Public health infrastructure and equity in the utilization of outpatient health care services in Peru

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This article analyzes the magnitude and nature of socioeconomic differences in the utilization of outpatient health care services in Peru. In particular, it explores the potential equity-enhancing effect of the expansion and improvements in the network of health centres during the 1990s. The Peruvian health reform made relatively little progress in terms of the reform agenda promoted internationally during the 1990s. Nevertheless, the expansion of health centres and the improvements in their equipment has been noteworthy during the same period. Using the 1997 survey of the Peruvian Living Standards Measurement Study (PLSMS), we find large differences in the utilization of outpatient health care services. The richest to poorest quintile ratio is 1.9, and even larger in rural areas. Estimating a probit model with random effects at the district level to control for the systematic geographic bias associated with the optimal public allocation of such infrastructure, we find the income effect to be very large, even after controlling for other socioeconomic characteristics. Finally, we also find that the expansion of the public network of health centres has indeed an equity-enhancing effect, but this is rather small. These results indicate that although the expansion of the public network of health facilities may be necessary, it is not sufficient to promote equity in the utilization of health care services by Peruvian adults, especially in rural areas. It is important to look deeper into the costs of consultations and drugs as economic barriers to the utilization of health services by the poor. In particular, the expansion of health insurance mechanisms for the poor should be carefully monitored and evaluated.

Key words: public infrastructure, health equity, health reform, Peru

Introduction and background

The need to reform the health care system to offer preferential attention to the poor has won great acceptance among different social groups during the last decade. Birdsall and Hecht (1995), for instance, state the argument dramatically: ‘the level of tolerance for the inequity in health is less than the tolerance for inequity in income or education since differences in health are, literally, a matter of life or death’.

The health reform agendas discussed in many developing countries during the 1990s shared several common features, influenced by the work of multilateral organizations (Whitehead et al. 2001). The prototype agenda included increases in public investments in basic health centres, massive educational programmes in preventive health care and family planning, the use of packages of basic health services, etc. (World Bank 1993). Reforming the financing of health care costs for the poor and the privatization of the provision of publicly funded health services were also very high on the agenda. The former was motivated by the idea of reducing the large income effect often found in the patterns of utilization of health care services. The latter was motivated by the idea of increasing competition to improve efficiency and quality in the provision of health services.

The implementation of that agenda varied substantially across countries, however, affected by particular social, economic and political conditions. In the Peruvian case, the stated goals of the needed health sector reform were to promote equity, efficiency and quality in the provision of health care (Aguinaga 1997). Nonetheless, it is not easy to establish the agenda that was actually being promoted by the Peruvian Ministry of Health (MOH), because the sector was as disarticulated as the rest of the Peruvian state. Most of the prototypical health reform agenda was promoted at some point by at least one of the special projects that worked within the MOH with funding provided by a particular international agency. Nevertheless, prototypical reforms in the financing and provision of health services were blocked from the very beginning given the serious political opposition of different stakeholders.

In that context, the most salient feature of the Peruvian health sector during the 1990s was the large increase in public and private health expenditures. Total health expenditures recovered from US$1.2 billion in 1993 to US$3 billion in 1996 (World Bank 1998). The MOH budget doubled in real terms during that period (Cotlear 2000). Following the guidelines of the prototype reform agenda, new funds were mostly assigned to targeted programmes focusing on preventive activities and primary care. Other government agencies made significant investments in public health infrastructure. The Peruvian social investment fund, FONCODES, alone built or rehabilitated 1200 basic health centres. The result was a significant expansion of the public network of health centres and hospitals, and improvement of the quality of equipment available in the older facilities (Cotlear 2000).
Other important aspects of the Peruvian health reform refer to a programme for the decentralization of management in the provision of health reform and the public funding of subsidized health insurance for children studying in public schools, pregnant women and the newborn. The first refers to the creation of a local administration committee (CLAS) to manage primary care health centres with participation of the facility's administrative officials, local authorities and civil society. By 1998, the programme involved about 10% of the MOH facilities and its initial evaluation has been very positive, with significant improvements in the access to health care services by the poor and patients' satisfaction with the quality of the services (Cortez 1998). Still, problems remain due to the lack of commitment from regional health authorities (DISAs) and lack of coordination between individual decentralized experiences (World Bank 1998). Some of these difficulties are still planned to be addressed by working with CLAS defined for a network of health centres rather than individual ones.

