Economics of self-medication: theory and evidence

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Summary

A pervasive phenomenon in developing countries is that self-prescribed medications are purchased from drug vendors without professional supervision. In this article we develop a model of self-medicating behavior of a utility-maximizing consumer who balances the benefits and risks of self-medication. The empirical investigation focuses on the role of income and health insurance on the use of self-medication. Our data are from the World Bank’s Living Standards Measurement Survey of Vietnam, 1997–1998. The results show that self-medication is an inferior good at high income levels and a normal good at low income levels, and it shows a strong and robust negative insurance effect. Copyright © 2003 John Wiley & Sons, Ltd.

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Introduction

Self-medication is a pervasive phenomenon in most economies. The World Health Organization (WHO) reports that sale of self-prescription products, i.e. purchase of prescription-only drugs without a prescription, is far more common than the sale of over-the-counter (OTC) drugs. In many less developed economies where both public and private health care delivery systems are relatively basic, self-medication is often the single most important form of health care. See, for example, Gertler and Litvack [2] and the World Bank’s Vietnam Health Sector Review [3]. With competitive promotion of self-medication products greatly aided by the Internet, and deregulation of OTC sales of pharmaceutical products with active ingredients, the scope of self-medication is expected to grow in the future.

Self-medication is often not a preferred alternative. However, it is practised as professional care may be unavailable or prohibitively expensive. The risks stem from: incorrect self-diagnosis and self-care; absence of specialized knowledge of alternative treatments; abuse of medications due to incorrect dosage and duration of use; dangers of side-effects and neglect of interactions with other drugs. Thus it seems likely that self-care is risky and of lower average quality than professional care.

A well-documented phenomenon is that many drugs that in the United States are available only by prescription, e.g. penicillin, are available as OTC drugs in less developed countries. In so doing the consumer can have the benefit of such drugs at low cost and, at the same time expose themselves to the aforementioned risks. Moreover, the dangers of self-medication abuse may go beyond an individual. For example, there may be an increase in drug-induced and/or drug-resistant strains of diseases as well as greater social acceptance of reliance on quick-fix remedies. See [3, pp. 119–120]. Others who have commented on the harmful aspects of self-medication include Hughes et al. [4],
Levy [5], Mudur [6] and Neu [7]. In many cases the extent of self-medication is so great that it raises the concern of international institutions such as the WHO [8] and the World Bank, who entertain the possibility of greater regulation of the pharmaceutical drugs and alternative forms of intervention in the drug market.

Despite its evident importance the topic of self-medication has not been seriously studied from either the viewpoint of theoretical or empirical economics. In this article we take a first step by developing a model of self-medicating behavior of a utility-maximizing consumer who balances the benefits and risks of self-medication. Then we provide an empirical study of the use of self-medication, for which our working definition is the use of self-prescribed medications purchased from drug vendors without professional supervision. We focus primarily on the role of income and health insurance on the use of self-medication. Our data are from Vietnam, and are derived from the World Bank’s Living Standards Measurement Survey of Vietnam 1997–1998. Trivedi [9] provides a more detailed empirical analysis of the pattern of health care utilization in Vietnam, but in this study we concentrate solely on economics of self-medication.

The remainder of this article is organized thus. In the next section, we develop our theoretical model. The third section provides an analysis of its implications. The fourth section provides some background information and data that both instigated and shaped our study and describes the Vietnamese data that are used in our empirical analysis. The fifth section outlines the key econometric issues that condition the empirical analysis that is reported in sixth section. In the last section we conclude.

The model

Let the utility function of a typical consumer be of the form $U(c, h)$, where $c$ is the consumption and $h$ is the health status. As usual, $U(c, h)$ is strictly increasing and strictly concave in $c$ and $h$. A consumer is endowed with income $y$ and initial health $h_0$. There are two ways to improve one’s health status: professional care or self-medication. The former is obtained by visiting a doctor’s office (or a hospital), denoted by $V$, while the latter is obtained by purchasing drugs at a pharmacy, denoted by $Q$. The consumer may choose self-medication because it is cheaper, even though it may be less effective and generally risky. To contrast the risky nature of self-medication and to make the model tractable, we assume that professional care is risk-free, i.e. deterministic. Then the consumer’s problem is a portfolio selection problem.

Given income $y$, the consumer allocates $c$ dollars on consumption, $V$ dollars on professional care and $pQ$ dollars on self-medication, where $p$ is the price of $Q$, so that the budget equation is

\[ y = c + V + pQ \]  

By setting the price of $V$ as unity, the parameter $p$ is also the relative price of self-medication $Q$ to professional care $V$. Let the return to professional care be $R > 0$ and the return to self-medication be $R - e$, where $e$ is a random variable with $E[e] > 0$. As such, self-medication is, on the average, less efficient than professional care. Furthermore, we allow for the possibility that $R - e < 0$ for some $e$, if wrong medicines were used that caused negative side effects. With the purchase of $V$ and $Q$, the actual level of health is

\[ h = h_0 + RV + (R - e)Q \]

which is a random variable. Then the consumer’s problem is

\[ \max_{V, Q} E[U(y - V - pQ, h_0 + RV + (R - e)Q)] \]  

The model can be generalized to include other health services.\(^b\)

The comparative statics of a bivariate decision problem such as (2) are generally ambiguous. To make the model tractable and to stress the risky nature of self-medication, we give problem (2) a mean–variance formulation. Let $\mu = E[R - e]$ be the expected return to the risky asset, and let $\sigma^2 = E[(R - e - \mu)^2]$ be the variance of $R - e$. Assume $U(c, h) = u(c) + v(h)$ and $v(h)$ is of the form

\[ v(h) = -(b - h)^2/2, \quad \text{with } 0 \leq h \leq b \]  

Then Problem (2) becomes

\[ \max_{V, Q} \{u(y - V - pQ) - (b - h_0 - RV - \mu Q)^2/2 - \sigma^2 Q^2/2\} \]  

The first order conditions for (4) are

\[ -u'(y - V - pQ) + (b - h_0 - RV - \mu Q)R = 0 \]  

and

\[ -pu'(y - V - pQ) + (b - h_0 - RV - \mu Q)\mu - \sigma^2 Q = 0 \]
In this case, the Hessian matrix is

\[
H = \begin{bmatrix}
h_{11} & h_{12} \\
h_{21} & h_{22}
\end{bmatrix} = \begin{bmatrix}
u''(y - V - pQ) - R^2 & pu''(y - V - pQ) - R\mu \\
pu''(y - V - pQ) - R\mu & p^3u''(y - V - pQ) - \mu^2 - \sigma^2
\end{bmatrix}
\]

It is straightforward to verify that \( h_{ij} < 0 \) for all \( i \) and \( j \), and the determinant of \( H \)

\[
D = [(\mu - pR)^2 + \sigma^2](-u'') + R^2\sigma^2 > 0
\]

Some comments on the modeling approach are in order. We believe that treating self-medication as a risky asset distinguishes it from many other self-help activities, which are typically time-intensive, often of low quality and sometimes providing individuals with a sense of satisfaction. For time-intensive self-help activities, a model of allocation of time (e.g. Becker [10]) seems more appropriate. Even though time may play a factor in health care, it is usually secondary rather than a dominant factor. For those self-help activities characterized by a lower quality work at a lower price, a model of hedonic price index (e.g. Rosen [11]) will be more suitable. For those ‘do-it-yourself’ activities that generate psychic income, a model in which this activity directly enters the utility function would be more appropriate.

**Theory implications**

Before we start, we need a key lemma from which many of the theory implications follow. From (5) and (6), we have

\[
b - h_0 - RV - \mu Q \equiv \sigma^2 Q
\]

Since \( b > h \) for any realization of \( \varepsilon \), we must have \( b > E[h] = h_0 + RV + \mu Q \). Therefore, \( \mu > pR \).

