Which Price is Right: Load or Premium?*

Rexford E. Santerre

Department of Finance, Center for Healthcare & Insurance Studies, School of Business, University of Connecticut, 2100 Hillside Road, Unit 1041, Storrs, CT 06269-1041, U.S.A.

E-mail: rsanterre@business.uconn.edu

This paper uses national time series data for the United States to investigate whether changes in the premium or loading fee offer a better explanation for variations in the percentage of the population with private health insurance from 1960 to 2004. The empirical results suggest that premium provides a better measure of price when estimating the demand for health insurance at the extensive margin. The empirical analysis also indicates that the aggregate short-run price and income elasticities of demand for health insurance are fairly close at -0.19 and 0.27 , respectively. One implication is that the percentage of the population with private health insurance in the United States should continue to decline in the future if real premiums persistently grow significantly faster than the overall economy.

The Geneva Risk and Insurance Review (2008) 33, 90–105. doi:10.1057/grir.2008.10

Keywords: load; premium; demand for health insurance

Introduction

Data from the Bureau of the Census, over the last decade or so, suggest that 15–16 percent of the U.S. population tends to be uninsured for the entire year (DeNavas-Walt *et al.*, 2005). Many health policy analysts believe that an uninsured population causes both inefficiencies and inequities in a health care system. For example, Miller *et al.* (2004) estimate that the uninsured population results in annual costs of about \$65 to \$130 billion largely because of premature deaths and lost productivity. Furthermore, data indicate that being uninsured falls disproportionately on low-income and non-white individuals. Thus, not surprisingly, a relatively large amount of attention has been directed at analyzing the demand for private health insurance as a means of determining the most effective method of reducing the number of uninsured individuals. Specifically, information on price and income elasticity is essential for predicting the responsiveness of demand to subsidized prices or income tax credits.

*I thank Laurie Bates and the anonymous referees of this journal for their comments on earlier versions of this paper.

When estimating demand, researchers have considered two alternative prices for health insurance: the *loading fee* or load and the *premium* payment. Initially, economists treated the loading fee as the appropriate price of health insurance (Sindelar, 1988; Phelps, 1997; Feldstein, 1999; Santerre and Neun, 2007). The loading fee represents the difference between the health insurance premium and the pure premium or expected medical benefits to be paid out. It also captures the amount that health insurers receive to cover claims processing, underwriting, marketing, and risk-bearing costs, as well as taxes and profits.

According to conventional theory, the loading fee reflects that risk-averse individuals are willing to pay an amount above the expected medical benefits paid out because they prefer losing a certain amount of income rather than an uncertain amount of an equal expected magnitude. That is, the load mirrors the degree to which a risk-averse consumer values financial security. According to conventional theory, the size of the load determines whether a person purchases health insurance because, when it rises and all other factors remain constant, the expected utility with health insurance declines relative to the expected utility without health insurance.

Coincidentally, it was the case that health insurers competed more on the basis of controlling the size of the load and not as much on controlling the amount of medical benefits paid out prior to the 1980s. Part of the reason for the emphasis on the size of the load was the mission of the then-dominant non-profit Blue Cross/Blue Shield plans to offer comprehensive and complete health insurance coverage to as many subscribers as possible (Hay and Leahy, 1987; Sindelar, 1988).¹

For several related reasons, however, various health economists have recently begun to regard the total premium payment as the proper price of health insurance instead of the loading fee (Nyman, 2003; Chernew *et al.*, 2005; Gilmer and Kronick, 2005). First, after the 1980s because of a growing concern over rising health care costs, managed care companies started competing on the basis of controlling the size of the load and also “managing” the medical benefits paid out. Since managing the magnitude of medical claims typically entails increasing the administrative portion of the load (e.g., utilization review and disease management activities), the size of the premium payment reflects

¹ This scenario is not exactly correct. Commercial insurers entered the market after the non-profit Blues showed that health insurance was an insurable risk. They entered the market and attracted low-risk employer groups away from the Blues by offering experience rated premiums. Experience rating eventually led to competition based on risk selection not just among commercial insurers but also between commercial insurers and the Blues. Indeed, the risk selection and experience rating proved to be unfavorable to the elderly and led, in part, to the enactment of the Medicare program (MacIntyre, 1965).

how successful managed care companies are at “delivering the goods”. If higher total administration costs are not matched by commensurate medical cost-savings, the resulting higher premium payment may cause some consumers to drop their health insurance coverage.