Health insurance systems are of recent creation and clearly aim at improving the access to health care services and medicines by three of the most vulnerable groups. They were initially separated into a public school-based health insurance and a mother–child health insurance, but were recently unified in 2001 under the new government. The school-based insurance (Seguro Escolar) was initiated in 1997 and by 1998 had already covered 2 million consultations. On the other hand, the mother–child insurance did not start until 2001, just before it was unified with the Seguro Escolar. To my knowledge, no rigorous evaluation has been published yet, but facilities tend to complain that reimbursements for the school-based insurance were slow and insufficient, excluding important costs and incentives to the medical staff which are important to avoid disincentives for the provision of services (World Bank 1998).

In summary, the evolution of the Peruvian health sector during the 1990s was marked by a significant recovery of public and private expenditures in health. New public funds were mainly devoted to targeted programmes focused on primary care and the expansion and rehabilitation of public health infrastructure. There is already some evidence that targeted programmes performed well in reaching the poor and had impact on the reduction of infant mortality rates among the poor. More questions remain on whether the expansion of health infrastructure has helped to promote equity in the access to health services in Peru. It could be equity-enhancing if the expansion focuses on poorer districts where individuals have to travel several hours or days to reach a health facility. Nevertheless, even in that case, the expansion of the public network could benefit the wealthier first or to a greater extent, especially for outpatient care and, if other economic barriers remain unchanged, in particular, the costs of consultation, hospitalization and medicines.

The health equity literature has increasingly focused on the relationship between income and health even in developed countries (Smith et al. 1999; Kahn et al. 2000; Kaplan 2000; Lynch et al. 2000; Miller and Paxson 2000). The estimation of the magnitude of such a relationship is obtained not by comparing the health status of individuals with different income, but using a multivariate approach that controls for the effect of other correlates. Several mechanisms can be identified to explain such a relationship. For developed countries, Smith et al. (1999) argue that healthy lifestyles are often less accessible for the less favoured, either due to lack of resources or lack of information on the benefits of certain eating and exercise habits. Also, the lack of resources generates stress on the individual, which, in turn, leads to his/her inability to assume a healthy lifestyle. However, in developing countries it is likely that the limitations set by the lack of resources to gain access to quality health care services are more important. This accessibility problem can be connected to the cost of obtaining those services, either for the monetary price charged for consultation and drugs, or for the time required to get to the location of the health facility.

Another line of the literature argues that it is not only absolute income levels that matter but also the relative levels (Lynch et al. 2000). Again, psychological effects may arise from the self-identification of the individual about his/her location in the society’s ladder. A material explanation argues that the relationship is based more on the under-provision of public goods in more unequal societies, in particular health infrastructure. The 1990s brought increasing evidence to support a direct relationship between health outcomes in a society and income inequality (Wilkinson 1992; Kaplan et al. 1996). However, more recent evidence has seriously questioned these results, suggesting that the previous evidence was mainly the result of a non-linear relationship between income and health at the individual level (Deaton 2001).

The relationship between income and the health status of adults is often difficult to assess in developing countries due to the lack of adequate databases (Savedoff and Schultz 2000). Household multi-purpose surveys tend to include proper income measures but very limited health status variables, the most common being self-reported morbidity indicators. On the other hand, most health surveys do include a wide range of health-related variables but poor indicators of the socioeconomic status (SES) of the individual or household. Given these data limitations, this article does not focus on the income effects on the health status of the individuals, but on their utilization of health care services. A recent study (Makenen et al. 2000) reports significant differences in the utilization of health care in developing countries, especially in Latin America (Guatemala and Paraguay). They also argue for the need to use a multivariate approach to further understand the socioeconomic differences in the use of health care services.

This article primarily presents empirical evidence on the nature and magnitude of the effect of socioeconomic status upon the utilization of outpatient health care services in Peru. Secondly, a multivariate approach is used to explore the magnitude of the income effect, controlling for other socioeconomic characteristics of the individual, his/her household and the district. Thirdly, I explore the role of the availability
of health infrastructure in the district and evaluate its potential equity-enhancing effect. The estimation is based on a random effects model considering the distortions that come from the systematic bias in favour of the richer districts in the distribution of the public network of health centres.