1. First, the income effect is derived from

\[
\left[ \frac{\partial V}{\partial y} \right] = \frac{1}{D} \begin{bmatrix} h_{22}u'' - h_{12}pu'' \\
h_{11}pu'' - h_{21}u''
\end{bmatrix}
\]

\[
\left[ \frac{\partial Q}{\partial y} \right] = \frac{1}{D} \begin{bmatrix} u''(-\mu^2 - \sigma^2 + p\mu R) \\
u'' R(\mu - pR)
\end{bmatrix}
\]

Since \( \mu > pR \), we have \( \partial V / \partial y > 0 \) and \( \partial Q / \partial y < 0 \). That is, \( V \) is a normal good and \( Q \) is an inferior good. Our empirical analysis provides support for this contention. Our theory predicts that, other things being equal, the practice of self-medication would diminish as the economy grows. Similarly, our theory predicts that the practice of self-medication will be more prevalent in poor countries than in rich countries.

2. The price effect is given by

\[
\left[ \frac{\partial V}{\partial p} \right] = \frac{1}{D} \begin{bmatrix} -h_{22}Qu'' - h_{12}(u' - pQu'') \\
h_{11}(u' - pQu'') + h_{21}Qu''
\end{bmatrix}
\]

\[
\left[ \frac{\partial Q}{\partial p} \right] = \frac{1}{D} \begin{bmatrix} -h_{12}u' \\
h_{11}u'
\end{bmatrix} - Q \begin{bmatrix} \frac{\partial V}{\partial y} \\
\frac{\partial Q}{\partial y}
\end{bmatrix}
\]

The first term is the substitution effect, which is negative. The second term is the income effect. Since the expenditure share of self-medication is small, the income effect could also be small. If so, the substitution effect dominates the income effect and the law of demand for self-medication holds, i.e. an increase in \( p \) lowers the demand for \( Q \). This implication is particularly relevant to our analysis of Vietnam health utilization data where the expenditure share is small.

There are three immediate implications from the law of demand. First, an increase in health insurance coverage lowers the price of professional care and is equivalent to an increase in the relative price \( p \). Then the theory predicts that better health insurance would discourage self-medication. We test this prediction in our empirical studies. Second, the ‘full cost’ of professional care should include the time cost of traveling. Other things being equal, rural patients face a lower relative price of self-medication than urban patients because of accessibility. This helps explain the fact that self-medication activities are more prevalent in rural areas than in urban areas. Third, the deregulation of OTC sales and increased availability of self-medication products amount to a substantial reduction in the price of self-medication. It provides an explanation for the growth of self-medication in Vietnam.

3. The effects of increased uncertainty are given by

\[
\frac{\partial V}{\partial \sigma^2} = \frac{Qh_{12}}{D} > 0 \quad \text{and} \quad \frac{\partial Q}{\partial \sigma^2} = \frac{Qh_{11}}{D} < 0
\]

The former says that a riskier self-medication would increase the demand for professional care.
This result is intuitively appealing because professional care and self-medication, by the functional form of (4), are substitutes in utility. The latter says that the demand for self-medication is lowered if self-medication becomes riskier. It also suggests that making pharmaceutical products safer for home use would encourage more self-medication.

4. The effects of initial health status are given by

\[
\begin{align*}
\frac{\partial V}{\partial h_0} &= \frac{1}{D} \left[ h_{22} R - h_{12} \mu \right], \\
\frac{\partial Q}{\partial h_0} &= \frac{1}{D} \left[ -p(\mu - pR)u'' - R\sigma^2 \right].
\end{align*}
\]

Clearly, \(\frac{\partial Q}{\partial h_0} < 0\), but the sign of \(\frac{\partial V}{\partial h_0}\) remains indeterminate. It simply says that the healthy tend to spend less on self-medication. However, no conclusion can be drawn about professional care.

5. To facilitate the analysis of corner solutions and to study the total medical expenditure, \(V + pQ\), we shall reformulate Problem (4) as a constrained maximization problem. Let

\[
F(c, V, Q) = u(c) - (b - h_0 - RV - \mu Q)^2/2 - Q^2\sigma^2/2
\]

Assume \((b - h_0 - RV - \mu Q) > \sigma^2 Q\) for all \(V\) and \(Q\) to ensure \(F_Q > 0\). Then \(F(c, V, Q)\) is strictly increasing and strictly concave in \((c, V, Q)\). Then, Problem (4) becomes

\[
\phi(y, h_0, p) = \max_{c, V, Q} F(c, V, Q), \text{ s.t. } (1)
\]

(7)

where \(\phi(y, h_0, p)\) is the indirect utility function.

We show in Appendix, paragraph 1, that the law of diminishing marginal utility of income holds, i.e. \(\frac{\partial \lambda}{\partial y} < 0\), where \(\lambda\) is the marginal utility of income, or the Lagrange multiplier of Problem (7). Then, from \(F_c = u'(c) = \lambda\), an increase in \(y\) lowers \(\lambda\) and hence increases \(c\), i.e. consumption is a normal good. The income effect on the total medical expenditure \(V + pQ = y - c\) is positive if the marginal propensity to consume is less than unity. There is some evidence to support this; e.g. in Trivedi [9], the income elasticity of medical expenditure is found to be around 0.6.

We show in Appendix, paragraph 2, that income and initial health are substitutes in utility in the sense that the marginal utility of income falls as the initial health status rises, i.e.

\[
\frac{\partial^2 \phi}{\partial y \partial h_0} - \frac{\partial \lambda}{\partial h_0} < 0
\]

A corollary follows immediately from this substitutability. From the first order condition, \(u'(c) = \lambda\), an increase in \(h_0\) lowers \(\lambda\) and hence increases \(c\), which in turn reduces \(V + pQ = y - c\), since \(y\) remains fixed. Together they imply that the healthier is the person, the smaller is the total medical expenditure, and the greater is the consumption.

6. Finally, we examine the nature of corner solutions, because there is strong evidence of such solutions in the data. We show in Appendix, paragraph 3, that, for any consumer who chooses zero self-medication, there is a critical price \(p_0\) above which the demand curve for self-medication is vertical, below which the demand has a finite slope. Note that some consumers may not choose \(V = 0\) for all \(p\). As \(p\) rises, some people who had positive quantity of self-medication may begin to choose zero self-medication. The theory thus predicts that the aggregate frequency of zero self-medication would not fall as \(p\) rises. This is consistent with the law of demand.

The income effect of corner solutions will be analyzed through studying the income expansion path. We show in Appendix, paragraph 4, that, if the income expansion path is continuous, then the expansion path stays on the \((c, V)\)-plane once it reaches there. In other words, once a consumer has chosen zero self-medication, she will stay with her choice if her income is increased. However, some may start to choose zero self-medication with increased income. Hence, the aggregate frequency of zero self-medication would not fall as income rises. This is consistent with self-medication being an inferior good.

Data and some issues

Our empirical analysis is based on the individual and household level data from the 1997–1998 Vietnam Living Standards Survey (VLSS). We focus on the major features and determinants of self-prescribed use of pharmaceutical drugs. The term ‘self-medication’ in the empirical section of this paper is used to indicate any pharmacy visit without a prior contact with another health care provider. This may include genuine prescription drugs (in the USA) that are OTC drugs in
Vietnam, as well as OTC drugs. The riskiness varies by what is under consideration. The major foci are price and income effects, but we control for the standard socio-demographic factors that are expected to be relevant. There are some important features and limitations of the VLSS data that require advanced econometric tools that we also employ in our analysis.

The VLSS data are the subject of a comprehensive but largely descriptive analysis in the World Bank’s *Vietnam Health Sector Review* [2] that has provided a great deal of statistical information on the structure and organization of health care delivery and on the broad pattern of health care use in Vietnam. Moreover, it provides many points of comparisons between the 1992–1993 VLSS and its 1997–1998 counterpart. However, the study includes no explicit or formal modeling of the data and relationships between variables are studied or interpreted on a (potentially misleading) bivariate basis. By contrast, this article will address the task of measuring and testing various hypotheses about health care utilization within a multivariate econometric framework.

