (1=yes) | −2.155 | | −5.652 | |
| | (1.48) | | (2.11) | |
| Free care (1=yes) | 1.584 | | −4.109 | |
| | (1.20) | | (1.75) | |
| Constant term | −29.992 | | −74.385 | |
| | (4.20) | | (5.819) | |
| $R^2$ | .426 | | — | |
| $F$ (20, 1501) | 57.836 | | — | |
| $X^2$ (d.f.=20) | — | | 852.66 | |

net income effect is found by solving

$$\frac{dC}{dI} = \frac{\partial C}{\partial I} + \frac{\partial C}{\partial \theta} \cdot \frac{\partial \theta}{\partial I}$$

$$= \frac{\partial C}{\partial I} + \frac{\partial C}{\partial \theta} \cdot \frac{\partial \theta}{\partial \mathrm{MTR}} \cdot \frac{\partial \mathrm{MTR}}{\partial I}$$

or

$$\eta_{CI} = \bar{\eta}_{CI} + \frac{I}{\theta} \cdot \eta_{C\theta} \cdot \frac{\partial \mathrm{MTR}}{\partial I/I}$$

Estimating the effect of income tax deductibility on demand for insurance requires several assumptions. First, for group insurance premiums paid by the employer, I assume that the entire amount is exempted from both income tax and payroll tax.[11] In 1963, personally paid premiums could be combined with out-of-pocket medical expense and deducted subject to a 3 per cent of adjusted gross income (AGI) limitation. Data from Mitchell and Phelps (1974) enable one to estimate how the marginal income tax rate changes with income—the estimate is that $\partial\mathrm{MTR}(\partial I/I) = 0.1$. This estimate applies only to personally paid premiums. For employer premiums, the progression of marginal tax rate is lower because of the incidence of social security taxes; data in Mitchell and Phelps allow an estimate of $\partial \mathrm{MTR}(\partial I/I) = 0.05$ for premiums paid by employers.[12]

For employer-group premiums, the total income elasticity is decomposed as

$$\eta_{CI} = \bar{\eta}_{CI} + \eta_{C\theta} \frac{\partial \mathrm{MTR}}{\partial I/I} \cdot \frac{I}{\theta}$$

where $\bar{\eta}_{CI}$ is the total income elasticity. I have estimated $\eta_{C\theta} = -0.7$ for policies in which $\bar{\theta} \approx 0.2$. Thus $\eta_{CI} = \bar{\eta}_{CI} - 0.7(0.05)(1/0.2) = \bar{\eta}_{CI} - 0.175$. Thus the tax subsidy is a substantial portion of the estimated income elasticity (see below).

For family-paid premiums, the effect is more difficult to estimate. At one extreme, if the family did not anticipate having premiums plus unreimbursed medical expenses exceeding a "deductible" of 3 per cent of AGI, there would be no effect. At the other extreme, if the family knew with certainty that such a deductible would be exceeded, the appropriate correction would be twice that computed for employer-paid premiums, since the marginal tax rate for individuals progresses faster ($\partial \mathrm{MTR}/\partial I/I = 0.1$) when the payroll taxes are excluded. Offsetting this would be higher loading fees paid on many family premiums. The calculation must be made on the anticipations of the family, and hence cannot be computed directly—certainly the expected frequency of taking the deduction measures with error the true perception, since the family has more information about itself than is revealed by average behavior.[13] Under 15 per cent of families in 1963 actually used the medical deduction, the rate generally falling with income, so the estimate that little or no effect is present in the income elasticity estimate is probably more accurate than using the full value.

As a general summary, since about 40 per cent of premiums were paid by employers in 1963, and since that induces an upward bias of 0.1 to 0.25 in the estimated income elasticity, a rough correction factor for the estimates presented might be 0.4 times 0.1 to 0.25, or a reduction of 0.04 to 0.1.

Demand for coverage is quite sensitive to the price of insurance. In these data, the instrument for price is the group size from which the insurance was obtained. Insurance texts contain data on how premiums fall with size of total group premium, from which one can compute how the loading fee falls with the size of the group. For the scaling of group size used in these regressions, it is approximately true over the entire range that $d\ln \theta / d\ln$ (group size) $= -1$, so that the group-size elasticities should be multiplied by $-1$ to obtain an estimate of the own-price elasticity with respect to $\theta$.[14] In the quadratic specification of demand for payment per day, the elasticity with respect to group size at the mean is 0.67, ($t$=23.53), so the elasticity with respect to $\theta$ is $-0.67$.

Of particular interest is the response of payment per day to the price of medical care faced by the consumer (an instrumental variable fitted from an OLS regression). In the linear equation (not shown), there is no response. In the quadratic specification, both the linear and quadratic terms are highly significant, suggesting that the relationship between demand for payment per day and hospital prices first rises and then falls as $p_h$ increases. Thus, as might be anticipated with the complex interactions shown in (27)–(29), the effect of $p_h$ on demand for $C$ varies with $p_h$.

The tobit regression for this equation shows few significant differences. The estimated income elasticity is slightly higher (0.23), and the estimated elasticity with respect to work-group size is 0.86, rather than the 0.67 found in the OLS equation.

## 5. DEMAND FOR MAXIMUM COVERAGE

Three variables measure maximum coverage in these data—the maximum number of hospital days covered, the maximum surgical payment allowed in the fee schedule, and the maximum payment under a major medical insurance plan. Tables 2 through 4 present estimates of demand for each of these variables, with each of the continuous variables entered in quadratic form.

Results are similar for all three. Each is highly elastic to the loading fee, with OLS estimates of the group-size elasticity at the means of the data equal to 0.63 (hospital days), 0.67 (surgical maximum), and 0.68 (major medical insurance maximum), corresponding to elasticities with respect to $\theta$ of the negative of those values. The income elasticity for two of these maximum payment parameters is estimated to be zero, although the wage-income elasticity is positive and significant for hospital days ($\eta$ = 0.29), maximum surgical payment ($\eta$ = 0.48), and for major medical maximum ($\eta$ = 1.07). With wage income omitted, the income elasticities are similar to those reported for wage-rate elasticities (Phelps, 1973).

The response of these variables to medical prices is interesting. The hospital maximum variable is in physical units, whereas the remaining two are in dollar units, so "no response" in the first case is a zero elasticity, and "no response" in the remaining two cases is a unitary elasticity, in which case a 1 per cent increase in $p_h$ would cause a 1 per cent change in the maximum, leaving the "real" maximum unchanged. For the hospital days, the response is curvilinear, first increasing and then decreasing. At the mean price, the elasticity is near zero and not significant. This result, using the same data but an imputed medical price that is area-specific, is similar to one reached by Phelps (1973). The surgical maximum payment has a positive elasticity of 0.64 ($t = 2.98$), but the dependent variable is in dollar terms, so this implies a reduction in the *real* coverage maximum as prices rise. In the major medical maximum equation, the effect of medical prices is estimated to be $\eta = 1.13$ ($t = 0.91$), which is similar to results using the earlier price measure, but less precise. This result suggests a slightly increasing demand for maximum coverage with price increases.