Second, Nyman (2003) offers an alternative theory regarding why people desire health insurance that is not dependent on risk aversion. According to his theory, health insurance offers access value and is purchased because people desire a wealth transfer, in the event they become ill, from those who remain healthy. Premium represents price in the Nyman model because it captures how much of all other goods and services a consumer knowingly gives up when healthy to have a claim on the wealth of others if she becomes sick. Most people purchase health insurance because the expected utility loss from paying the premium and remaining healthy is far less than the expected utility gain from the wealth transfer when sick. However, high premium payments resulting from monopolistic price distortions in various medical care markets, for example, may create an incentive for some individuals to not purchase health insurance.²

Third, Gilmer and Kronick (2005) take a simpler approach and argue that, because of budgetary pressures, rising premiums mean that consumers must sacrifice increased consumption of other goods and services to remain insured. Hence low- and moderate-income individuals may drop their coverage as health insurance premiums rise. Finally, Chernew *et al.* (2005) point out that rising premiums over time may reflect growing moral hazard costs and increased uncompensated care for the uninsured because of costly new medical technologies. If so, some people may drop their health insurance because they no longer value their coverage as much as they value the goods and services that they would have otherwise given up.

Thus, researchers have defended the use of two different measures of the price of health insurance and studies estimating the demand for health insurance have used either the loading fee (or some proxy variable thereof like size of employer group or marginal tax rate) or premium as the appropriate measure of price. No study to date, however, has been concerned with determining which measure of price is the right one. Use of the wrong price may result in incorrect estimates of the price and income elasticities of demand and thereby lead analysts to draw erroneous policy implications.

This study fills this void in this literature by examining whether the loading fee or premium is a better measure of the price of health insurance. The

² To some extent, medical care can be considered as an input in the production of health insurance. Monopolistic practices in medical care markets increases the price of an input and thereby results in higher insurance prices and less people insured.

analysis is conducted by using national time series data to estimate an aggregate demand for health insurance in the United States over the period from 1960 to 2004.³ The relative superiority of the two measures of price is determined with the *J*-test for non-nested models. In the next section, the data are discussed and the empirical model is specified. The subsequent section involves a discussion of the findings and some exploration of the models. Concluding comments are offered in the last section.

Data and empirical specification

The dependent variable used in the following demand estimation is the percentage of the population in the United States with private health insurance coverage from 1960 to 2004. Systematic estimates of the percent privately insured can be obtained from the Bureau of the Census beginning in 1987. For years prior to 1987, private health insurance enrollment data are gathered from the Health Insurance Association of America (HIAA) and the Centers for Medicare and Medicaid Services (CMS). HIAA began independently gathering and reporting estimates of enrollment in private health insurance plans beginning in 1940 (HIAA, 1999). HIAA's estimates of private health insurance enrollment represent net policy estimates, which attempt to eliminate any double counting among private plans. Double counting occurs when health insurance plans are not mutually exclusive and can result because of double spousal coverage, for example. HIAA (now America's Health Insurance Plans) ended its public reporting of health insurance data in 1995.

CMS (then the Health Care Financing Administration) began estimating the number of people with private health insurance in 1966 for calculations involving the National Health Accounts. Except for a couple of years over the overlapping period of 1966–1986, HIAA reported higher private enrollment figures than CMS. Based upon periodic estimates of health insurance coverage by the Public Health Service prior to 1966, Reed (1967) points out that the HIAA tended to overestimate the private insurance enrollment rate by about 6 percentage points. Thus, to err on the side of conservatism, the lower of the two private insurance enrollment series is used, which means that most of the figures on private enrollment are drawn from CMS for the period from 1966 to 1986.

Combining the HIAA (1960–1966), CMS (1967–1986), and Census (1987–2004) data, Figure 1 provides estimates of the fraction of the U.S. population

³ One advantage of using the aggregate data is that they reflect over 200 million choices made each year over a 45-year period. Another is that the requisite data on the expense load are unavailable at the micro-level so this may be the best one can do.