**Background**

It is useful to begin with a summary of some important features of health care utilization in Vietnam.

Several studies (in English) have examined various aspects of the health care system in Vietnam, including Gertler *et al.* [12], Gertler and Litvack [2], World Bank [3, 13], and Trivedi [9]. We now review the institutional background and some descriptive features of our self-medication data. The reported summaries are based on the health section of the 1997–1998 VLSS data provided by the World Bank.

Apart from self-medication care can be obtained through the following providers: private health facilities; outpatient services in government hospitals; commune health centers; regional polyclinics and other government health facilities; traditional medical practitioners. The data show that the last two are a small part of the total number of visits. Drug vendors, government hospitals, private health facilities and commune health centers account for bulk of the total usage. There are two main types of private health providers: (i) full time providers who own private facilities, and (ii) part-time providers who are employed by the public health facilities but engage in private practice during off-hours [3, p. 101]. Both licensed and unlicensed practices are included in this category. Nearly 70% are estimated to be in urban areas.

The broad pattern of health care usage in Vietnam can be summarized by the observation that the relatively higher income groups get treated in government hospitals, the poorest groups in commune health centers, and all groups practice self-medication to a high degree. However, as noted below, those who are covered by Vietnam Health Insurance (VHI) program are serviced by government hospitals. And insurance coverage under VHI is more extensive for the relatively better off groups.

All types of utilization have grown since 1993, but the use of drug vendors has shown a particularly fast rate of growth – it has more than tripled. This attests to the overwhelming importance of self-medication. The World Bank points out that the increase in average number of pharmacy visits is accompanied by a decline in out of pocket expenditure on drugs [3, p. 56]. Previous analyses have attributed this result to two factors. First, the improvement in the supply and availability of pharmaceuticals and drugs between 1993 and 1998, following the deregulation of the retail markets and liberalization of the pharmaceutical industry in 1989. Evidence suggests that the real price of drugs may have declined over the 1993–1998 period. The second factor is the ease of access of medicines relative to the alternatives. In rural area distance from government health facilities and poor quality of health services at commune health centers have been cited as possible reasons for the continued growth of self-medication.

Table 1 provides figures on annualized health service contact rates for 1993 and 1998 by type of provider. The contact rate is highest for pharmacies and drug vendors. Moreover, it has grown more than threefold in the 5 year period. Average contact rates with health care providers can be misleading indicators of accessibility if not supplemented with other information. In any sample, one is likely to observe zero contact frequency, in part because the individual was healthy in the survey period and did not need health care. It is, therefore, an interesting preliminary calculation to compare the overall contact rates with those who were classified as sick or injured, and to further compare the same across different insurance status.
The results of such a calculation are shown in Table 2.

Table 3 shows that commune health care is sought at a 3–4 times higher rate by those in the sick in the lowest 25% income class compared with the sick in the highest 25% income class. The situation is reversed in the case of government hospitals. The contact rate in that case is 3–4 times higher for the sick rich than for the sick poor, indicating that these hospitals are more important providers of health care for the relatively better off. This differential usage is also in evidence for private health providers, but the difference multiple is closer to 2 than 3 or 4. The differential usage is the smallest for pharmaceutical providers. This appears to indicate that access to private drug vendors is roughly equal for the sick, whether low- or high-income. However, this needs a caveat because government hospitals that are favored by the high-income groups also dispense pharmaceutical drugs.

Table 3 also shows average contact rates with providers by health status.

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1997/98 VLSS survey

The survey collected information on standard socio-demographic variables including age, sex, marital status, educational level, and household income.

Coverage of health care questionnaire. The health component of the 1997–1998 household questionnaire collected information on the following variables for 4 weeks of the survey period.

- Current health status such as illnesses (ILL) or days of limited activity (ACTDAYS) in the preceding 4 weeks, injuries (INJ), the number of days of illness (ILLDAYS) or injury. Unfortunately, there is no information on long term health status such as the presence of limiting and nonlimiting chronic conditions.
- Utilization data on self-medication (PHARVIS). The questionnaire sought information on number of contacts, total expenditure, and the amount spent on medicines, and transportation and other costs associated with the visits. For the self-medication part the questionnaire also sought information on whether the visit to the drug vendor was self-initiated or initiated by another provider. The responses to the following two questions are analyzed in this paper: ‘1. In the past 4 weeks did you come directly to a pharmacist or medicine peddler to buy medicine or seek advice on treating illness or injury without visiting another health care provider first? 2. In the past 4 weeks how many times did you come to the pharmacy in this way?’
- Health insurance status (HLTHINS) and the amount spent on health insurance in the previous 12 months.
- Health expenditures during the 4-week survey period and in the 12 months prior to the survey.

The data definitions and descriptive summary statistics are given in Table 4.

These data can support investigations both at the individual and the household level. The available frequency-of-use data can be used in event count models of utilization.

For each type of service contact, the questionnaire collected responses on the total cost of transportation, room and board, and other related costs. These data can be used to estimate average extraneous cost of health service. Although this is useful information, it cannot be used here for modeling individual choice of the type of service. Standard economic theory suggests that in choosing between two types of providers, e.g. commune health center and government hospital, the relative

<table>
<thead>
<tr>
<th>Variable</th>
<th>Definition</th>
<th>Mean</th>
<th>Std. dev</th>
</tr>
</thead>
<tbody>
<tr>
<td>PHARVIS</td>
<td>Number of pharmacy visits</td>
<td>0.51</td>
<td>1.31</td>
</tr>
<tr>
<td>PHARDUM</td>
<td>0/1 dummy for pharmacy visit</td>
<td>0.26</td>
<td>0.44</td>
</tr>
<tr>
<td>GOVHOSP</td>
<td>Number of hospital outpatient visits</td>
<td>0.049</td>
<td>0.41</td>
</tr>
<tr>
<td>PHF</td>
<td>Number of visits to private health facility</td>
<td>0.11</td>
<td>0.80</td>
</tr>
<tr>
<td>CHC</td>
<td>Number of visits to commune health centers</td>
<td>0.04</td>
<td>0.34</td>
</tr>
<tr>
<td>HOSPADM</td>
<td>0/1 dummy for hospital admissions</td>
<td>0.051</td>
<td>0.22</td>
</tr>
<tr>
<td>HOSPnite</td>
<td>Number of hospital nights</td>
<td>13.53</td>
<td>21.82</td>
</tr>
<tr>
<td>MEDEXP</td>
<td>Total medical expenditure (4 weeks plus 12 months)</td>
<td>1520</td>
<td>5139</td>
</tr>
<tr>
<td>HHEXP</td>
<td>Total nominal household expenditure</td>
<td>15273</td>
<td>13020</td>
</tr>
<tr>
<td>ln(INC)</td>
<td>log(HHEXP) or log(income)</td>
<td>2.60</td>
<td>0.62</td>
</tr>
<tr>
<td>HLTHINS</td>
<td>0/1 dummy for health insurance status; 1 for insured</td>
<td>0.16</td>
<td>0.37</td>
</tr>
<tr>
<td>AGE</td>
<td>Age in years</td>
<td>29.7</td>
<td>9.67</td>
</tr>
<tr>
<td>SEX</td>
<td>0/1 dummy for gender; 1 for female</td>
<td>0.51</td>
<td>0.49</td>
</tr>
<tr>
<td>MARRIED</td>
<td>0/1 dummy for marital status; 1 for married</td>
<td>0.40</td>
<td>0.49</td>
</tr>
<tr>
<td>EDUC</td>
<td>Completed level of education</td>
<td>3.38</td>
<td>1.94</td>
</tr>
<tr>
<td>ILL</td>
<td>0/1 dummy for illness in 4 weeks before survey; 1 for ill</td>
<td>0.41</td>
<td>0.49</td>
</tr>
<tr>
<td>INJ</td>
<td>0/1 dummy for injury in 4 weeks before survey; 1 for injury</td>
<td>0.009</td>
<td>0.098</td>
</tr>
<tr>
<td>ILLDAYS</td>
<td>Number of days of illness/injury in 4 weeks before survey</td>
<td>2.80</td>
<td>5.45</td>
</tr>
<tr>
<td>ACTDAYS</td>
<td>Number of days of limited activity in 4 weeks before survey</td>
<td>0.06</td>
<td>1.11</td>
</tr>
</tbody>
</table>
extraneous cost of the services of the two providers is relevant. The survey data only pertain to the average extraneous cost of the service actually chosen by the patient. By itself it cannot be used to construct a relative price for each user, which is what one needs for modeling purposes. Even a simple measure such as distance from different types of providers may be used to construct a more appropriate measure of the extraneous costs under the assumption that such costs are closely related to the distance.