Some secondary evidence is available on the effect of medical prices on demand for insurance. Some of the families in the sample were eligible (for various reasons) for reduced-rate or free medical care, either through welfare or for professional courtesy. Although these variables may be confounded with income (permanent income is held constant in these regressions), the tobit demand curves show systematic strong negative relationships between obtaining some free care and the demand for insurance.[15] It is possible to interpret these free-care sources as very low medical prices, in which case (at low prices) one would expect a positive relationship between medical prices and demand for coverage. This is exactly the case with the actual medical price variables, of course—the effect of price increases is first to increase, and then decrease, demand for coverage. At the average of the sample, as has been mentioned, the effect is essentially zero on demand for $C$, and there exists some evidence for a positive effect on demand for $h^*$.

The remaining variable of interest in these equations is the effect of estimated illness on demand for coverage. Particularly in community-rated insurance, it can be shown that persons with higher *anticipated* illness levels should have demands for more insurance. The estimated loss variable is an instrumental variable regression, showing in essence the average illness level for persons in a given age-sex-income class. Although this variable is of general interest, it is not a sufficient variable with which to test the problem of adverse self-selection against insurance companies. The crucial question for self-selection analysis is whether residuals in a regression study of insurance demand are correlated with residuals in a regression study of medical care demand, and it is this latter

**TABLE 2 Maximum Number of Hospital Days (HOSDAY) (n = 1491)**

| Variable | OLS Coefficient (*t* statistic) | OLS Elasticity (*t* statistic) | TOBIT Coefficient (*z* statistic) | TOBIT Elasticity (*z* statistic) |
|---|---|---|---|---|
| Income | .788 E-02 | | 0.174 | |
| | (1.81) | .39 | | .36 |
| Income squared | −.2085 E-06 | (2.36) | −.7176 E-06 | (2.36) |
| | (.72) | | (1.46) | |
| Wage | .0597 | | −.0425 | |
| | (.12) | .29 | (.051) | .23 |
| Wage squared | .691 E-03 | (1.27) | .0016 | (1.16) |
| | (.34) | | (.474) | |
| Work-group size | 23.526 | | 65.971 | |
| | (3.69) | .63 | (6.98) | .78 |
| Group size | | (8.33) | | (12.13) |
| squared | −1.147 | | −4.703 | |
| | (1.57) | | (4.39) | |
| Hospital price | | | | |
| index | 435.80 | | 854.282 | |
| | (3.76) | .07 | (4.23) | .02 |
| Hospital price | | (.34) | | (.09) |
| squared | −258.54 | | −512.607 | |
| | (3.91) | | (4.48) | |
| Estimated illness | −66.298 | | −114.218 | |
| | (2.47) | −.67 | (2.64) | −.42 |
| Illness squared | 7.442 | (2.38) | 14.315 | (1.64) |
| | (1.93) | | (2.17) | |
| Education | | | | |
| of head | 14.953 | | 43.874 | |
| | (1.63) | −.35 | (2.69) | −.04 |
| Education | | (1.78) | | (.21) |
| squared | −1.984 | | −4.354 | |
| | (2.08) | | (2.67) | |
| Age of head | −3.997 | | −6.349 | |
| | (.44) | .21 | (.424) | .34 |
| Age squared | .800 | (2.03) | 1.754 | (3.50) |
| | (.84) | | (1.113) | |
| Race | | | | |
| (1 = nonwhite) | −19.854 | | −56.043 | |
| | (2.16) | | (3.44) | |

**TABLE 2 (concluded)**

| Variable | OLS Coefficient (*t* statistic) | OLS Elasticity (*t* statistic) | TOBIT Coefficient (*z* statistic) | TOBIT Elasticity (*z* statistic) |
|---|---|---|---|---|
| Sex (1=female) | 22.094 | | 34.811 | |
| | (1.97) | | (1.94) | |
| Family size | 2.640 | .10 | 1.202 | .03 |
| | (1.12) | | (.302) | |
| Urban/Rural (1=rural) | 3.336 | | 5.936 | |
| | (.50) | | (.56) | |
| Welfare care (1=yes) | −21.134 | | −68.860 | |
| | (1.36) | | (2.39) | |
| Free care (1=yes) | −23.531 | | −56.413 | |
| | (1.65) | | (2.22) | |
| Constant term | 184.02 | | −675.867 | |
| | (2.36) | | (4.94) | |
| $R^2$ | .205 | | — | |
| F (20, 1470) | 18.909 | | — | |
| $X^2$ (d.f.=20) | — | | 544.47 | |

problem that TSLS estimators should be helpful in solving. Unless some measure of anticipated health status on each individual were available, studies such as those I have performed here are not likely to show strong effects of illness on demand for insurance. Notice that community-rated insurance is tantamount to hiding from the insurance company information on individual illness, so that if community rating dominated the insurance industry, even the age-sex-race-income specific illness measure I use here should have some association with demand for insurance. To the extent that experience rating predominated, one would expect illness measures to have but slight effects on insurance demand.

In the hospital maximum payment per day equation, the effect at the mean illness level is small, but the quadratic formulation shows the effect to be increasing with illness levels, so that for high expected losses, more insurance would be chosen. In the maximum hospital coverage equation, the effect is negative at mean values of expected illness, but again is increasing with losses. In the surgical maximum equation, the estimated effect is zero at mean values, but again increasing, and there appears to be no significant effect on

**TABLE 3 Demand for Maximum Surgical Payment (SURMAX) ($n$ = 1493)**

| | OLS | | TOBIT | |
|---|---|---|---|---|
| Variable | Coefficient ($t$ statistic) | Elasticity ($t$ statistic) | Coefficient ($z$ statistic) | Elasticity ($z$ statistic) |
| Income | .0150 | | .0362 | |
| | (1.80) | −.03 | (2.33) | .16 |
| Income squared | −.1269 E-05 | (.23) | −.2411 E-05 | (1.00) |
| | (2.25) | | (2.38) | |
| Wage | −.5171 E-02 | | .4018 | |
| | (.00) | .48 | (.23) | .50 |
| Wage squared | .403 E-02 | (2.83) | .0038 | (2.28) |
| | (1.03) | | (.55) | |
| Work-group size | 77.971 | | 177.332 | |
| | (6.41) | .67 | (9.04) | 1.06 |
| Group size squared | −4.572 | (17.62) | −13.0585 | (14.98) |
| | (3.29) | | (5.88) | |
| Surgical price index | −398.51 | | −1049.040 | |
| | (1.15) | .64 | (1.73) | .55 |
| Surgical price squared | 248.52 | (2.98) | 565.093 | (2.03) |
| | (1.51) | | (1.94) | |
| Estimated illness | −142.77 | | −243.17 | |
| | (2.75) | .03 | (2.56) | −.09 |
| Illness squared | 21.235 | (.12) | 35.295 | (.31) |
| | (2.84) | | (2.60) | |
| Education of head | 75.111 | | 126.773 | |
| | (3.30) | .16 | (3.07) | .35 |
| Education squared | −1.46 | (1.02) | −12.423 | (1.83) |
| | (2.91) | | (2.63) | |
| Age of head | −13.121 | | −37.153 | |
| | (.76) | .13 | (1.23) | .27 |
| Age squared | 2.0496 | (1.74) | 5.600 | (2.56) |
| | (1.12) | | (1.75) | |