Figure 1. Fraction of the U.S. population with private health insurance coverage 1960–2004.

with private health insurance coverage for each year over the period from 1960 to 2004. According to the estimates, a little over two-thirds of the population was covered by private health insurance in 1960. Despite the introduction of Medicare and Medicaid in 1966, the rate of private health insurance coverage increased unabated until 1970 when it stood at roughly three-quarters of the U.S. population. After 1970 the private health insurance coverage rate declined but rose again to its highest level of slightly over 76 percent in 1975. Thereafter, the percentage of population with private health insurance declined and hovered around 75 percent over the period from 1975 to 1990. From 1990 to 2004 the private health insurance rate declined at a much greater pace than during the previous period but with some fluctuation. Amazingly, by the end of the 45-year period, the percentage of the U.S. population with private health insurance stood at virtually the same rate in 2004 as it was in 1960.

It is this fluctuation in the fraction of the U.S. population with private health insurance that this study uses to test if the loading fee or premium represents a better measure of price. It is important to keep in mind that the composition of the health insurance product has changed substantially over time. For example, Johnson (1965) reports that only one out of every five persons in the U.S. civilian population had major medical protection in 1962, because most people were covered only by hospital insurance. Today, nearly all insured individuals purchase major medical insurance, which includes coverage for not only hospital stays but also surgeries, outpatient visits, and home care services. For this reason, a control is introduced for the changing health insurance product in the forthcoming multiple regression analysis.

Figure 2. Average real load payment and premium payment per privately insured individual 1960–2004.

For measures of price, data necessary for determining the size of the average loading fee and premium payment are obtained from CMS’s National Health Accounts.⁴ The average load payment for each year is determined by subtracting total medical claims paid out from aggregate premium income and dividing by the number of privately insured individuals. Similarly, the average premium rate is calculated by dividing aggregate premiums by the number of individuals with private health insurance coverage. Both prices are expressed in real terms by dividing by the general consumer price index for each year.⁵

Figure 2 shows how both health insurance prices have trended over time. Over the entire 45-year span, both prices increased but average premium increased at a much more rapid pace than average load. Of course, the rate of growth of the difference between the two, medical claims paid out, accounts for the faster growth of average premium. Nevertheless, both price measures appear to possess sufficient variability over time to capture any effect of a changing price on the quantity demanded of health insurance at the extensive margin.

⁴ These data can be obtained at the Research and Statistics section of the website at CMS, or <http://www.cms.gov>.

⁵ The general rather than medical consumer price index is used in the calculation of real values because we want to know the overall quantity of goods and services that are given up when buying health insurance.

Both conventional theory and Nyman's (2003) model predict that people purchase health insurance when the expected utility with insurance exceeds the expected utility without insurance. It follows from both of these models that out-of-pocket price and consumer income play an important role in the demand for private health insurance. Other variables from these models, such as health status, degree of risk aversion, or magnitude of the expected loss, are difficult to measure or are endogenous, particularly at the aggregate level. Therefore, a bare-boned model, with a few key variables, is used to explain variations over time in the fraction of the U.S. population with private health insurance coverage, Q_t . A basic model seems appropriate given that the objective of this study is to determine the best measure of price and not necessarily to specify a full-fledged model of the demand for private health insurance.

The few key explanatory variables selected to explain health insurance coverage are the two alternative measures of price, P_t , real gross domestic product per capita, GDP_t , as a measure of income, the unemployment rate, U_t , to reflect cyclical uninsurance, and the fraction of medical care costs covered by insurers, $FCOV_t$, to capture the changing health insurance product over time.⁶ Data for the fraction of medical costs reimbursed by health insurers come from the National Health Accounts. Data for GDP and the unemployment rate come from the Bureau of Economic Analysis.

The empirical model is assumed to take the following log-linear form, where u_t is the typical error term:

$$\begin{aligned} \log Q_t = & \beta_0 + \beta_1 \log P_t + \beta_2 \log GDP_t \\ & + \beta_3 \log U_{t-1} + \beta_4 \log FCOV_t + u_t. \end{aligned} \quad (1)$$

Note that the various coefficient estimates can be interpreted as elasticities. Theory and prior research allows us to anticipate the signs of most of these elasticity estimates. The law of demand suggests that the coefficient estimate on the log of price should be found negative ($\beta_1 < 0$). Prior studies indicate that health insurance represents a normal good (e.g. Santerre and Neun, 2007), so a positive parameter estimate is expected on the log of real GDP per capita ($\beta_2 > 0$).