**Health insurance**

A major new development in the health sector since the 1992–1993 VLSS survey is the emergence of a national health insurance program, Vietnam Health Insurance (VHI), that was initiated in late 1992, and which began effective operation in 1993. Three health insurance programs in Vietnam are provided under government sponsorship, comprising a compulsory national health insurance program and the two voluntary programs. During its first phase insurance coverage was provided to current and retired civil servants and to salaried employees of state-owned and large private enterprises. Benefits include the full cost of pharmaceuticals, ambulatory and inpatient care at governmental facilities to which enrollees are referred; a district or provincial hospital acts as a primary care provider. The mandatory VHI coverage does not extend to dependents of employees; see Behrman and Knowles [14,15].

A second voluntary VHI plan provides for coverage with varying benefits, on a group-basis, to VHI dependents and some other groups such as communes. A third tier of national health insurance is the voluntary plan, called Comprehensive Student Insurance (CSI). The benefits of the compulsory VHI are less variable than those of the voluntary component. The CSI plans and premiums are locally designed and administered, and show substantial variation in premiums and benefits among localities.

One estimate of the number of total (compulsory and voluntary) enrollees in the VHI program comes from VHSR (1999) that estimates this at 9.8 million in 1998, which includes about 38% voluntary enrollees. This covers roughly 12% of the population. The coverage of the target population for the compulsory component is around 77%, but it is much lower for the voluntary component and consists largely of students.

Having health insurance is positively associated with income. In the lowest 25% of the income level, 9.2% of the sample has health insurance. A high proportion of these may be those enrolled in the voluntary scheme. In the top 25% of the income distribution, 24.5% have health insurance, see Table 5.

Because the VHI premium for the compulsorily insured is a fixed percentage of the employee’s base salary, the cost of insurance varies and income serves as a partial proxy for the cost of insurance.

One of the objectives of the empirical investigations in this article is to estimate the impact of health insurance on self-medication. The foregoing account raises an important technical issue regarding the treatment of health insurance. For those compulsorily enrolled in the program, i.e. the majority, we may treat insurance status as exogenous, but for those who are voluntarily enrolled, there may be an element of discretionary choice that may be an argument for treating the variable as endogenous (see Cameron et al. [16]). However, as was noted above, enrollment is on a group (not individual) basis, which diminishes the role of individual preferences in the choice of health insurance. There is also a related data problem. The health component of the questionnaire asked only two questions about insurance; first whether the respondent had health insurance, and second, the cost of health insurance in the previous 12 months. Hence, one cannot distinguish between those who were enrolled in the compulsory insurance program and those who were not. Therefore, insurance status will be treated as exogenous in the health care utilization equation. Under this restriction it becomes possible to identify the causal parameter that links insurance and self-medication. The validity of the restriction should be scrutinized in future work based on better data that are currently not available. However, the robustness of empirical results can be analyzed by estimating separate equations for adults, females, and the young population.

### Table 5. Health insurance and income status

<table>
<thead>
<tr>
<th>Income class</th>
<th>Lowest 10%</th>
<th>Lowest 25%</th>
<th>Highest 25%</th>
<th>Highest 10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>Percentage insured</td>
<td>8.7</td>
<td>9.2</td>
<td>24.5</td>
<td>27</td>
</tr>
</tbody>
</table>

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Econometric issues

First we consider some important statistical and modeling issues. The individual utilization data are available for 27 765 cases. However, there are certain features of these data that influence econometric analysis in a crucial manner. First, because the utilization period is only 4 weeks, there are many instances of zero utilization, see Table 6. For PHARVIS the zero proportion is about 74%. The observed frequency distribution shows positive probability mass at a few other integer values, such as 1, 2, and 3, with very small mass at higher integer values. For example, for integer values, such as 1, 2, and 3, with very small shows positive probability mass at a few other about 74%. The observed frequency distribution Table 6. For PHARVIS the zero proportion is there are many instances of zero utilization, see

Table 6. Frequency distribution of pharmacy visits

<table>
<thead>
<tr>
<th>Number of contacts</th>
<th>Observed frequency (%)</th>
<th>Fitted value (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>20639 (74.43)</td>
<td>70.16</td>
</tr>
<tr>
<td>1</td>
<td>3827 (13.79)</td>
<td>15.99</td>
</tr>
<tr>
<td>2</td>
<td>1716 (6.18)</td>
<td>8.60</td>
</tr>
<tr>
<td>3</td>
<td>776 (2.79)</td>
<td>3.58</td>
</tr>
<tr>
<td>4</td>
<td>359 (1.29)</td>
<td>1.19</td>
</tr>
<tr>
<td>5</td>
<td>174 (0.62)</td>
<td>0.33</td>
</tr>
<tr>
<td>6</td>
<td>64 (0.23)</td>
<td>0.08</td>
</tr>
<tr>
<td>7</td>
<td>43 (0.15)</td>
<td>0.02</td>
</tr>
<tr>
<td>8</td>
<td>16 (0.05)</td>
<td>0.00</td>
</tr>
<tr>
<td>9</td>
<td>4 (0.01)</td>
<td>0.00</td>
</tr>
</tbody>
</table>

Notes: sample size: 27 731.

The standard count data framework of the extended Poisson regression can naturally accommodate the presence of a significant probability mass at zero. There are two standard solutions for the excess zeros problem. First the hurdles model, also called the two-part model (TPM), may be used. The second accounts for zero inflation by extending the Poisson Model. Disaggregation into sub-samples, such as the under-18 group, and/or elimination of clustered observations due to the presence of, for example, commune fixed effects can also reduce the seriousness of the problem.

In the hurdle (or two-part) model the first part is a binary outcome model, and the second part is a truncated count model. This framework accommodates corner solutions. Such a partition has the interpretation that positive observations arise from crossing the zero hurdle or threshold (Mullahy [17]; Gertler and Litvack [2]). Continuous probability distributions cannot allow for nontrivial probability mass at zero frequency. The TPM model allows the process generating zeros to be different from that generating the positives. Note also that the zeros have two possible interpretations. The first is that they correspond to 'corner solutions' in the consumer choice problem. That is, they indicate non-consumption, given current income, price, and health status. A second interpretation is that zeros indicate that the overall level of consumption, conditional on an interior solution to the choice problem, while the second part models the level of consumption, conditional on an interior solution. Hence both parts can yield estimates of economically interesting parameters. For this reason it is worthwhile to develop a binary outcome model of zero and positive number of visits to pharmacy.

The second approach based on the zero-inflated Poisson or negative binomial model. For example,
consider the following:

\[
\Pr[y_i = 0] = \varphi + (1 - \varphi) e^{-\mu_i},
\]

\[
\Pr[y_i = r] = (1 - \varphi) \frac{e^{-\mu_i} \mu_i^r}{r!}, \quad r = 1, 2, \ldots \tag{10}
\]

where \( \mu_i = \exp(x_i' \beta) \). This distribution can be interpreted as a finite mixture with a degenerate distribution whose mass is concentrated at zero. In (10) the proportion of zeros, \( \varphi \), is added to the Poisson\[\mu\] distribution, and other frequencies are reduced by a corresponding amount. One possible justification for this is the case of misrecorded observations, where the misrecording is concentrated exclusively in the zero class.