**TABLE 3 (concluded)**

| Variable | OLS Coefficient (*t* statistic) | OLS Elasticity (*t* statistic) | TOBIT Coefficient (*z* statistic) | TOBIT Elasticity (*z* statistic) |
|---|---|---|---|---|
| Race (1 = nonwhite) | −49.507 | | −136.813 | |
| | (2.79) | | (4.09) | |
| Sex (1 = female) | 33.091 | | 70.733 | |
| | (1.53) | | (1.89) | |
| Family size | −2.485 | −.04 | −4.288 | −.05 |
| | (.54) | | (.51) | |
| Urban/Rural (1 = rural) | −8.719 | | −11.038 | |
| | (.70) | | (.51) | |
| Welfare care (1 = yes) | −43.181 | | −121.331 | |
| | (1.41) | | (2.04) | |
| Free care (1 = yes) | −7.614 | | −68.743 | |
| | (.27) | | (1.31) | |
| Constant term | 44.332 | | −251.054 | |
| | (.22) | | (.68) | |
| $R^2$ | .319 | | — | |
| $F$ (20, 1472) | 36.428 | | — | |
| $\chi^2$ (d.f. = 20) | — | | 702.25 | |

demand for major medical maximum payment. Thus, the simultaneous aspect of medical demand and insurance demand would appear to be particularly acute for persons with very high expected illnesses, but not severe at mean values of illness in this population. Nevertheless, simultaneous equation methods are indicated in medical demand studies because of this association.

The tobit regressions in general show similar results for each of these three equations. The differences will be highlighted here. First, for hospital days the estimated income elasticity is positive (0.36) and significant ($z$ = 2.36), where $z$ is a unit normal variate. The elasticity with respect to work-group size is slightly higher. One surprising result is that the demand for hospital days coverage falls as a function of estimated increase in illness; the elasticity is −0.42 ($z$ = 1.64). In demand for maximum surgical coverage, the

## TABLE 4 Demand for Major Medical Maximum Payment (MM-MAX)

| Variable | OLS Coefficient (*t* statistic) | OLS Elasticity (*t* statistic) | TOBIT Coefficient (*z* statistic) | TOBIT Elasticity (*z* statistic) |
|---|---|---|---|---|
| Income | −.2572 | | .6963 | |
| | (1.13) | −.15 | (.45) | .47 |
| | | (.45) | | (.97) |
| Income squared | .1713 E-04 | | .3879 E-04 | |
| | (1.11) | | (.03) | |
| Wage | 17.775 | | 218.9687 | |
| | (.67) | 1.07 | (1.37) | 1.67 |
| Wage squared | .9947 E-02 | (2.31) | −.4118 | (2.63) |
| | (.10) | | (.70) | |
| Work-group size | −172.75 | | 5438.822 | |
| | (.52) | .68 | (3.65) | 1.45 |
| | | (4.55) | | (7.82) |
| Group size squared | 96.89 | | −249.850 | |
| | (2.57) | | (1.53) | |
| Medical price index | 5241.1 | | 21494.78 | |
| | (.59) | 1.13 | (.55) | 1.77 |
| | | (.91) | | (1.09) |
| Medical price squared | −1429.6 | | −3252.15 | |
| | (.41) | | (.23) | |
| Estimated illness | 882.560 | | 4039.55 | |
| | (.63) | .69 | (.46) | 1.01 |
| Illness squared | −67.399 | (1.19) | −220.22 | (1.26) |
| | (.33) | | (.18) | |
| Education of head | −503.46 | | −807.502 | |
| | (1.00) | .11 | (.25) | −.14 |
| | | (.27) | | (.26) |
| Education squared | 52.827 | | 55.503 | |
| | (1.06) | | (.19) | |
| Age of head | 179.00 | | −1764.978 | |
| | (.38) | −.48 | (.67) | −.87 |
| Age squared | −41.137 | (2.25) | 22.441 | (2.59) |
| | (.83) | | (.08) | |

**TABLE 4 (concluded)**

| Variable | OLS Coefficient (*t* statistic) | OLS Elasticity (*t* statistic) | TOBIT Coefficient (*z* statistic) | TOBIT Elasticity (*z* statistic) |
|---|---|---|---|---|
| Race (1 = nonwhite) | 368.12 | | 1668.83 | |
| | (.77) | | (.59) | |
| Sex (1 = female) | 840.80 | | 9120.07 | |
| | (1.45) | | (2.98) | |
| Family size | −141.54 | −.21 | −463.855 | −.18 |
| | (1.14) | | (.66) | |
| Urban/Rural (1 = rural) | 198.55 | | 129.116 | |
| | (.59) | | (.07) | |
| Welfare care (1 = yes) | −77.644 | | −9637.624 | |
| | (.09) | | (1.55) | |
| Free care (1 = yes) | −541.21 | | −2133.249 | |
| | (.71) | | (.48) | |
| Constant term | −6289.0 | | −1.007 E-07 | |
| | (1.07) | | (3.24) | |
| $R^2$ | .171 | | — | |
| $F$ (20, 1558) | 16.096 | | — | |
| $X^2$ (d.f. = 20) | — | | 314.64 | |

only difference is a 20 per cent increase in the estimated own-price elasticity of demand for coverage, the elasticity with respect to work-group size being 1.06 ($z$ = 14.98).

In demand for major medical insurance, the tobit regression should be and is superior to OLS because in a large fraction of the observations the dependent variable is clustered at the limit. The estimated regressions are considerably different. The income elasticity of demand for major medical maximum is 0.47 ($z$ = 0.97), and the estimated elasticity with respect to wage rate is 1.67 ($z$ = 2.63). Taken together, these suggest an extremely high income elasticity of demand for major medical insurance. Demand for major medical coverage is also quite sensitive to the price of medical care, with an elasticity of 1.77 ($z$ = 1.09), an extremely large elasticity compared to other results found here even if precision is low. Demand for major medical coverage also rises with estimated illness, the elasticity being 1.01 ($z$ = 1.26).

Taken together, these results suggest that persons facing high medical care prices, particularly those with high incomes, shift insurance coverage from basic medical care coverage (such as hospital or surgical insurance) into a major medical insurance plan. This accounts for the relatively low response of the basic hospital and surgical coverage parameters to medical care prices and the quite high response of major medical coverage. It should be recalled that these estimates are derived from data obtained for 1963, at which time less than 20 per cent of the population had major medical insurance. That ratio now exceeds 40 per cent, so these values may have changed considerably between 1963 and the present.

## 6. OBSERVED COINSURANCE RATES

The 1963 CHAS-NORC data also permit estimation of the observed coinsurance rates resulting from insurance policies held by the families, and from their actual medical purchases during the year. Unlike the insurance parameter equations presented above, these equations contain a considerably random component induced by variation in actual health outcomes during the year; hence, their precision is much lower than the previous estimates. The dependent variable takes on values from 0 to 1 inclusive, expressed as benefits/expenses = coverage rate. Observations with zero expenses in each category studied were deleted, since no data are available to compute a coverage rate for them. Table 5 shows maximum likelihood logistic regressions for coverage rates on total expense, hospital expense, and physician office expenses. The maximum likelihood estimator relates the dependent variable coverage to explanatory variables in a fashion analogous to normal regression, the major difference being that the error term in this maximum likelihood regression has a logistic distribution, rather than the usual normal distribution. The partial derivative is $\partial y/\partial X_i = \beta_i y\ (1-y)$ and the elasticity is $\eta_{yX_i} = \beta_i \cdot X_i \cdot (1-y_i)$.