⁶ The poverty rate as a measure of income distribution and various age distribution variables such as percent young, percent middle aged, and percent elderly, the total marginal tax rate, and the percentage of the population in health maintenance organizations were also experimented with as additional explanatory variables. The poverty rate proved to be highly correlated with real GDP per capita and the rest of the variables proved to have no statistical impact on the results.

Because employer-sponsored health insurance remains the dominant form of private health insurance in the United States, the unemployment rate is specified to capture any changes in the degree of cyclical uninsurance over time (Santerre and Neun, 2007). That is, changes in the uninsurance rate that result from swings in the business cycle. Experimentation showed that changes in the unemployment rate influence the percentage of the population with private health insurance with a 1-year lag. The 1-year lag may reflect the length of time it typically takes for COBRA benefits to run out and/or for a family to become eligible for Medicaid coverage. A negative coefficient estimate is expected on the logarithm of the unemployment rate ($\beta_3 < 0$).

Benefit coverage changes when insurers reimburse for a thinner or wider array of medical products including such items as hospital care, outpatient care, home health care, dental care, and prescription drugs. Coverage also changes when insurers sell plans with higher or lower amounts of consumer out-of-pocket payments in the form of deductibles, coinsurance, or co-payments. To capture the extent of the medical benefit coverage that health insurers provide at a particular point in time, FCOV is measured by dividing total private insurance payments to health care providers by total private consumer medical care expenses. When this ratio increases, insurers offer more generous benefits. As a result of any greater coverage, the demand for health insurance at the extensive margin should increase at a given price. Hence a positive parameter estimate is expected on this fraction of insurance coverage variable ($\beta_4 > 0$).

Empirical results

Evidence of serial correlation necessitated first differencing of the time series data. First differencing of the data seems particularly appropriate because diagnostic tests reveal unit roots in several of the time series (e.g., the two price measures and real income per capita). Unit roots can lead to spurious correlations among time series variables (Granger and Newbold, 1974; Phillips, 1986). First differencing of the data offers an added appeal, particularly in this case, because technological changes in medicine, and thus real benefits covered by insurance, are neutralized or netted out as a result of the average year-to-year parameter estimates obtained. These parameter estimates can be interpreted as short-run elasticities.

The first differenced versions of Eq. (1) are estimated by the ordinary least squares (OLS) procedure. Multiple regression results are shown in Table 1 for the two different specifications involving the price of health insurance. Given that the data have been first differenced, the adjusted R^2 s are fairly respectable for both equations. The Breusch–Godfrey and Q -statistic tests fail to detect the

Table 1 Ordinary least squares results: dependent variable is percentage change in the fraction of the population with private health insurance

<i>Variable</i>	<i>Regression results with load as a measure of price</i>		<i>Regression results with premium as a measure of price</i>	
	<i>Estimated coefficient</i>	<i>Absolute value of t-statistic</i>	<i>Estimated coefficient</i>	<i>Absolute value of t-statistic</i>
Constant	−0.004	1.36	0.000005	0.02
Percentage change in price	−0.036	3.18	−0.177	4.53
Percentage change in real GDP per capita	0.243	3.14	0.238	3.56
Percentage change in previous year unemployment rate	−0.028	2.37	−0.021	1.91
Percentage change in fraction of medical costs covered by insurance	0.085	0.84	0.254	2.54
Adjusted R^2		0.35		0.47
DW statistic		1.73		1.62

presence of any remaining first or higher order serial correlation. Despite the different measures of the health insurance price, the two estimated equations yield similar qualitative findings. Both results suggest that the aggregate demand for health insurance is price and income inelastic. In addition, both results imply that health insurance coverage responds inversely to the unemployment rate, as expected. One qualitative difference between the two sets of results is that the generosity of health insurance coverage only influences the demand for health insurance in the specification involving premium as the measure of price.

However, from a quantitative perspective, the two specifications do yield a dramatically different finding. Specifically, the specification involving the loading fee as a measure of price suggests that price elasticity of demand equals -0.036 whereas the specification with premium as a measure of price indicates that price elasticity of demand equals -0.177 . These two estimates are statistically different from one another ($t = 5.64$) and the magnitude of the difference is not trivial when evaluated within a policy context. For example, the results imply that the fraction of the population with private health insurance would grow 1.4 percent faster if price declined by 10 percent and premium is the correct price rather than the loading fee.