**Problem of modeling clustered outcomes**

The VLSS sample is likely to exhibit potential intra-cluster correlation within each commune because it is based on stratified sample survey methodology, with communes as the primary sampling unit. The correlation must be taken into account to obtain correct sampling variances of the regression coefficients, and without such a correction the statistical significance of estimated coefficients may be overstated [19]. The sample used in this article covers 194 communes, with between 50 and 210 observations per commune.

**Statistical assumptions.** There are two broad statistical approaches for dealing with the effects of clustering considered in this paper. The cluster-specific approach to handling intra-cluster correlation is to model the expectation of the outcome as a function of covariates and a cluster-specific term which may be either a fixed or random effect.

The Poisson fixed effects cluster model Graubard and Korn [20], specifies

\[
y_{ic} \sim P[\mu_{ic} = \alpha_c \exp(x_{ic}' \beta)],
\]

\[c = 1, \ldots, C, \quad j = 1, \ldots, N_c \]

where \( P[\cdot] \) denotes the Poisson distribution, and \( x_{ic} \) excludes an intercept and any cluster-invariant regressors. This is the standard Poisson model, except that the usual conditional mean \( \exp(x_{ic}' \beta) \) is scaled multiplicatively by the cluster specific fixed effect (FE) \( \alpha_c \). For this particular model the conditional maximum likelihood approach leads to elimination of the parameters \( \alpha_c \). We refer to this formulation as cluster-specific fixed effects (CSFE) modeling.

The alternative is the random effects (RE) approach which assumes that the cluster-specific intercept is randomly distributed across communes with a common ‘exchangeable’ parametric distribution, often the gamma distribution with mean 1 and unknown variance.

The two approaches involve different statistical assumptions and will not necessarily yield the same results. The CSFE model permits correlation between the fixed effects and the random equation error, whereas the RE model assumes that the cluster-specific intercept is uncorrelated with the covariates in the model. In this article we have used both approaches to evaluate the robustness of our empirical results. The CSFE approach follows the conditional maximum likelihood (CML) approach (Andersen [21]) in which inference is carried out conditionally on sufficient statistics for the fixed effects. However, the cluster-specific fixed effects cannot always be eliminated. To do so, we require that there exist sufficient statistics for the fixed effects. While this is possible for some distributions it is not so for all. For example, this is not possible for a truncated Poisson distribution or for the zero-inflated Poisson.

Fixed effects nonlinear panel data frameworks relevant to our analysis include binary logit and the Poisson. Hsiao [22] has discussed fixed effects panel logit models, and Haussman et al. [23] have developed the fixed effects Poisson regression for counts and some variants thereof.

Consider a sample with total \( N \) observations, which are distributed in \( C \) clusters with each cluster \( N_c \) \((c = 1, \ldots, C)\) observations and \( \sum_{c=1}^{C} N_c = N \). If the number of observations per cluster varies, the data correspond to an unbalanced panel. Even in such a case the fixed effects panel logit model can be readily adapted to accommodate the case of clustered observations by replacing the \((y_{it}, x_{it}; i = 1, \ldots, N; t = 1, \ldots, T)\) in panel data by \((y_{ic}, x_{ic}; i = 1, \ldots, N_c; c = 1, \ldots, C)\). With this modification, the expressions for CML of fixed effects Poisson and logit models can be readily derived. The details are given in the next subsection.

**CSFE, RE Poisson, and logit models.** Introducing commune-specific fixed effects is an attractive way of controlling for several factors that are relevant to health care utilization but for which we do not have data. An example is the cost of access to a public provider. We can expect systematic differences in access costs between rural and urban
areas, and large and small communes. As previously discussed the incentives for self-medication are expected to be greater in rural areas where the cost of good quality health care will be high. One can also expect systematic differences by region and location.

In the first two models, the dependent variable are assumed non-negative integer, while in the conditional logit it is 0/1 binary variable.

**Conditional fixed effect Poisson model:** Assume \( y_{cn} \) is i.i.d Poisson distributed \( P(\lambda_{cn}) \) and \( \lambda_{cn} = \exp(\mathbf{x}_c \beta) \), and \( \mathbf{x}_c \) excludes an intercept (for identification purpose), we have conditional joint probability for the \( c \)th cluster as

\[
\Pr \left( y_{c1}, \ldots, y_{cN_c} \mid \sum_{n=1}^{N_c} y_{cn} \right) = \frac{\Gamma \left( \sum_{n=1}^{N_c} y_{cn} + 1 \right)}{\prod_{n=1}^{N_c} \Gamma \left( y_{cn} + 1 \right)} \prod_{n=1}^{N_c} \left( \frac{\exp(\mathbf{x}_n \beta)}{\sum_{d \in X_c} \exp(\mathbf{x}_n \beta)} \right)^{y_{cn}} \tag{11}
\]

where \( \Gamma(\cdot) \) denotes the gamma function. The CML estimator of \( \beta \) maximizes the conditional log-likelihood function based on the preceding equation.

**Conditional fixed effect logit model:** In this model, dependent variable \( y \) takes only two values, \( 1 \) or \( 0 \). The joint conditional probability for the \( c \)th cluster with \( N_c \) observations is

\[
\Pr \left( y_{c1}, \ldots, y_{cN_c} \mid \sum_{n=1}^{N_c} y_{cn} \right) = \frac{\exp(\beta \sum_{n=1}^{N_c} \mathbf{x}_n y_{cn})}{\sum_{d \in X_c} \exp(\beta \sum_{n=1}^{N_c} \mathbf{x}_n d_{cn})} \times \frac{\Gamma \left( \sum_{n=1}^{N_c} y_{cn} + 1 \right) \Gamma \left( N_c - \sum_{n=1}^{N_c} y_{cn} + 1 \right)}{\Gamma(N_c + 1)} \tag{12}
\]

where \( X_c = \{ (d_{c1}, \ldots, d_{cN_c}) \mid d_{cn} = 0 \text{ or } 1, \sum_{n=1}^{N_c} d_{cn} = \sum_{n=1}^{N_c} y_{cn} \} \), \( \mathbf{x}_c \) excludes an intercept.

The random effects Poisson panel model with gamma distributed intercept has been analyzed in Hausman et al. [23] who provide the closed form expression for the likelihood. This can be adapted for clustered data. Further, the random effects panel logit model has been analyzed under the assumption that the distribution of the random intercept is normal. Under this assumption there is no analytical expression for the likelihood function so numerical quadrature is used to obtain the coefficient estimates; see Pendergast et al. [24].

An approach that does not requires a distributional assumption for the random component is to simply use robust variance estimation, i.e. ‘cluster robust’ standard errors obtained by adapting the so-called Eicker–White robust variance estimator to handle clustered data\(^4\). Specifically, if \( f_i (i = 1, \ldots, N) \) denotes the density for the \( i \)th observation, \( \theta \) denotes the vector of unknown parameters, then the cluster-robust variance estimator evaluated at the maximum likelihood estimate \( \hat{\theta}_{MLE} \) is given by

\[
V_C = \left[ \sum_{j=1}^{C} \sum_{i=1}^{N_j} \frac{\partial^2 \ln f_{ij}}{\partial \theta \partial \theta'} \right]^{-1} \times \left[ \sum_{j=1}^{C} \sum_{i=1}^{N_j} \sum_{k=1}^{N_j} \frac{\partial \ln f_{ij}}{\partial \theta} \frac{\partial \ln f_{ik}}{\partial \theta'} \right] \times \left[ \sum_{j=1}^{C} \sum_{i=1}^{N_j} \frac{\partial^2 \ln f_{ij}}{\partial \theta \partial \theta'} \right]^{-1} \hat{\theta}_{MLE} \tag{13}
\]

**Results**

The main focus of the discussion of results will be on the role of income and health insurance, after conditioning on a set of relevant covariates. We shall discuss several alternative variants of models specifications that were estimated for the number (frequency) of visits to pharmacy (PHARVIS), and the probability of visit to pharmacy (PHARDUM). As is customary, the income variable is proxied by logarithm of total household expenditure, denoted ln(INC). The only ‘price variable’ is the HLTHINS dummy, and note that there is no further breakdown by type of health insurance, e.g. full or dependent coverage versus student coverage. In modeling health care utilization it is usual to include measures of health status, but the available measures (ILL, INJ, ILLDAYS, ACTDAYS) largely reflect current rather than long term health status. Other conditioning covariates are AGE, SEX, MARRIED, EDUC; all variables are defined in Table 4.