**Methods**

The empirical analysis is based on the combination of three datasets, merged at the district level. The main source of information is the survey of the 1997 round of the Peruvian Living Standards Measurement Survey (LSMS) with a full sample of 19,480 individuals (3,843 households) and 11,852 of them being at least 16 years old. The LSMS includes a wide range of household and individual socioeconomic characteristics, and individual information on the occurrences of illness and the utilization of outpatient health services during the 4 weeks prior to the interview. The second source is the Population Census, from which we use the proportion of individuals with at least one unsatisfied basic need (UBN). Finally, we use the Census of Health Infrastructure, from which we obtain the information on the number of health centres by district. The two census databases include information for the 1,813 districts formed in Peru by 1996, but we only use the information of the 242 districts included in the LSMS database.

I focus here on outpatient consultations by adults (16 years or older) who reported being sick during the recall period of the survey for health care-related variables (4 weeks). These restrictions imply a national sample of 4,183 individuals. The restriction to outpatient care is induced by the database used in this study, which includes very few individuals reporting inpatient care. Also, the analysis is restricted to consultations with health professionals, excluding cases in which individuals report a consultation with a pharmacist or unlicensed or traditional practitioner.4

The socioeconomic differences in the utilization of health care are presented using two different indicators. The first is the richest/poorest quintile ratio for health care utilization, a simple and useful measure but one that could hide important features outside the extremes. The second measure is the concentration index. This index is an extension of the Gini coefficient, widely used in the literature of economic inequalities. For current purposes, it is based on the proportion of consultations concentrated among individuals of different socioeconomic status.5 This index can take values between –1 and 1. A positive (negative) number means that the share of consultations for the poorer is smaller (larger) than its proportion in the population, for different definitions of poor. The larger the positive number, the less equitable the society or area under analysis is.

In addition, a multivariate analysis is used to properly establish an income effect in the utilization of outpatient health care services. Differences in health care utilization across quintiles cannot be attributed exclusively to differences in income, since there are other characteristics such as education or age that also change across quintiles and affect the need and access to health care services. The purpose of this exercise is then to see the extent to which that rich/poor ratio changes when a regression that controls for these other characteristics is used.6

The other purpose of the multivariate analysis is to evaluate the equity-enhancing effect of an expansion in health infrastructure. In order to do this correctly, we face the problem that the distribution of the health infrastructure is not random, but follows an optimization process by the government. Such optimization may imply a systematic bias of that distribution in favour of either the poorer or richer districts, depending on the relative prevalence of altruistic motives or the weight of interest groups (Rosenzweig and Wolpin 1986). Such bias would, in turn, lead to a biased estimation of the parameters of interest under a pooled regression in the presence of unobservable district and individual characteristics. If health infrastructure has a pro-rich (pro-poor) bias, then the income effect under a pooled regression will be biased upwards (downwards). The optimal procedure would be to control for individual unobservable characteristics by estimating a random effects model with longitudinal data, because individuals would use the district information to decide on their utilization of health care. The limitation here is that I only have a cross-sectional dataset, so that each individual is observed only once. In this case, however, we can control for district unobservable characteristics by estimating a probit model with random effects at the district level, taking advantage of the fact that we observe those units more than once since there are several individuals per district.

Formally, the following model is estimated:

\[
c_i = X_i \beta + y_i \gamma + Z_i \delta + \mu_i + \epsilon_i
\]

where \(c_i\) is a dichotomic variable that takes the value of 1 if the individual \(i\), who resides in district \(j\), does consult a formal health care provider given the event of an illness or accident during the 4 weeks prior to the survey. \(X\) denotes the vector of individual characteristics such as age, education, gender, among others, and \(y\) denotes the income or SES variable, which in this case is proxied by the log of household per capita expenditures. \(Z\) denotes the vector of observable characteristics of district \(j\), including the availability of health infrastructure and the marginality index, as measured by the proportion of individuals with at least one unsatisfied basic need (UBN).7 Finally, \(\mu_i\) denotes the vector of unobservable characteristics of district \(j\).8

The estimated coefficient \(\gamma\) can tell us if the income effect is positive and significant, but being a probit regression, it does not give us an idea of the magnitude of the effect. To do so, we would have to use the estimated coefficients to calculate predicted probabilities under different assumptions. The predicted probabilities \(\hat{p}_j = \Pr(c = 1|y, X, Z)\) can be estimated, holding constant the vectors \(X\) and \(Z\), say at sample mean values. With it, we can calculate the richest quintile to poorest quintile ratio, call it \(A_j\). We can call this indicator the income effect in the sense that, having held constant other correlates, the estimated difference in consultation rates can be solely attributed to income differences.
Finally, to evaluate whether the expansion of health infrastructure in the district has a significant equity-enhancing effect, an interaction term is added between income and health infrastructure to (1), and obtain the income effect for different fixed levels of the health infrastructure variable.