Recall that three separate data series for the percentage privately insured have been spliced together. The combination of the series could affect the reliability of the findings. As a result, the two demand equations are also estimated for the shorter periods 1967–2004 and 1987–2004. The resulting

estimated price elasticities are found to be very robust with respect to the specific time period examined, equaling -0.032 and -0.031 for the load demand model and -0.183 and -0.178 for the premium demand model. The estimated income elasticities are more varied but fall within a similar range. Specifically, the estimated income elasticities are 0.20 and 0.32 for the load demand equation and 0.21 and 0.25 for the premium demand equation. Consequently, it does not appear that the three data series are wildly inconsistent with one another.⁷

Another concern is that the negative coefficients on both price variables simply reflect that both the dependent variable and each insurance price are constructed with a similar measure – the number of people with private insurance. As a result, the inverse relationship is almost guaranteed given that this number shows up in the numerator of the quantity variable and the denominator of the price variable. To examine this possibility, real aggregate load and real aggregate premiums are substituted in the place of the per unit prices in the respective equations. The estimated coefficients on the aggregate price terms are found to be -0.031 and -0.116 for the load and premium demand equations suggesting that the regression results do not merely reflect the manner in which the variables are constructed.⁸

Unfortunately, the possibility of endogeneity bias raises a question about the validity of this comparison of the two OLS results regarding the price elasticity of demand. That is, one or both of the two measures of health insurance price and the fraction of the population privately insured may be jointly influenced by some omitted variables and/or reverse causality may play a role. For example, people facing above average health risks may represent the marginal purchasers of insurance and thereby raise insurance price for everyone as the size of the insured group increases. In addition, the fraction of medical costs reimbursed by health insurers may be endogenous because this variable, as a measure of the generosity of insurance, could be interpreted as the quantity demanded of health insurance at the intensive margin. If these variables are endogenous, the OLS estimates are biased and inconsistent.

As a result, two stage least square estimates (2SLS) are derived. To implement the 2SLS procedure, a set of instrumental variables must be found that are correlated with the insurance prices or FCOV but not with the error term of the health insurance demand function in Eq. (1). Within a time series framework, it is generally agreed that lagged values of the suspected endogenous variable serve as good instruments if the error terms are uncorrelated (Maddala, 1992; Kennedy, 2003). As discussed previously, serial

⁷ These regression results will be provided to the interested reader.

⁸ These regression results will also be provided to the interested reader.

correlation can be ruled out after first differencing of the data. Therefore, lagged first differenced values of the two prices and FCOV serve as instruments in the first stage of the estimation procedure.

Another instrument, the logarithm of the annual number of new drug approvals, is specified in the FCOV first-stage equations. According to the new drug offset effect theory of Lichtenberg (2007) and Zhang and Soumerai (2007), among others, a new drug tends to pay for itself at the margin by reducing overall health care spending. Simply put, new drugs are argued to keep people out of expensive hospitals and from seeing costly specialists. If this theory holds as the empirical work seems to suggest, then insurer reimbursements on hospital and outpatient expenses should decline with a greater number of new drugs. Privately insured individuals typically purchase relatively high rates of coverage for these two products given their “big ticket” nature. However, fewer privately insured typically purchase coverage for prescription drugs and/or higher co-payments are normally involved. Thus, the insurer proportion of medical expenses declines with a greater number of new drugs because of variations in coverage rates across medical products. For that reason a negative coefficient estimate is expected on the logarithm of the number of new drugs in the FCOV first-stage equations.⁹

Table 2 provides the first-stage results where all of the exogenous factors in the system have been regressed on the suspected endogenous variables for each of the two health insurance demand models. A Wald *F*-test for the significance of each set of instruments is reported at the bottom of the table for each estimation equation. The appropriateness of using the specified instruments rests on the plausibility of the signs and the statistical significance of the coefficient estimates on the selected instruments. With this in mind, notice that the coefficient estimates on the two lagged health insurance price terms are statistically different from zero at conventional levels of significance in both first-stage price models.¹⁰ It is very likely that current period price growth is directly related to prior year price growth (Gabel *et al.*, 1991). It is also interesting to note that the opposite signed coefficients on the lagged price

⁹ The specification of previous year new drug approvals assumes it takes 1 year for a new drug to launch. The new drug variable is not first differenced for theoretical and empirical reasons. Theoretically, the number of new drugs represents the change in pharmaceutical knowledge from one year to the next so specifying its level rather than its first difference is consistent with the other variables being in difference form. Empirically, the number of new drugs does not trend upward like most of the other variables but instead cycles over time. Data for the number of new drug approvals come from the Federal Food and Drug Administration.