**Frequency of pharmacy visits**

Detailed regression results are give in Table 7. In this case the regression models refer to frequency
Because of the overwhelming importance of expenditure on purchased medicines, the results for the frequency (counts) of pharmacy visits are of special interest.

To model counts, we have used as a starting point the standard Poisson regression, with robustly estimated variances of parameters. Given correct specification of the conditional mean function, this estimator is consistent. The results are shown in the first two columns of Table 7. A comparison between the observed and fitted frequency distribution of visits is in Table 6. This shows that the discrepancy between sample and the fitted distributions of visits is small. However, there is some evidence of overdispersion, which suggests some unaccounted unobserved heterogeneity. To reduce variance inflation the standard errors are adjusted to allow for intra-cluster correlation using the ‘cluster robust’ variance formula to due to clustering.

We also consider the possibility that clustering is due to the omission of unobserved fixed effects. Last four columns of Table 7 provide estimates of the FE Poisson model formulation for two versions of the conditional mean specification. The Poisson regression provides a good overall fit to the data. The estimates are plausible in regard to the role of the demographic and health status variables. Females have on average a higher number of visits, as do those who are or have been recently sick and limited in activity. The completed years of education variable is negatively related to PHARVIS. The role of other factors – such as being female, being married, having illness or injury, and the length of illness – is similar to that which has been found for other types of health care, that is, they all increase the frequency of pharmacy visits. One difference is that AGE does have a positive effect.

Whereas the basic Poisson regression suggests that having health insurance coverage reduces the number of pharmacy visits, it also suggests a positive income effect. The latter result about the role of income is sensitive to adjustment for cluster effects. On the other hand, FE Poisson estimates suggest that the income effect is negative and statistically significant. This result is plausible and consistent with lower income households relying overwhelmingly on self-medication in the event of illness, injury, and activity limitation. Evidence cited earlier indicated an increasing reliance on self-medication as the supply of drugs improved and the retailing of drugs became deregulated. Drugs are also dispensed at public hospitals in

Table 7. Models for frequency of pharmacy visits

<table>
<thead>
<tr>
<th>Variables</th>
<th>Poisson cluster robust standard errors</th>
<th>Poisson fixed effects</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cons</td>
<td>-3.824</td>
<td>0.099</td>
</tr>
<tr>
<td>ln(INC)</td>
<td>0.078</td>
<td>0.047</td>
</tr>
<tr>
<td>INC1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>INC2</td>
<td></td>
<td></td>
</tr>
<tr>
<td>INC3</td>
<td></td>
<td></td>
</tr>
<tr>
<td>INC4</td>
<td></td>
<td></td>
</tr>
<tr>
<td>HLTHINS</td>
<td>-0.253</td>
<td>0.061</td>
</tr>
<tr>
<td>SEX</td>
<td>0.080</td>
<td>0.033</td>
</tr>
<tr>
<td>AGE</td>
<td>0.002</td>
<td>0.001</td>
</tr>
<tr>
<td>MARRIED</td>
<td>0.106</td>
<td>0.047</td>
</tr>
<tr>
<td>ILLDAYS</td>
<td>0.042</td>
<td>0.003</td>
</tr>
<tr>
<td>ACTDAYS</td>
<td>0.009</td>
<td>0.019</td>
</tr>
<tr>
<td>INJ</td>
<td>0.164</td>
<td>0.221</td>
</tr>
<tr>
<td>ILL</td>
<td>0.560</td>
<td>0.028</td>
</tr>
<tr>
<td>EDUC</td>
<td>-0.051</td>
<td>0.015</td>
</tr>
<tr>
<td>-log-lik</td>
<td>25274</td>
<td></td>
</tr>
<tr>
<td>N</td>
<td>27765</td>
<td></td>
</tr>
</tbody>
</table>

Notes: cluster effect standard errors are adjusted for clustering by communes; robust standard errors are Eicker–White standard errors.
Vietnam, but this particular channel is available to, and more likely to be used by, the high income insured individuals.

To more closely investigate the connection between income and pharmacy visits, and to specifically allow for differences in response at different levels of income, additional regressions were fitted for probability and number of pharmacy visits, using a variants of the Poisson model based on a flexible spline specification for income variable. This specification permits the response in each income quartile to be different. Hence \( \ln(\text{INC}) \) gets replaced by four variables, denoted \( \text{INC1}, \text{INC2}, \text{INC3} \) and \( \text{INC4} \). The fixed effects Poisson regression is reestimated with this new functional form. The model is also estimated without the fixed effects but with cluster-corrected standard errors. Both sets of estimates suggest that pharmacy use is an inferior good, in the top income quartile. The model without fixed effects also suggests that pharmacy use is a normal good, with a positive income elasticity, in the lower income quartile. The estimated positive income elasticity for the lowest income quartile is at most of the order 0.30. The estimated negative income elasticity for the richest quartile is of the order –0.20. The weighted sum of these two is positive. These income elasticities are not large, but they imply that pharmacy visits appear to be an inferior good for the rich, but a normal good for the poor. This result is different from that in some previous analyses based on the 1992–1993 VLSS data (Gertler et al. [12]) that suggest that pharmacy drugs are a normal good at all income levels. Unfortunately, the relatively large standard errors on the coefficients preclude a more definitive statement.

The impact of \( \text{HLTHINS} \) on self-medication is found to be negative, as predicted by the theoretical model, and statistically significant and also sign-wise robust across a range of alternative specifications. The size of the impact is the largest in the basic Poisson regression, somewhat smaller in the FE Poisson regression, and the smallest in the cluster-adjusted model. The average negative impact on the number of visits over a four week period due to having insurance is measured by \( \exp(\text{coeficient of HLTINS}) \). For values of coefficient of around –0.3 this value is about 0.75, and for a coefficient value of about –0.10, the value is around 0.90. In all cases the estimate is unambiguously negative and statistically significant. This implied negative impact on average number of visits is large in the relevant sense – about one visit per individual. Observe from Table 4 that the average number of visits for the 4 week period is about 0.5. Thus, having health insurance eliminates the pharmacy visit for an ‘average’ individual. Therefore, this mechanism, unlike the income effect, seems a powerful drag on the extent of self-medication.

In Table 8 we present estimates of income and insurance coefficients (other covariates are not shown) based on subset FE regressions for four groups: under-18, over-22, male, and female. These results confirm the robustness of the qualitative result that on average pharmacy visits decline in number with rising income and with having health insurance. The insurance effect is robustly negative, and, except in the case of young males, the income effect is also negative. Yet these estimates do not suggest a major revision of the quantitative impact discussed in preceding paragraphs.

The negative impact of insurance is consistent with the interpretation that self-medication is a risky form of health care and it is avoided as the relative price of higher quality (less risky) professional care falls. In Vietnam higher quality care comes from public hospitals and having health insurance causes substitution towards such care. Since insured patients do not have the alternative

| Table 8. Income and insurance impact estimates from disaggregated FE models |
|-----------------------------|-----------------------------|-----------------------------|-----------------------------|-----------------------------|
|                             | Male                        | Male                        | Female                      | Female                      |
|                             | Age < 18                    | Age < 18                    | Age > 22                    | Age > 22                    |
| Sample size                 | 5432                        | 5239                        | 6635                        | 7671                        |
| ln(INC)                     | 0.006 0.055                 | –0.185 0.055                | –0.152 0.038                | –0.108 0.029                |
| HLTINS                      | –0.056 0.079                | –0.153 0.081                | –0.244 0.0522               | –0.136 0.048                |

of being treated in private facilities, and are required to obtain their treatment in public hospitals, we suggest that the observed effect of insurance ‘underestimates’ the true effect of health insurance (relative to the unconstrained choice model) by Le Chatelier principle. To fully test the hypothesis that demand is diverted from pharmacies towards government hospitals, one should also check whether the coefficient of HLTHINS in the government hospital equation is positive and statistically significant. The results reported in Trivedi [9] show that this is indeed so.