In general, the estimation of the parameters in (1) might require the use of the Heckman two-step procedure to correct for the selectivity problem, since the question is only asked to individuals who reported an illness or accident during the reference period, whose characteristics might differ significantly from the healthy group (Greene 1997). Such a procedure was applied but is not reported here because the selectivity bias was found not important.9

Results

The health infrastructure in Peru has traditionally been concentrated in richer urban areas, as can be seen from Table 1. In 1992, Peru had a total of 4318 health facilities, but the richer districts had 1462 health facilities, three times the number located in the poorer districts. Although this distribution could be explained by the distribution of the population itself, it does not change the fact that individuals living in poorer districts have a harder time reaching a health centre. The other interesting information in Table 1 relates to the bias shown by the expansion of health infrastructure between 1992 and 1996. During that period, the number of health facilities grew by almost 50% (2055) and about 20% (394) of that expansion concentrated on districts of the poorest quintile. For these districts, though, such expansion implied a growth of 82% in their number of facilities, compared with an increase of only 20% in the richer districts. These figures confirm that the expansion in health infrastructure was large, but additionally show that it had a bias towards the poorer areas. The question is whether this expansion has helped to reduce socioeconomic differences in the utilization of outpatient health care services by adults.

Table 2 shows clearly that the utilization of outpatient consultation services in Peru among those who report an illness has a clear trend in favour of the wealthier. Only 25% of adult individuals in the poorest quintile who reported being sick consulted a formal health facility, whilst for the richest quintile the figure reaches 48%. That is, the rate of outpatient consultation for the richest quintile is 1.9 times larger than for the poorest quintile. This result is consistent with the importance of education and income in the use of these services. Note that the consultation rate is smaller and inequality slightly larger in rural areas (2.0) with respect to urban areas, especially Lima (1.3). These differences would be consistent with a greater availability of public health infrastructure, as well as information about the need for formal medical attention, in large urban areas like Metropolitan Lima.

Another way to show the relationship between income inequality and inequality in access to health care is by using the concentration indexes (CI) associated with the utilization of these services. Table 3 shows, indeed, that the use of outpatient services by adults is concentrated in the richer segments of the population (CI = 12.2 > 0). The interesting additional information provided by Table 3 is that such bias against the poorer is reversed when looking at health care services in the public facilities run by the MOH, especially in urban areas and in primary level centres. The CI for the utilization of MOH health centres by urban residents is negative (–11.1), indicating a pro-poor bias. The CI in rural areas is still positive (CI = 1.2 > 0), indicating a pro-rich bias, but much smaller than that in MOH hospitals (CI = 11.1 > 0).

However, the pattern observed in Table 2 cannot be regarded as an income effect since it may also include other socioeconomic factors such as age, gender and the level of education of the individual. A multivariate analysis is needed

<table>
<thead>
<tr>
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<th></th>
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</thead>
<tbody>
<tr>
<td>Peru</td>
<td>4318</td>
<td>6373</td>
<td>2055</td>
<td>48%</td>
</tr>
<tr>
<td>Poorest quintile</td>
<td>483</td>
<td>877</td>
<td>394</td>
<td>82%</td>
</tr>
<tr>
<td>Richest quintile</td>
<td>1462</td>
<td>1761</td>
<td>299</td>
<td>20%</td>
</tr>
</tbody>
</table>