For total expense (representing all medical outlays by the family), the average coverage rate in 1963 was 13 per cent. The equation was estimated in linear form and with quadratic terms for continuous variables. A chi-square test on the vector of quadratic terms is 24.83 (7 d.f.), rejecting the hypotheses that these coefficients are

**TABLE 5 Observed Coverage Percentage (Logistic Regressions)**

| Explanatory Variables | All Medical Care: Coefficient (z value) | All Medical Care: Elasticity (z value) | Hospital Care[a]: Coefficient (z value) | Hospital Care[a]: Elasticity | Medical Office Visits: Coefficient (z value) | Medical Office Visits: Elasticity (z value) |
|---|---|---|---|---|---|---|
| Nonwage income | .796 E-04 | | −.298 E-04 | −.01 | .6848 E-04 | |
| | (1.45) | .05 | (1.11) | | (.97) | .05 |
| Nonwage income squared | .1966 E-08 | (1.30) | | | −.1247 E-08 | (.99) |
| | (.58) | | | | (.49) | |
| Wage/Week | .0181 | | .00918 | .31 | .0254 | |
| | (1.51) | .71 | (1.66) | | (1.26) | .34 |
| Wage squared | .452 E-04 | (1.76) | — | | −.8464 E-04 | (.63) |
| | (1.01) | | | | (1.12) | |
| Work-group size | .5011 | | .2532 | .31 | .6409 | |
| | (4.02) | .58 | (6.13) | | (3.54) | .64 |
| Group size squared | −.0419 | (5.94) | — | | −.0563 | (4.11) |
| | (3.06) | | | | (2.89) | |
| Medical price index | −10.24 | | −.840 | −.20 | .0347 | |
| (service-specific) | (.56) | −1.58 | (1.05) | | (1.67) | .24 |
| Price index squared | 4.190 | (1.28) | — | | −.2148 E-03 | (2.08) |
| | (.48) | | | | | |
| Estimated illness level | 1.376 | | −.3073 | | .1222 | |
| | (1.78) | .40 | (1.15) | −.32 | (.10) | .29 |
| Estimated illness squared | −.1682 | (.82) | | | −.0306 | (.38) |
| | (1.57) | — | | | (.19) | |

| | | | | | | |
|---|---|---|---|---|---|---|
| Education of head | .789 | | .0400 | .05 | .8974 | |
| | (2.58) | .09 | (.37) | | (1.90) | −.33 |
| Education squared | −.0759 | (.25) | — | | −.0868 | (.61) |
| | (2.71) | | | | (1.97) | |
| Age of head | −.4078 | | −.0176 | −.02 | .2067 | |
| | (2.06) | .04 | (.22) | | (.64) | .13 |
| Age squared | .0427 | (.20) | — | | −.0188 | (.39) |
| | (1.94) | | | | (.52) | |
| Race | −.164 | | −.40 | | −.0292 | |
| (nonwhite = 1) | (.62) | | (.91) | | (.07) | |
| Sex | .186 | | .160 | | .008 | |
| (female = 1) | (.68) | | (.39) | | (.02) | |
| Family size | .0865 | .25 | .0677 | .07 | .1012 | .33 |
| | (1.53) | | (.75) | | (1.23) | |
| Urban/Rural | .0245 | | −.109 | | −.1088 | |
| (rural = 1) | (.16) | | (.44) | | (.50) | |
| Welfare care | −.87 | −1.48 | −1.545 | −.85 | −.3645 | |
| (= 1 if yes) | (1.80) | | (2.63) | | (.56) | |
| Free care | .28 | | .108 | | .3195 | |
| (= 1 if yes) | (.85) | | (.20) | | (.73) | |
| Constant term | −3.622 | | −2.416 | | −9.000 | |
| | (.39) | | (1.60) | | (3.09) | |
| $x^2$ (d.f.) | 107.30 (21) | | 75.948 (14) | | 54.217 (21) | |
| (probability) | (.00000) | | (.00000) | | (.00009) | |
| $E(y)$ | .129 | | .717 | | .068 | |
| # Observations | 2328 | | 489 | | 1839 | |

[a]Non-quadratic equation given $-X^2$ test on quadratic coefficients is $X^2_7 = 7.88$ ($P \approx 0.4$). Hence, quadratic specification is rejected.

jointly zero; hence, the quadratic specification is presented. Three variables have effects significantly different from zero—the loading fee instrument (group size), with an elasticity at the mean of 0.58 ($z = 5.94$); receipt of welfare care, with an elasticity of $-1.48$ ($z = 1.80$); and wage income, with an elasticity of 0.71 ($z = 1.76$). The price of medical care had no apparent effect on demand for coverage, either in the linear or quadratic specifications. Nonwage income had an elasticity of 0.05 ($z = 1.30$), so the combined income elasticity is near 0.75. Age and education both show significant coefficients in the quadratic specification (but not the linear), but the combined effect of these variables in combination is not significant.

The coverage of hospital expenses is similar. Here, the quadratic terms are not significantly different from zero, so the linear specification is presented. The price of insurance is again the major predictor of demand, with an elasticity of 0.31 on the group-size variable ($z = 6.13$). The wage income elasticity is 0.31, under half that for total expense ($z = 1.66$), and receipt of welfare care has a negative effect ($\eta = -0.85, z = 2.63$). The price of hospital care had a negative but insignificant effect on demand for coverage ($\eta = -0.2, z = 1.05$), as is true for the expected illness variable.

In 1963 medical office visits were covered at a much lower rate than hospital care and surgical care, the average coverage being slightly under 7 per cent. However, as with hospital care, the major predictor both in terms of statistical precision and magnitude is the work-group size. This instrument for the loading fee of insurance has an elasticity of 0.64 ($z = 4.11$). This implies an own-price elasticity of demand for physician office coverage of about $-2/3$, quite elastic compared with estimates of demand for physician office care itself (Newhouse and Phelps, 1974; Phelps and Newhouse, 1974). The major difference between demand for physician office coverage and for hospital coverage is that the price of office care is positively and significantly related to demand for coverage. The elasticity is estimated to be 0.24 at the mean ($z = 2.08$). This difference between demand for physician office coverage and other types of care may well be related to the average existing coverage. As pointed out in the theoretical discussion earlier in this paper, the response of the demand for coverage to medical prices may well be nonlinear, and this presents some evidence that indeed the coverage is positively related to price of medical care when the coverage level was initially low. As coverage rises, as for example in the case of hospital care, the response to medical price may fall off.