Next we consider the robustness of the above results to a change in the distributional assumptions of the model, and compare the variability of results across different demographic groups. Table 9 contains results for income and insurance variables. For brevity the details of coefficients for other conditioning covariates have been left out. The new specification is the zero-inflated Poisson model (with standard errors adjusted for clustering by communes) that allows the probability mechanism for the zero count to be different from that for the positive counts. The coefficient of ln(income) is positive for the full sample. When ln(INC) is entered in the spline form, again the results are somewhat mixed. The positive coefficient for the lowest income quartile and a negative coefficient for the highest are unambiguous although the precision of the estimates is variable. However, for the second and third income quartiles, the results are mixed; for the third quartile the evidence suggests positive rather than the hypothesized negative response to income. As before the effect of health insurance on the number of pharmacy visits is significantly negative.

It is interesting to estimate separate equations for children and adults. Table 9 also gives results for several subsamples. Overall, therefore, the results from the ZIP model are similar to those presented earlier. The regression for females suggests that pharmacy visit is an inferior good at all income levels; however, the coefficient estimates have large standard errors so this interpretation should be taken with caution.

Previous analyses [2, 3] have suggested that a high level of self-medication is in part a consequence of low quality of care available to the lower income groups, especially in the communes. A variety of policy prescriptions have been put forward. The VHSR [3, pp. 120–121] mentions the proliferation of counterfeit and sub-standard drugs available on the market and goes on to describe the recently initiated Vietnam National Drug Policy also aimed at rational and safe use of drugs. Gertler and Litvack [2, pp. 246–247] suggest that improving the quality of service at the commune health centers, for example by improving the supply of low-priced generic drugs, would reduce self-medication. Our results suggest that an expansion of the voluntary health insurance program would also have a qualitatively similar effect on the leftward shift of the demand curve. However, exploratory models of HLTHINS showed that enrollment in the insurance program is very income sensitive, so the poor sections would be less affected by such an expanded enrollment; see Trivedi [9].

Table 9. Zero-inflated models for frequency of pharmacy visits

<table>
<thead>
<tr>
<th>Sample</th>
<th>Model</th>
<th>HLTHINS</th>
<th>ln(INC)</th>
<th>INC1</th>
<th>INC2</th>
<th>INC3</th>
<th>INC4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Full</td>
<td>ZIP</td>
<td>-0.223</td>
<td>0.087</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.029)</td>
<td>(0.015)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Full</td>
<td>ZIP</td>
<td>-0.221</td>
<td></td>
<td>0.272</td>
<td>-0.043</td>
<td>0.351</td>
<td>-0.160</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.050)</td>
<td>(0.114)</td>
<td>(0.205)</td>
<td>(0.154)</td>
<td>(0.083)</td>
<td></td>
</tr>
<tr>
<td>Age &gt; 22</td>
<td>ZIP</td>
<td>-0.248</td>
<td></td>
<td>0.224</td>
<td>-0.116</td>
<td>0.333</td>
<td>-0.167</td>
</tr>
<tr>
<td>N = 14,428</td>
<td></td>
<td>(0.059)</td>
<td>(0.123)</td>
<td>(0.213)</td>
<td>(0.168)</td>
<td>(0.090)</td>
<td></td>
</tr>
<tr>
<td>Age &lt; 18</td>
<td>ZIP</td>
<td>-0.133</td>
<td></td>
<td>0.572</td>
<td>0.121</td>
<td>0.186</td>
<td>-0.126</td>
</tr>
<tr>
<td>N = 11,142</td>
<td></td>
<td>(0.075)</td>
<td>(0.197)</td>
<td>(0.252)</td>
<td>(0.224)</td>
<td>(0.130)</td>
<td></td>
</tr>
<tr>
<td>Female; age &gt; 22</td>
<td>ZIP</td>
<td>-0.165</td>
<td>0.028</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>N = 5,690</td>
<td></td>
<td>(0.050)</td>
<td>(0.023)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Female; age &gt; 22</td>
<td>ZIP</td>
<td>-0.164</td>
<td></td>
<td>0.212</td>
<td>-0.228</td>
<td>0.317</td>
<td>-0.163</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.050)</td>
<td>(0.081)</td>
<td>(0.162)</td>
<td>(0.138)</td>
<td>(0.050)</td>
<td></td>
</tr>
</tbody>
</table>

Probability of pharmacy visit

A binomial model of the decision to visit pharmacy throws light on the determinants of the probability that a corner solution occurs. Results of logit and fixed effects logit model are given in Tables 10 and 11. Random effects models were also estimated but the detailed estimates are not reported here.

Table 10 provides estimates of the full logit model, with standard errors adjusted for clustering by commune. The main variables of interest are ln(INC) and HLTHINS. Interestingly, when ln(INC) is introduced as a single variable, its coefficient was insignificantly different from zero. As was found in frequency regressions, this result may be a consequence of some income groups reacting positively and others negatively to increases in income. The reported result allows for the spline version of the income variable. In this case once again the income effect is positive for the two lowest quartiles and negative for the two highest quartiles. However, all coefficients have large standard errors. The HLTHINS variable has a well-determined negative impact on probability of pharmacy visit.

Table 11 provides results for the fixed effects version of the logit model. Once again, with the exception of the young male group, these results provide clear evidence to support the hypothesis that the income impact on probability of pharmacy visit is negative, for the full sample and for the subsamples. For the spline version of the same model, evidence supports the view that for the non-adult subsample, income effect is positive for the poorest section of the population and negative for all demographic groups in the richest section of the population. The qualification is that sampling variability is large. For all groups an increase in insurance coverage reduces the probability of self-medication. Analogous models without fixed effects produced qualitatively similar but less clear-cut evidence. In summary, these results are consistent with those from models of frequency of visits. The results are also consistent with those from the random effects specifications that were estimated for full and disaggregated samples. Those from the full sample are relatively the most precise and these are very similar to those from the FE specification. In all cases discussed in this paragraph, the standard model is easily rejected against the alternative of either FE or RE models.

The results need some econometric qualifications. Both health insurance and household income are treated as weakly exogenous. Further, different categories of enrollment into the health insurance program have not been distinguished, resulting in possible aggregation bias. No allowance has been made for measurement errors in income.

Concluding remarks

From previous analyses of 1992–1993 VLSS data a stylized pattern of health care utilization in Vietnam has emerged. Broadly, according to this stylized view, the richer sections of the population get their health care at public hospitals, the poorer at the commune health centers, and all groups use self-medication heavily, causing the latter to dominate as the principle source of health care. In this picture, other health care providers play a relatively minor role. The stylized description says little about the impact of health insurance on the health care use pattern.

This article has attempted to develop an economic model of self-medication based on choice behavior under uncertainty. Some features
of this model have been empirically tested. The empirical results suggest that, except for children from low-income households, self-medication is an inferior good, i.e. the demand for it declines with rising household incomes. Evidence that self-medication is an inferior good is relatively clear for high income households. In the aggregate, however, the net effect of income on self-medication is estimated to be close to zero. Empirical evidence also strongly suggests that health insurance diverts demand away from self-prescribed medication and towards higher quality care available in government hospitals. These results are consistent with the view that self-medication is a low-quality and risky form of health care in Vietnam. A corollary of the preceding conclusion is that professionally provided health care services may have sizeable income elasticities. Hence, rising incomes may put additional pressure on health care services as individuals and households move away from self-medication.