Table 2. Use of outpatient medical services by expenditure quintiles

<table>
<thead>
<tr>
<th>Quintile</th>
<th>I</th>
<th>II</th>
<th>III</th>
<th>IV</th>
<th>V</th>
<th>Total</th>
<th>V/I</th>
</tr>
</thead>
<tbody>
<tr>
<td>Peru</td>
<td>25.1</td>
<td>35.2</td>
<td>40.0</td>
<td>45.3</td>
<td>48.0</td>
<td>39.1</td>
<td>1.9</td>
</tr>
<tr>
<td>Urban</td>
<td>29.5</td>
<td>39.8</td>
<td>43.9</td>
<td>46.3</td>
<td>49.6</td>
<td>42.3</td>
<td>1.7</td>
</tr>
<tr>
<td>Rural</td>
<td>19.9</td>
<td>27.8</td>
<td>27.6</td>
<td>37.4</td>
<td>38.8</td>
<td>31.0</td>
<td>2.0</td>
</tr>
<tr>
<td>Metropolitan Lima</td>
<td>35.9</td>
<td>40.7</td>
<td>43.1</td>
<td>44.4</td>
<td>48.1</td>
<td>42.8</td>
<td>1.3</td>
</tr>
</tbody>
</table>

a proportion of adult individuals (>16 years) who reported an illness and received medical attention. Different quintiles are constructed for each sub-sample.

Table 3. Concentration index in the use of medical services, 1997 (outpatient consultation)*

<table>
<thead>
<tr>
<th>Source</th>
<th>System</th>
<th>All MOH facilities</th>
<th>MOH hospital</th>
<th>MOH centres</th>
</tr>
</thead>
<tbody>
<tr>
<td>Global</td>
<td>12.2</td>
<td>–7.4</td>
<td>–0.7</td>
<td>–12.7</td>
</tr>
<tr>
<td>Urban</td>
<td>9.1</td>
<td>–9.3</td>
<td>–7.5</td>
<td>–11.1</td>
</tr>
<tr>
<td>Rural</td>
<td>12.3</td>
<td>4.5</td>
<td>11.1</td>
<td>1.2</td>
</tr>
</tbody>
</table>

*sample of adult individuals (>16 years) who reported an illness and received medical attention.

to control for these other factors and obtain an effect that is more directly related to the lack of economic resources by individuals. As indicated in the previous section (methods), a pro-rich (pro-poor) bias in the allocation of health infrastructure would in turn bias the estimated income effect upwards (downwards). The distribution of health infrastructure in Table 1 leaves us uncertain about the size of the bias because public health infrastructure in Peru has had, and still has, a strong pro-rich bias, but the recent expansion has a strong pro-poor bias. Only the estimation of the probit model with random effects at the district level will offer insights on which of the effects dominate.

Table 4 presents the results of the probit model with and without district random effects for the whole sample and then for the urban and rural sub-samples separated. The control variables include individual characteristics such as education, age and gender, household per capita expenditures (LPCE), and district characteristics such as the per capita number of public health facilities (PPHF) and the percentage of households with at least one unmet basic need (PUBN). These two variables are also included in interaction with LPCE to evaluate whether the income effect varies with the level of those variables.

Several interesting insights come out of the analysis of these results. First of all, the likelihood ratio test does reject the hypothesis of non-random effects when using the full sample and the rural sample, but it does not reject it for the urban sample. With the full sample, changes in the coefficients for the per capita public facilities (PPHF) and their interaction with the LPCE indicate that district unobservables imply that the pooled regression biases downwards the parameters of interest, but not by much.

In the full sample, more educated, female and older individuals tend to use these health services more, but again important differences arise when separating urban and rural samples. Education is more significant and gender has also a larger effect among urban individuals. The increasing age effect is only a pattern for urban individuals; no differentiated effect is observed for the rural sample.

The coefficient for the LPCE is much larger in rural areas (0.74 vs. 0.08 in urban areas), and even not significant in urban areas. However, in the presence of interaction terms, the effect of income on the utilization of health care services also includes the two interaction terms multiplied by a particular level for the interacting variables, PPHF and PUBN. The last row of Table 4 reports that effect when using sample mean values for the interacting variables, indicating that it is largely significant in both samples, and not that different. The coefficient in urban areas is 0.24 while for rural areas it is 0.28.12

The district health infrastructure (PPHF) and the marginality variables are significant for the urban sample, but not for the rural one. The same occurs for the interaction terms. This implies that the expansion of PPHF may have an equity-enhancing effect only in urban areas.