How do these estimates compare with those using insurance policy parameters as dependent variables? In both sets of estimates, the primary predictor in terms of both explanatory power and magnitude is the price of insurance. Maximum payment parameters have an own-price elasticity of between −0.5 and −1.0, and the "coverage" types of parameter HOSMAX has an own-price elasticity of about −0.7. The coverage parameters obtained from the logistic estimates using expenditure and benefit data show own-price elasticities of about −0.4 for total medical expense, with smaller values for component expenditures (hospital, M.D. office, etc.). The value of −0.4 is about half that obtained from annual time series of aggregated benefits and expenditures (Phelps, 1973), although the mean loading fee in the time series is considerably lower than in the 1963 data used here, and the elasticity is a function of $\theta$ in the logistic specification of the dependent variable. From the time series estimates in Phelps (1973), the fitted own-price elasticity for 1963 for coverage was −0.26, considerably closer to the above estimate than the time series value taken at the mean. Similarly, the income elasticity at the mean from the time series estimate was 0.39, somewhat lower than the combined wage and nonwage elasticities taken from these results. The time series results on the effects of medical prices on demand for coverage are inconclusive, being highly sensitive to functional form employed. In the lowest mean-squared-error equation (Phelps, 1973, Table 18, LOGIT equation), the elasticity at the mean of coverage with respect to $p_h$ was 0.83 ($z = 2.53$), suggesting a strong effect of medical prices on insurance demand. The value using 1963 levels of the explanatory variables is 0.79. This result differs considerably from the estimates found here—that medical prices have no measurable effect on insurance demand except for M.D. office coverage. The question can be raised, of course, concerning measurement error in the medical price variable used here, since it is fitted from a regression estimate of persons with positive use of medical care, and represents, in effect, the average price paid by persons in a given region and of specific socioeconomic grouping. Other medical price variables used in Phelps (1973) yield results similar to those found here, although those other variables are conceptually similar to the instrument used in this paper—they are regional averages, adjusted for income differences.

Fuchs and Kramer investigated demand for physician office care insurance using interstate aggregated data wherein the dependent variable is the absolute level of benefits (BEN). They found BEN to be highly related to income ($\eta = 0.76$ to $1.61$) and also highly

**TABLE 6 Summary Statistics—Insurance Policy Regressions**

| Variable | Mean | Standard Deviation |
|---|---|---|
| HOSMAX[a] | 13.15 | 12.51 |
| HOSDAY[b] | 79.15 | 111.36 |
| SURMAX[c] | 203.26 | 242.15 |
| MM-MAX[d] | 2108.17 | 5709.57 |
| Income | 5897.19 | 3100.88 |
| Education of head | 5.16 | 1.67 |
| Age of head | 4.74 | 1.91 |
| Medical price index | 1.00 | .09 |
| Physician price index | 1.06 | .16 |
| Hospital price index | .83 | .18 |
| Estimated illness | 3.34 | .752 |
| Work-group size (scaled) | 3.13 | 2.92 |
| Free care | .068 | .25 |
| Welfare care | .059 | .24 |
| Nonwhite | .139 | .34 |
| Female head | .217 | .41 |

[a] $n = 1522$
[b] $n = 1491$
[c] $n = 1493$
[d] $n = 1579$

own-price sensitive—on the order of −1.58 to −1.74. They also found benefits strongly negatively related to medical prices ($\eta$ = −0.74, see their Table 18, Equation (D.4)). Since actual medical purchases also fall with medical prices ($\eta$ = −0.1 to −0.3, tending around about −0.25), this means that coverage (BEN/EXPENSE) falls with medical prices as well, in contradiction to the large time series estimate in Phelps (1973) for total medical expenses, and to the smaller positive effect found here for M.D. office coverage. These conflicting results suggest that there still remains a considerable amount to be learned about the effects of medical prices on demand for insurance, and unfortunately, the theory provides little guidance in this matter.

Other researchers have estimated demand curves for insurance as well. M. Feldstein (1973) concluded that "a rise in the price of hospital services causes an increase in the proportion enrolled and in the total quantity of insurance." His estimated elasticity of "quantity" with respect to medical prices is 0.39, with a long-run

**TABLE 7 Summary Statistics for Observed Coinsurance Regressions**

| Variable | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|
| Race (1 = nonwhite) | .134 | .341 | 0 | 1 |
| Age of head (scaled) | 4.66 | 1.846 | 1 | 8 |
| Education (scaled) | 5.32 | 1.617 | 1 | 8 |
| Sex of head (1 = female) | .191 | .393 | 0 | 1 |
| Welfare care (1 = yes) | .043 | .204 | 0 | 1 |
| Free care (1 = yes) | .0567 | .231 | 0 | 1 |
| Family size | 3.33 | 1.892 | 1 | 16 |
| Rural/Urban (rural = 1) | .308 | .462 | 0 | 1 |
| Permanent income | 6349.2 | 3047 | −3795 | 16898 |
| Wage income/week | 116.16 | 37.97 | −12.743 | 220.24 |
| Work-group size (scaled) | 3.802 | 3.048 | 1 | 8 |
| Estimated illness level | 3.411 | .730 | 1.157 | 6.316 |
| Nonwage income | 758.42 | 2169 | 0 | 7000 |
| Coverage percentage | .1278 | .239 | 0 | 1 |
| Medical prices (estimated) | 1.00 | .085 | .767 | 2.382 |
| Sample size: 2,328 families with positive medical expense (2,367 original) | | | | |
| For M.D. office visits, add the following: | | | | |
| M.D. office coverage percentage | .068 | .200 | 0 | 1 |
| M.D. office price index | 7.210 | 7.208 | .25 | 125.00 |
| Sample size: 1,839 families with positive M.D. office expense | | | | |
| For hospitalizations, add the following: | | | | |
| Hospital coverage percentage | .717 | .361 | 0 | 1 |
| Hospital price index | .835 | .160 | .393 | 1.269 |
| Sample size: 489 families with positive hospital expense | | | | |

NOTE: For M.D. office and for hospital coverage percentage equations, summary statistics on independent variables differ slightly from those given for the sample of 2,328 families.

elasticity calculated to be 1.21. He also estimates an own-price effect on demand for insurance of −0.09 (and insignificant), although he estimates a strong positive relationship between an instrument for group insurance availability (percentage of employees in each state who work in manufacturing or government), and collinearity between this variable and his measure of

**TABLE 8 Estimating Equations for Medical Price Variables**

| Variable | Medical Price Coefficient (*t* statistic) | Hospital Price Coefficient (*t* statistic) | Physician Price Coefficient (*t* statistic) |
|---|---|---|---|
| New England | −.1191 | −.2621 | −.0946 |
| | (3.01) | (2.95) | (.48) |
| Mid-Atlantic | −.1143 | −.2819 | −.1265 |
| | (4.53) | (4.93) | (1.16) |
| E. N. Central | −.1138 | −.1642 | −.2236 |
| | (4.35) | (2.91) | (1.98) |
| W. N. Central | −.0595 | −.2414 | −.0800 |
| | (1.92) | (3.68) | (.56) |
| S. Atlantic | −.0491 | −.2573 | .0258 |
| | (1.76) | (4.17) | (.22) |
| E. S. Central | −.0636 | −.3911 | −.1588 |
| | (1.69) | (4.40) | (.61) |
| W. S. Central | −.0573 | −.1418 | −.0580 |
| | (1.81) | (2.06) | (.40) |
| Mountain | −.0644 | −.0138 | −.2022 |
| | (1.43) | (.14) | (.96) |
| Wage income | .8274 E-03 | .5911 E-03 | −.8416 E-04 |
| | (3.29) | (.98) | (.06) |
| Nonwage income | .1942 E-04 | −.2823 E-05 | −.3536 E-05 |
| | (2.70) | (.16) | (.10) |
| Education of head | .07210 | .0534 | −.2523 |
| | (3.08) | (.96) | (1.66) |
| Education squared | −.0056 | .0040 | .2267 |
| | (2.46) | (.75) | (1.61) |
| M.D.'s/100,000 | .7919 E-03 | .00244 | .00213 |
| | (4.27) | (5.98) | (2.47) |
| Beds/1,000 | .0013 | .00918 | −.00414 |
| | (.34) | (1.21) | (.26) |
| Constant | .6717 | .5650 | 1.5618 |
| | (9.14) | (3.20) | (3.18) |
| $R^2$ | .05 | .17 | .09 |
| $F$(d.f.) | 8.036 (14,222) | 8.24 (14,556) | 1.65 (14,229) |
| $E(y)$ | 1.00 | .83 | 1.04 |