The net impact of health insurance on self-medication is negative. That is, under the current organization of health care delivery, having health insurance diverts patients away from self-medication and mainly towards public hospitals (Trivedi [9]). The growth of services in the higher quality health care facilities is therefore likely to be affected in opposite directions by rising income and rising proportion of insured population.

The discussion of the preceding two paragraphs generates an apparent puzzle. During the 1993–1998 period Vietnam has experienced a high average growth rate and an expansion of its health insurance coverage – the two factors that should according to our theoretical and empirical results dampen the extent of self-medication. Yet the opposite is the case. Why? To resolve the apparent puzzle note that in the aggregate the income effect is quite weak. The ‘price effect’ emanating from the greater coverage and the scope of health insurance coverage in Vietnam is more substantial but it impacts only on a small subpopulation of the insured (about 16%). More important, these dampening effects were in all likelihood overwhelmed by a huge reduction in the price of self-medication as a result of the deregulation of the pharmaceutical sector in Vietnam. Indeed, the data [3, p. 117] show that there was around 25% reduction in the price of medicines in the five year period after 1992–1993. Continuing the current policy of expanding coverage of health insurance to a broader group of citizens and at the same time imposing some form of safety regulation on the pharmaceutical industry should slow down the use of self-medication.

Table 11. Fixed effects logit model for probability of pharmacy visit

<table>
<thead>
<tr>
<th>Sample</th>
<th>Coefficient of</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>HLTHINS</td>
<td>ln(INC)</td>
<td>INC1</td>
<td>INC2</td>
<td>INC3</td>
<td>INC4</td>
</tr>
<tr>
<td>Full</td>
<td>0.269</td>
<td>0.2433</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.059)</td>
<td>(0.046)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Male; age &gt; 22</td>
<td>0.384</td>
<td>0.261</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>N = 6635</td>
<td>(0.115)</td>
<td>(0.095)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Female; age &gt; 22</td>
<td>0.414</td>
<td>0.244</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>N = 7671</td>
<td>(0.110)</td>
<td>(0.073)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Female; age &lt; 18</td>
<td>0.268</td>
<td>0.389</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>N = 5239</td>
<td>(0.171)</td>
<td>(0.128)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Male; age &lt; 18</td>
<td>0.065</td>
<td>0.142</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>N = 5432</td>
<td>(0.167)</td>
<td>(0.122)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Finally, we note that anecdotal evidence about (say) pill-popping parents is often advanced in support of a behavioral interpretation of self-medication. Reducing harmful types of behaviorally-induced self-medication may require a different approach from one that is appropriate when self-medication is essentially due to low income levels. For example, if it can be established that heavy drug promotion, following deregulation in the retail pharmaceutical sector, stimulated inappropriate and harmful self-medication, then controls on such promotional activity may be considered. The model presented in this paper is based on the conventional theory of consumer choice and does not address the role of drug promotion in increasing individual self-medication. Of course, in both cases there may be a common source of the problem, e.g. lack of awareness of the harmful aspects of self-medication – this is a situation in which state intervention is justified on the grounds that ‘individuals are not the best judges of their own interest’ (Besley [25]). However, caution in interpreting the results is desirable because the phenomenon we have studied has both economic and behavioral aspects, and their relative importance has important implications for both the extent and type of state intervention that is appropriate. Behavioral approach may emphasize, for example, the role of habit persistence even in the absence of economic constraints and/or clinical evidence that certain types of self-medication are appropriate and effective. The cross section data used in this study are not suitable for testing habit persistence type of hypotheses. The final qualification is that the period we have studied is characterized by a transitional situation in which there is ‘a lot going on’ in the economy, and cross section data may not adequately pick up the dynamics of change.

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Notes

a. World Health Organization [1]. This paper cites a consumer interview study carried out in six Latin American countries that found that only 34% of the dispensed medicines were classified as OTC.

b. In practice, consumers can choose treatment for their illness from private hospitals, government-sponsored hospitals, government-sponsored infirmaries, local pharmacies and ‘alternative’ medicines, to name just a few possibilities. Assume all but one, say, the private hospital care, are risky investments in that the return to each treatment is a random variable \( R - e_i, i = 2, 3, \ldots, \ell \). Then the problem becomes

\[
\max_{v, Q_2, \ldots, Q_{\ell}} E \left[ U \left( y - V - \sum_{j=2}^{\ell} p_i Q_j, h_0 + RV \right. \right.
\]

\[
\left. \left. + \sum_{j=2}^{\ell} (R - e_i) Q_j \right] \right]
\]

This is a multi-variate extension of Problem (2).

c. The contents of this largely descriptive section overlap with Section 3.2 of Trivedi [9].

d. Most variance estimators used in this study can be adapted to handle clustered observations in this way. For example, this is possible in the computer program STATA 7.0 for a variety of estimators.

e. See Cameron and Trivedi [18, Chapters 3 and 9] for a detailed discussion of the count data models used here.

Appendix

The law of diminishing marginal utility of income

We note that the first order conditions for (7) are

\[ F_c(c, V, Q) = u'(c) = \lambda \]

\[ F_V(c, V, Q) = (b - h_0 - RV - \mu Q)R = \lambda \]

\[ F_Q(c, V, Q) = (b - h_0 - RV - \mu Q)\mu - \sigma^2 Q = \lambda p \]

where \( \lambda \) is the Lagrange multiplier. Let \( D_b < 0 \) be the determinant of the bordered Hessian matrix. A direct computation shows that

\[
\frac{\partial^2 \lambda}{\partial y} = \begin{vmatrix}
1 & u'(c) & 0 & 0 & 0 \\
0 & -R^2 & -\mu R & 0 & 0 \\
0 & -\mu R & -\mu^2 & -\sigma^2 & 0 \\
-1 & -1 & -p & -1 & 0
\end{vmatrix}
\]

\[
= \frac{-u''(c)R^2 \sigma^2}{D_b}
\]
Since $D_b < 0$, $\partial \lambda / \partial y < 0$. Thus, the law of diminishing marginal utility of income holds.

Substitutes in utility

A straightforward exercise shows that

\[
\begin{align*}
\overline{\partial \lambda / \partial h_0} &= 1 \begin{vmatrix}
0 & 0 & 0 & 0 \\
0 & -R^2 & -\mu R & R \\
-1 & -1 & -\sigma^2 & \mu \\
-\mu & 0 & 0 & 0
\end{vmatrix} \\
&= -u'(c)R\sigma^2 / D_b \\
&= \frac{-u'(c)R\sigma^2}{D_b}
\end{align*}
\]

Since $D_b < 0$, $\partial \lambda / \partial h_0 < 0$. In other words, income and initial health are substitutes in utility.

The existence of $p_e$

Suppose the optimal solution is $(c, V, 0)$. An increase in $p$ makes the budget set $\{(c, V, Q) : c + V + pQ \leq y\}$ smaller, while leaving the indifference surfaces of $F(c, V, Q)$ invariant. Therefore, if $(c, V, 0)$ is the optimal solution to Problem (7), it would remain so with increased $p$. On the other hand, a substantial decrease in $p$ will make $(c, V, 0)$ suboptimal, and imply an interior solution. Therefore, there exists a critical $p_e$ below which the optimal solution is interior. Thus, the demand curve for $Q$ is vertical for $p > p_e$, and has a finite slope when $p < p_e$.

The income expansion path

Let $(c, V, 0)$, with $c > 0, V > 0$, be the corner solution to Problem (7) when the income level is $y$. Suppose the expansion path may exit the $(c, V)$-plane with increased income. Then there exists an income $y' > y$ such that the optimal solution is $(c', V', Q')$, with $c' > 0$, $V' > 0$, and $Q' > 0$. Then a reduction in income from $y'$ to $y''$ with $y \leq y'' < y'$ increases self-medication, since we have shown that self-medication is an inferior good for interior solutions. Hence, the corresponding optimum $Q''$ satisfies $Q'' \geq Q > 0$, which is bounded from below. Tracing the income expansion path backward from $y'$ to $y$, we will never return to $(c, V, 0)$. This is a contradiction.