¹⁰ Experimentation shows that the coefficient estimates on any further lagged values of price are statistically insignificant.

Table 2 First-stage regression results

<i>Independent variables^a</i>	<i>Regression results with load as a measure of price</i>		<i>Regression results with premium as a measure of price</i>	
	<i>Endogenous variables</i>		<i>Endogenous variables</i>	
	<i>Percentage change in real load</i>	<i>Percentage change in FCOV</i>	<i>Percentage change in real premium</i>	<i>Percentage change in FCOV</i>
Constant	-0.282 (2.38)	0.044 (3.09)	0.006 (0.19)	0.042 (2.87)
Percentage change in real GDP per capita	2.15 (2.80)	-0.266 (2.86)	0.148 (0.69)	-0.243 (2.67)
Percentage change in previous year unemployment	0.224 (1.83)	0.001 (0.08)	0.089 (2.59)	0.006 (0.44)
Logarithm of prior year new drug approvals	0.074 (2.05)	-0.009 (2.15)	-0.008 (0.77)	-0.009 (1.98)
Percentage change in fraction of medical costs covered by insurance in the previous year	3.81 (3.46)	0.455 (3.42)	0.84 (2.46)	0.504 (3.47)
Percentage change in price during the previous year	0.384 (3.74)	-0.009 (0.70)	4.90 (3.80)	-0.023 (0.42)
Percentage change in price 2 years ago	-0.365 (3.25)	-0.014 (1.07)	-0.425 (3.63)	-0.016 (0.33)
Adjusted R^2	0.67	0.46	0.60	0.43
DW Statistic	2.26	2.03	1.98	2.07
Wald F -test for instruments	8.91	10.78	9.1	10.42
	(prob. = 0.0007)		(prob. = 0.0002)	
	(prob. = 0.0007)		(prob. = 0.0003)	

^aEstimated coefficient with the absolute value of the t -statistic in parentheses.

terms produce a cycling of premium growth which analysts have dubbed the underwriting or profitability cycle (Gabel *et al.*, 1991; Grossman and Ginsburg, 2004).

The first-stage regression results for the FCOV equations perform equally well. The coefficient on the lagged FCOV is positive and statistically significant.¹¹ Moreover, the estimated coefficient on the logarithm of new drugs is negative and different from zero at conventional levels of statistical significance. Finally, the selected instruments jointly provide a significant amount of explanation for each endogenous variable as indicated by the

¹¹ Further lags for FCOV offer no additional information.

Table 3 Two-stage least squares results: dependent variable is percentage change in the fraction of population with private health insurance

<i>Variable</i>	<i>Regression results with load as a measure of price</i>		<i>Regression results with premium as a measure of price</i>	
	<i>Estimated coefficient</i>	<i>Absolute value of t-statistic</i>	<i>Estimated coefficient</i>	<i>Absolute value of t-statistic</i>
Constant	-0.010	1.89	-0.002	0.53
Percentage change in price	-0.050	2.69	-0.192	3.20
Percentage change in real GDP per capita	0.341	2.92	0.271	3.10
Percentage change in previous year unemployment rate	-0.028	2.08	-0.022	1.75
Percentage change in fraction of medical costs covered by insurance	0.297	1.60	0.377	2.13
Adjusted R^2		0.23		0.43
DW statistic		1.97		1.72

relatively high F -statistics. All in all, these first-stage results attest to the plausibility of the selected instruments for the 2SLS analysis.¹²

The results of the 2SLS estimation are depicted in Table 3. The 2SLS procedure generally yields coefficient estimates that are larger in value than their OLS counterparts. Different from the OLS results, the coefficient estimate on FCOV in the load demand equation is now statistically different from zero at better than the 10 percent level for a one-tailed test. Most importantly for the purpose at hand, the health insurance price elasticities differ across the two demand models and thus from a public policy perspective it is important to know which price is superior to the other.