own-price may preclude precise estimation of either.[16] Given the conflicting results of my cross-section survey data studies, my aggregate time series studies, and the state cross-section studies of Feldstein, Frech, and Fuchs and Kramer, it is clear that much remains to be learned about demand for insurance, particularly about the effects of medical prices on demand. The question is of special significance because of possible supplementation of universal (if only partial) coverage that may be introduced under national health insurance. If supplementation is highly responsive to medical prices, then large demand shifts in the medical care system could be exacerbated by additional demand induced through purchase of supplemental insurance policies.

## NOTES

1. It is easy to show that $h(l)$ is an increasing function of $l$—more illness heads for more medical care purchase—as long as the income elasticity of demand for health ($H$) is positive. See Phelps (1973) for proof.
2. The proof of this is in Phelps (1973).
3. It is not the case that $h(l) = \bar{h}$, the average medical care consumption in insured states.
4. In Ted Frech's comment, he points out that "moral hazard" (I interpret this to mean non-zero price elasticity) acts like a tax on the loading fee. This is true in one sense, but the analogy is not perfect. At least in formulations of an insurance premium that I have pursued, the "tax" of the non-zero elasticity applies only to a portion of a premium—that for insured states of the world under $h^*$ (obviously), and that portion of demand that is added when C is lowered. Reducing $C$ would increase the premium (i.e., $R_c$ would be negative) even if $\eta_{hC}$ were zero, since the insurer would pay a larger fraction of the bills. Notice also that the "tax" of the price elasticity is related to the level of insurance—it is highest when $C$ approaches zero and when $h^*$ is very large.
5. The optimality always holds in demand for $h^*$ if the consumer is risk averse. That is, if (5) is zero, then the denominator of (8) is negative. However, in demand for $C$, the optimality is guaranteed for solutions of (4) equal to zero only if demand for $h$ is perfectly price inelastic—the case considered by most previous analyses of insurance. When $\partial h/\partial C$ is negative, the denominator of (9) must be assumed to be negative, even if the consumer is risk averse.
6. It can be shown that $R_{h^*}$ is not a function of income, with the insurance premium functions I consider below.
7. For notational simplicity, the dependence of $\lambda$ and $r(I)$ on $l$ is dropped here.
8. This has been observed in a multiple regression study of state market shares of the Blues; state income is positively and significantly related to market shares of Blue Cross and Blue Shield, *ceteris paribus*. (Phelps, unpublished paper.)
9. I make the simplifying assumption that

$$\frac{\partial^2 h}{\partial C\, \partial p_h} = 0$$

10. In Phelps (1973), the measure of price was derived from BLS statistics on expenditure for medical care in forty-six cities and rural areas. These data were used to construct a price index for each of the primary sampling units (PSU) from which the 2,367 families were drawn. The measure attributed the same regional price to each person living in a PSU, a measure that is considerably less person-specific than the regression estimation of price I use here.

11. If the person earns more than the base amount of the payroll (social security) tax, the appropriate marginal tax rate does not include the payroll tax rate. In 1963, the maximum earnings taxed by the payroll tax were $4,800.
12. The first equation estimated $y$ = average marginal income tax rates. The equation estimate was $y = -0.61 + \underset{(t\,=\,9.52)}{0.09} \log(\text{income})$ $R^2 = 0.84, N = 19$ income groups. Data on income groups were obtained from IRS files. The second equation estimated used interval brackets reporting total marginal tax rates (income plus payroll), with the midpoint of income intervals used as the explanatory variable. The estimated equation was $y = -0.09 + 0.04 \times \log$ (income) $R^2 = 0.46, N = 8$.

    These data reflect 1970 incomes. Because the social security cutoff income in 1963 was lower, there may have been more progressiveness in 1963 than is indicated by the second estimate. For this reason, I use 0.05 as an estimate of $\partial \text{MTR}/\partial I/I$.
13. This problem is analogous to how a family behaves when faced with random medical expenses and is covered by an insurance policy with a deductible. For a discussion, see Keeler, Newhouse, and Phelps (1974).
14. See Mitchell and Phelps (1974) for details of this computation. In brief, the calculation is this: A relationship between loading fee and work-group size is derived from an insurance textbook. The data are then fitted with an OLS equation of the form $ln(\theta) = a_0 + a_1 ln$ (GROUPSIZE). The estimated $\hat{a}_1$ was $-0.998$ ($t$ =9.39), which is $\partial\theta/\partial G \cdot G/\theta$. The inverse of this is the correction factor to convert group-size elasticities to loading fee for elasticities.
15. These coefficients were systematically lower and had smaller $t$ ratios in the OLS regressions than in the tobit regressions presented here. In the four demand equations shown, none of the free-care or welfare-case coefficients were significant at usual significance levels in the OLS regressions, whereas the welfare coefficients are all significant at least at $p = 0.12$ and the free-care coefficients are significant in the hospital insurance demand equations with $p < 0.10$ in the tobit regression.
16. Goldstein and Pauly, in a paper presented at this conference, estimate a loading fee elasticity of $-0.33$, where the dependent variable is total premiums. Their data do now allow an estimate to be made of the effect of medical prices on insurance demand. Frech, in a comment on this paper, reports studies using aggregate state data, estimates an own-price elasticity for insurance very near zero, and an income elasticity of 0.17. He further reports an estimated effect of hospital prices on demand for insurance to be $-0.38$, quite in contrast to the figure derived by M. Feldstein using similar data. The dependent variable in Frech's work is the proportion of hospital expenses paid for by insurance, directly analogous to the hospital equation in my Table 5.

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# 4 COMMENTS

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This is a very interesting paper. It represents an attempt to properly formulate a model of demand for health insurance, making use of the fundamental insight from the Arrow-Pauly discussion of the 1960s that health insurance is characterized by moral hazard and is *not* merely an aggregation of contingent claims. This is an important enterprise because all insurance and many other financial contracts contain some sort of moral hazard. So the theory of moral hazard is important for the analysis of financial markets in general.