To examine which price is better at explaining the percentage of the population with private health insurance, the J -test is used. The basic idea behind the J -test is that one of two non-nested models is the correct one if the fitted value from model #1 offers no explanatory power when estimating model #2 but the predicted value from model #2 provides some explanation in model

¹² Following the approach of Davidson and MacKinnon (1993), a Hausman (1978) test is conducted to determine if the null hypothesis of exogeneity could be rejected for the two prices and FCOV. The null hypothesis could be rejected for FCOV in both cases and the load measure of price. However, the null hypothesis of exogeneity for the premium price could not be rejected. However, the premium price is treated as being endogenous in the 2SLS procedure given the strong theoretical likelihood that it cannot be exogenous.

#1. For example, the loading fee would represent the right price if its predicted value offers some additional explanation for the percentage of the fraction of the population with private health insurance in the premium demand model but the opposite scenario does not hold.

Within a 2SLS context it is computationally easier to use a procedure outlined by Maddala (1992), which he demonstrates is similar to a *J*-test. In terms of this study, the relevant test equation is:

$$d(\log Q_t)^A - d(\log Q_t)^{1P} = \pi[d(\log Q_t)^{2P} - d(\log Q_t)^{1P}]. \quad (2)$$

The expression on the extreme left of Eq. (2) represents the actual percentage change in the fraction of the population with private health insurance at time *t*, whereas the expression to the immediate left of the equality sign is the predicted percentage change from either one of the demand models, say the load demand model. The expression on the left within the brackets on the right-hand side of the equation is the predicted percentage change in the fraction of population with private health insurance based upon the second price, for example, the premium demand model. If the equation is estimated and π is not significantly different from zero, then the premium demand model does not add any information to explaining the actual percentage change over and above what the load demand model offers. But if π is significantly different from zero, the premium demand model provides information on changes in the actual fraction with private health insurance over and above the load demand model. The process is performed for both scenarios. But we cannot rule out that π may not differ from zero or that π may be different from zero under both scenarios. Consequently, the test can be inconclusive.

The regression results associated with this version of the *J*-test finds that π equals 0.94 with a *t*-statistic of 3.82 and the adjusted R^2 for the estimation equation equals 0.26 when the load demand model is treated as the null hypothesis and the premium demand model is treated as the alternative hypothesis. On the other hand, π equals 0.06 with a *t*-statistic of 0.23 and the adjusted R^2 for the estimation equation equals 0.001 when the situation is reversed. Thus, the results can be interpreted to mean that the load demand model offers no statistically significant explanation of the fraction of the population privately insured above and beyond that of the premium demand model. However, the predicted value from the premium demand model provides additional explanation for the fraction of population with private health insurance beyond that of the load demand model. Thus, it can be inferred from the *J*-test that the health insurance premium represents a better measure of the price of health insurance than the load.

Conclusion

Finding effective ways of providing health insurance coverage to nearly 16 percent of the population remains one of the most critical social issues facing the United States today. Previous studies estimating the price and income elasticities of demand for health insurance typically have had that goal in mind. However, these studies have been uncertain about whether the loading fee or premium represents the right price, which has raised some concern about the validity of the findings. This study addresses this concern by finding empirically that premium represents a better measure of the price of health insurance using aggregate data for the United States over the period from 1960 to 2004.

The Lewin Group predicts that 300,000 to 400,000 workers in employer-sponsored plans would lose their private health insurance coverage if premiums increase by 1 percent (GAO, 1998). For comparative purposes, our -0.19 estimate of the premium price elasticity along with sample data on the privately insured for 1998 can be used to predict how many people would drop their private health insurance coverage given a 1 percentage increase in premiums. The simulation suggests that a 1 percent increase in the average premium translates into 360,000 fewer people with private health insurance coverage on an annual basis. This estimate falls within the range of the Lewin Group but also includes people who previously purchased health insurance on an individual basis (e.g., the self-employed).¹³

The aggregate short-run income elasticity of demand is estimated to be roughly 0.27. Given that the premium and income elasticities are fairly close in magnitude, the implication is that the percentage of the population with private health insurance increases as long as average premium grows at a slower rate than income grows, assuming all other factors remain constant. If the immediate future resembles the recent past, this scenario does not bode well for those holding the view that the market should be able to insure most of the currently uninsured. Specifically, real premium growth averaged 6.7 percent over the last 5 years. At the same time, a market economy tends to grow by 3 percent per year over the long run. Given the estimated price and income elasticities, that means the percentage of the population with private health insurance will decline by 0.46 percent each year, *ceteris paribus*, given these health insurance premium and income growth rates. It also means that the U.S. health care system must be able to seriously slow the growth of premiums on a sustained basis to less than 4 percent or it may be time to replace its primarily voluntary private health insurance system with an alternative arrangement.

¹³ Chernew *et al.* (2005) and Gilmer and Kronick (2005) estimate that a 1 percent increase in premiums would lead to 164,000 and 246,000 fewer people with health insurance, respectively, after considering public health insurance alternatives.

References

- Chernew, M., Cutler, D.M. and Keenan, P.S. (2005) 'Increasing health insurance costs and the decline in insurance coverage', *Health Services Research* **40**: 1021–1039.
- Davidson, R. and MacKinnon, J.G. (1993) *Estimation and Inference in Econometrics*, New York: Oxford University Press.
- DeNavas-Walt, C., Proctor, B.D. and Lee, C.H. (2005) *Income, poverty, and health insurance coverage in the United States: 2004*, U.S. Census Bureau, Current Population Reports, P60-229, U.S. G.P.O., Washington, DC.
- Feldstein, P.J. (1999) *Health Care Economics*, Albany, NY: Delmar Publishers.
- Gabel, J., Formisano, R., Lohr, B. and Di Carlo, S. (1991) 'Tracing the cycle of health insurance', *Health Affairs* **10** (winter): 49–61.
- GAO (1998) 'Private health insurance: Impact of premium increases on the number of covered individuals is uncertain', GAO/HEHS-98-203R.
- Gilmer, T. and Kronick, R. (2005) 'It's the premiums, stupid: Projections of the uninsured through 2013', *Health Affairs* Web Exclusive, 5 April, pp. w5-143–w5-151.
- Granger, C. and Newbold, P. (1974) 'Spurious regressions in econometrics', *Journal of Econometrics* **2**: 111–120.
- Grossman, J.M. and Ginsburg, P.B. (2004) 'As the health insurance underwriting cycle turns: What next?', *Health Affairs* **23**: 90–102.
- Hausman, J.A. (1978) 'Specification tests in econometrics', *Econometrica* **46**: 1251–1271.
- Hay, J.W. and Leahy, M.J. (1987) 'Competition among health plans: some preliminary evidence', *Southern Economic Journal* **50**: 831–846.
- Health Insurance Association of America (HIAA) (1999) *Source Book of Health Insurance Data*, Washington, DC: HIAA.
- Johnson, H.M. (1965) 'Major medical expense insurance', *Journal of Risk and Insurance* **32**: 211–236.
- Kennedy, P. (2003) *A Guide to Econometrics*, Cambridge, MA: The MIT Press.
- Lichtenberg, F. (2007) 'Effects of new drugs on overall health spending: Frank Lichtenberg responds', *Health Affairs* **26**(3): 887–890.
- MacIntyre, D.M. (1965) 'Rate philosophy and health insurance competition', *Journal of Risk and Insurance* **32**: 525–537.
- Maddala, G.S. (1992) *Introduction to Econometrics*, New York: MacMillan Publishing Company.
- Miller, W., Vigdor, E.R. and Manning, W.G. (2004) 'Covering the uninsured: what is it worth?', *Health Affairs* Web Exclusive, 31 March, pp. 157–167.
- Nyman, J.A. (2003) *The Theory of Demand for Health Insurance*, Stanford, CA: Stanford University Press.
- Phelps, C.E. (1997) *Health Economics*, Reading, MA: Addison-Wesley.
- Phillips, P.C. (1986) 'Understanding spurious relations in econometrics', *Journal of Econometrics* **33**: 311–340.
- Reed, L.S. (1967) 'Private health insurance: Coverage and financial experience, 1940–1966', *Social Security Bulletin* **30**: 3–22.
- Santerre, R.E. and Neun, S.P. (2007) *Health Economics: Theories, Insights, and Industry Studies*, Mason, OH: South-Western Publishers.
- Sindelar, J.L. (1988) 'The declining price of health insurance', in H.E. Frech III (ed.) *Health Care in America*, San Francisco: Pacific Institute for Public Policy, pp. 259–291.
- Zhang, Y. and Soumerai, S.B. (2007) 'Do newer prescription drugs pay for themselves? A reassessment of the evidence', *Health Affairs* **26**(3): 880–886.