For policy purposes, it is important to know the relative extent of moral hazard (wherein more insurance increases the expected medical expenditure) versus adverse selection (wherein higher expected medical expenditures increase the extent of insurance). If adverse selection is relatively important, incomplete insurance and interpersonal variation in extent of insurance may reflect inefficient use of resources in trying to detect individual differences in risk and adjusting insurance contracts to minimize adverse selection. If so, a mandatory program of reasonably complete insurance for everyone may be efficient. It offers the hope of avoiding the resources spent in merely discovering individual differences in risk and adjusting contracts to such differences. Such private expenditures have a private return, but no social return. If moral hazard is relatively important, optimal insurance would be incomplete and vary a great deal on the basis of varying tastes for risk and, especially, varying elasticities of demand for the insured service (medical care).

Moral hazard works like a tax or extra loading charge on the insurance, which increases with the extent of insurance. Thus, it leads to the private choice of incomplete insurance, given the choice of various alternative insurance contracts. This private choice is socially optimal if (1) there is no alternative method of organizing the sale of insurance that will eliminate the moral hazard—the usual case—and (2) insurance is supplied competitively. Thus, moral hazard destroys the optimality of complete insurance, even if the insurance is priced at expected loss. A mandatory program of complete insurance will lead to inefficient moral hazard losses.

In the Phelps paper, moral hazard enters the model because of the form of the benefits—a subsidy for the purchase of medical care, not a simple contingent claim. This is clear in the budget constraint under insurance, his equation (2):

$$I = x + Cp_h h + R$$

Most of Phelps' results for insurance demand under moral hazard are quite reasonable. However, there is one result that seem a bit odd to me. The result would seem to hinge on the way in which the moral hazard welfare loss varies over the extent of insurance.

The result that concerns me is: The optimal limit $h^*$ is finite if the insurance load is positively related to the size of the expected benefits. This result indicates that the marginal moral hazard loss is increasing in the extent of insurance.

The easy way to see this is to first imagine a situation in which the insurance is simply an aggregate contingent claim. In this case, the true price of deepening the coverage (raising the $h^*$) is simply the expected value of the loss plus some proportional load. The consumer will keep increasing $h^*$ without limit as long as risk aversion does not fall as income decreases. This is the basis for the standard argument that the optimal insurance, in the presence of transactions costs, involves a deductible and otherwise full coverage—placing a floor on *ex post* income.

For Phelps' result, the rising real price of insurance as $h^*$ increases must be ascribed to rising marginal moral hazard, since load is held constant or is falling. It seems clear that his model is characterized by a marginal moral hazard loss that begins at zero when there is no insurance and rises as insurance becomes more complete, finally choking off further insurance purchases at a finite upper limit.

This is counter to the views of such scholars as Mark Pauly and Martin Feldstein, who favor insurance with large deductibles and complete insurance for large losses. Furthermore, it is counter to the generally held notions about the nature of medical care. The usual view is that the elasticity of demand for relatively minor elective medical problems is large, in part because there are many reasonable ways of treating the conditions. For more serious problems, the technological possibilities are much more limited so that the ill effects of moral hazard are much reduced. Furthermore, for really large losses, the possibilities for more than usual treatment often involve experimental techniques. For these cases, a richer model than simple consumer choice of level of medical care may be useful.

In any case, the usual view is that marginal moral hazard *decreases* with increasing loss. If that view is correct, then Phelps' proof of the optimality of a finite insurance limit is no help. We are left with arguments about (1) the ability to use private and government charity as the size of loss increases and (2) a low marginal utility of wealth and/or medical care for consumers who have experienced heavy losses. The latter argument is made persuasively in Dick Zeckhauser's recent article on catastrophic insurance (1973).

As an aside, I might mention that Phelps' proof here is more general than he supposes. He states that the optimal $h^*$ is finite only if part of the insurance load is proportional to the expected benefits. However, his proof turns on the sign of the partial derivative of the true price of insurance with respect to $h^*$. From his Equation (17) it is clear that his result holds if the loading, $\theta$, is equal to zero. Thus, the optimal $h^*$ would be finite even if insurance were available at actuarily fair rates.

In my own work, I have formulated a model of the interrelationships of health insurance regulation, health insurance markets, and health care markets that run roughly as follows. A major effect of health insurance regulation is to give nonprofit medical provider-controlled insurers (Blue Cross and Blue Shield) a competitive advantage over commercial insurers. Because of their legal status, these insurers cannot retain the monopoly rents that these regulatory advantages make possible. However, these nonprofit insurers can benefit the providers by raising the demand curve for medical services by offering only relatively complete insurance in the market. This overly complete (in terms of consumer preferences) insurance leads to increased demand for health services, as it is intended to.

In order to empirically examine the model, it is necessary to construct a model of the interactions of the insurance and medical care markets—otherwise alternative explanations (to the regulatory one) cannot be disputed. I have used state data for the year 1969, the last year for which complete data are available. The econometric model has five equations and is estimated by two-stage least squares. I will present some preliminary results that bear on Phelps' work. This is especially valuable, since econometric analyses so often are not robust over different data sets. And I basically corroborate his findings.

In this work, I have simply taken the average proportion paid by third-party payers as the measure of the extent of insurance. Also, my work is limited to hospital care. The demand equation determining the extent of insurance in my system is (standard errors in parentheses):

$$\mathrm{I} = 0.518 - \underset{(0.404)}{0.0540}\ \mathrm{PRINS} + \underset{(0.00004)}{0.00012}\ \mathrm{INC} - \underset{(0.0032)}{0.0065}\ \mathrm{PHOS} + \underset{(0.210)}{0.404}\ \mathrm{BCMSR}, \tag{1}$$

$n = 46$

where

I = proportion of hospital expenses paid by insurers
PRINS = price of insurance
INC = per capita disposable income
PHOS = price of hospital care (endogenous)
BCMSR = Blue Cross market share (endogenous)

The coefficient of variation is not reported because it is meaningless in two-stage estimation.

Phelps found an income elasticity of demand for health insurance of about 0.23. In my equation, that elasticity is slightly lower, about 0.17. Furthermore, the standard error is less than half the estimated coefficient.

Turning to the estimated price elasticity of demand for insurance, the estimated elasticity in my data is very low, −0.004 when defined the same as Phelps' variable. Feldstein (1973) also found a very low price elasticity in his work, using somewhat similar data. Phelps finds a much higher price elasticity of −0.5. The reconciliation seems to lie in the high collinearity between income and price in state data. High income states have larger employment groups and thus lower insurance prices.

I find a strong negative influence of the price of hospital care on the extent of insurance, with an elasticity equal to about −0.38. This is the opposite of what Feldstein (1973) found. The standard error is less than half the estimate in absolute value. Phelps found essentially no effect when he entered the price of hospital care in a linear fashion, but a significant effect when he used a more flexible quadratic formulation. I tried that in my equation, but the squared term did not do well, in part because of collinearity problems.

Another variable that I tested was the expected illness, measured by the average hospital expenditures of the state. There was virtually no effect on insurance demand. My results, as do Phelps', clearly show that the adverse selection problem in health insurance is not terribly important. This finding makes the case for compulsory relatively complete health insurance a more difficult one to make. This may be the most important finding in all this work.

At this point I cannot fail to mention that the Blue Cross market share, holding constant market influences, tends to increase the extent of insurance held by consumers. This is consistent with the main argument of my work.

Phelps raised the question of how insurers price their product. He believes that load is roughly proportional to the amount of the expected benefits, but does not have the necessary data to examine the question. The question is of some interest since it will certainly affect consumer choice. Information on insurance pricing might also be useful in evaluating some of the differences between Blue Cross and commercial insurer behavior. My data allow me to directly investigate the matter. Following are regressions of per capita insurance load (LD, LDBC, LDC) on expected benefits (BEN, BENBC, BENC) for all insurers, Blue Cross insurers, and commercial insurers, respectively. The equations were estimated across states for 1969. In the overall regression, BCMSR is also entered to standardize for the lower load of Blue Cross firms. Both BCMSR and the benefit measures are treated as endogenous variables.

$$\text{LD} = \underset{(3.434)}{21.396} + \underset{(0.076)}{0.178}\ \text{BEN} - \underset{(9.414)}{39.763}\ \text{BCMSR} \qquad n = 41 \tag{2}$$

$$\text{LDBC} = \underset{(4.540)}{-4.607} + \underset{(0.032)}{0.103}\ \text{BENBC} \qquad n = 41 \tag{3}$$

$$\text{LDC} = \underset{(4.590)}{9.657} + \underset{(0.072)}{0.216}\ \text{BENC} \qquad n = 41 \tag{4}$$

The results seem quite reasonable and bear out Phelps' hunches. Much of the loading charge is proportional to the expected benefits. In the case of Blue Cross insurance, for which selling expenses are low, the bulk of the loading charge seems to be proportional.[1] A more interesting study, for which I do not have adequate data, would examine the marginal load for large group insurance alone. One would expect very low marginal loads and great similarity between Blue Cross and the commercial insurers.

## NOTES

1. Measurement error in the number of insureds would lead to an upward bias in the marginal load. Checking for this by running the regressions in terms of total load, rather than per capita, showed the problem to be nonexistent for Blue Cross insurance and of minor importance for commercial insurance. A more important problem, especially with commercial insurance, is the aggregation of insurance plans with great variation in selling cost, which shows up as load in my data.

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# Lester B. Lave

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The paper I am to discuss is a good one, displaying authoritative knowledge in economic theory, econometrics, and health care. It is entirely appropriate, if somewhat unfortunate, that the author should be gaining firsthand knowledge of the health care industry, as well as of the health care insurance industry, instead of being able to attend the conference. In our conversation yesterday, he assured me that his back problem was actually a way of gaining better information about his subject, for which he is to be commended.

The paper is actually two papers. There is a theoretical paper, which looks at how an individual maximizes utility when confronted with an insurance policy that has a fixed coinsurance rate and a maximum payment. There also is an empirical paper that looks in detail at a 1963 survey and attempts to isolate the price and income elasticities of medical care as well as a host of other effects.

It is unfortunate that the two papers are combined. In order to keep the size of the paper down, Phelps has had to skimp on the details in presenting both his theoretical and empirical results. This means that each of them is extremely difficult to understand. The other problem is that, frankly, the empirical part of the paper has little or nothing to do with the theoretical part. For example, insurance is regarded as compensating an individual for the financial penalty of illness in the theoretical part, yet the financial penalty of illness includes not only the payment to providers for medical care received, but also lost income stemming either from absence from work or permanent disability. Although the income in the model is a permanent income, there are still significant losses owing to disability or death; yet, the insurance that he considers in the empirical part of the paper is merely insurance to compensate for medical expenses. There is no provision for replacing lost income.

I find myself vaguely dissatisfied with the results of the theoretical analysis. Without being able to pinpoint exactly where the dissatisfaction arises, I would point out that some of the conclusions seem to contradict what I see in the world. For example, if an individual is very wealthy, relative to the risk that he bears, he will insure only if the tax incentives of insuring are greater than the loading on the insurance policy or if his individual probability of loss means that the insurance costs less than its actuarial value. Notice that there are in fact tax breaks that attend the purchase of medical insurance and also that almost all medical insurance policies are community rather than individual rated. It seems to me that these facts are central in any theoretical examination of the purchase of insurance.

The purpose of insurance is to mitigate the financial loss stemming from an untoward event. To what extent does an individual (or society) desire to isolate his wealth from the effects of illness? How far is society willing to go to ensure that no individual ever has to live with a correctable health problem? Only the most risk-averse individual would insure so that his wealth was completely independent of untoward events. However, insofar as society bears residual responsibility for mitigating the financial and other consequences of disaster, it might choose to increase the incentives to purchase insurance.

Aside from income redistribution effects, the purpose of health insurance is to remove any financial barrier that might deter an individual from having a correctable medical problem treated. Notice that there are a host of difficulties in determining which health states are "correctable," especially since treatment involves risk. More important, there is no way of rationing medical care so that it goes only to those people it can help. To get 100 people with correctable problems who would not currently seek care to seek it, one would have to lower access and other costs so much that perhaps 5,000 extra people would seek care. Nor is it obvious that iatrogenic disease would not harm more than 100 of the 5,000 people induced to seek care. Although it might be good political rhetoric to declare that no individual shall ever have to live with a correctable health state, it does not make good sense to attempt to implement such a platitude.

Some years ago, Robert Solow and John Kenneth Galbraith argued in the *Public Interest* about how competitive the economy was, the welfare implications of consumer sovereignty, and other matters. Although I thought the discussion was amusing, I was struck the day before yesterday while driving on the Bayshore Freeway by how well Janis Joplin had said it all—"Lord, I need a Mercedes Benz." How much medical care do we *need*? Unlike the Mercedes Benz, we presumably believe that medical care is necessity more than luxury. The theoretical models presented in the papers at this conference assume that medical care is efficacious and a matter for public concern. As Dr. Berg commented, some parts of medical care such as screening and asymptomatic check-ups are of doubtful value; it is equally doubtful that other types of medical care are efficacious on the margin. Would one extra office visit a year improve health? Thus, I wonder if, on the margin, medical care is not more palliation than cure. If so, the Grossman investment model that approaches medical care as restoring the health stock lost to disease (or as improving the health stock) is true on average, but not on the margin. On the margin, I equate Joplin's Mercedes Benz with medical care as being pure consumption. The analogy extends to the unintended harmful effects of each in the form of accidents and iatrogenic disease.

It behooves us to look not only at what people believe they need, but also at whether they are correctly informed. Medical care is perhaps the extreme case in which an individual's observations on the amount of care that he needs are shaped by the providers of that service. It seems to me that we are misleading ourselves and society by happily grinding away at complicated, maximum problems that are based on the assumption that medical care is highly efficacious on the margin. Before we collaborate in pushing expenditures to $200 billion per year, I think we must inquire about whether additional expenditures on medical care would be effective.

I have no quarrel with the empirical contributions of the various papers presented at this conference. Alas, it seems that policymakers are going to go ahead with some national health insurance plan whether medical care is efficacious or not. (My quarrel is with pumping more money into the current system, not with national health insurance *per se*.) The least that we can do for policymakers in these circumstances is to give them some estimates of the increased demand that the lower prices will call forth. Need I add that while we're giving them these positive results, we should continually warn them not to delude themselves that this additional expenditure on medical care is other than a U.S.-built Mercedes Benz.