The coefficients in Table 4 do not readily tell us much about the magnitude of the effect of each variable, including the income effect, considering that we use a probit model, unstandardized independent variables and a log formulation for household expenditures. As described in the previous section (methods), we can estimate \( \beta_1 = \Pr(c = 1|y, X, Z) \) and calculate the richest quintile to poorest quintile ratio for the consultancy rate. This procedure leads, at the national level, to an income effect of 1.8 (see first column in Table 5). The income effect is slightly smaller in urban areas (1.6) and slightly larger (1.7) in rural areas. In any event, income differences alone make adult individuals of the richest quintile about 1.8 times more prone to use formal health services when they are sick. This result indicates that controlling for other characteristics of the individual, the household and the district does not reduce the differences in the utilization of outpatient health care services across quintiles.

The nature and magnitude of the interaction coefficients are the key parameters in evaluating the effect of public health infrastructure on equity. The positive estimate for the coefficient of the interaction term LPCE*PPHF indicates that the income effect is larger in poorer districts, that is, in districts with a higher proportion of households with at least one unmet basic need. On the other hand, the negative estimate for the coefficient of the interaction term LPCE*PUBN in Table 4 suggests that the income effect is smaller in districts with more health facilities. That is, the expansion of the network of health facilities would have an equity-enhancing effect. To evaluate the magnitude of this equity-enhancing effect, the estimates in Table 4 can be used to calculate the income effect for different levels of per capita health infrastructure in the district. Table 5 evaluates the effect of a pro-poor expansion of health infrastructure, in the sense that new facilities are concentrated in less endowed districts to increase their per capita number of health facilities up to different higher levels starting at 0.3, the average level, and then increasing it up to 0.9.13 At the national level, the income effect estimated when setting this variable to its average level is 1.83, while the one simulated by tripling the health infrastructure availability falls only to 1.78. This simulation shows that a policy that focuses on the expansion of health infrastructure is indeed equity-enhancing, but its effect is rather small.

Discussion and conclusions

This study analyzes the nature of the socioeconomic differences in the utilization of outpatient health care services in Peru using the information provided by the 1997 LSMS survey. In particular, the analysis focuses on the magnitude of the income effect, when controlling for other correlates, and the equity-enhancing effect of improvements in the geographical distribution of public health centres, considering that such infrastructure was significantly improved during the 1990s (Table 1).

The study has found that differences in the utilization of health care by socioeconomic status are larger in Peru than in other countries of the region. The rich to poor ratio for the proportion of ill people that used outpatient consultations in Peru is 1.9 (Table 2), larger than those reported for other Latin American countries by Makinen (2000) (1.5 in Guatemala and 1.6 in Paraguay) or Sapelli and Vial (1998) for Chile (1.4). On the other hand, rich to poor ratios for other health services in Peru are even larger, especially those for women. The World Bank (2000) reports a rich to poor ratio of 2.8 for the proportion of pregnant women with more than two antenatal care visits, and of 7.0 for deliveries attended by a medically trained person.

The income effect, obtained through a multivariate analysis, is also very large (1.8), suggesting the persistence of important economic barriers to the utilization of outpatient health care services in Peru. Finally, improving access to health infrastructure in less endowed districts is found to be equity-enhancing, but its effect is rather small (Table 5). This result is actually similar to that found by Hutchinson (1999) in the Ugandan case, and is very important in the context of the Peruvian health sector reform, since the expansion of the public health infrastructure was one of the major developments during the 1990s.

This latter result indicates that although the expansion of the public network of health facilities may be necessary, it is not...
sufficient to promote equity in the utilization of health care services by Peruvian adults, especially in rural areas. It is important to look deeper into the costs of consultations and drugs as economic barriers to health services utilization by the poor. In particular, the expansion of health insurance mechanisms for the poor should be carefully monitored and evaluated.

Finally, it would be useful to verify these results with a longitudinal database that can control for individual unobservables. For this, we need to combine the information from utilization surveys with that of the geographical distribution of the expansion and improvements in health centres. It is still possible to design an evaluation strategy along these lines for health insurance offered now to children of school age and women of reproductive age, since these programmes have not yet expanded nationally.

### Endnotes

1. The specific proposals moved from a system of user fees for the use of public medical services with exemptions for the poor, in the late 1980s and early 1990s, to broadening access to health insurance in the late 1990s (see Whitehead et al. 2001).

2. See, for instance, the survey questionnaires for the Living Standards Measurement Study (LSMS) developed by the World Bank, or those for the Encuesta Nacional de Hogares (ENAHO) being promoted by Inter-American Development Bank in Latin America.

3. This would be the case, for instance, with the Demographic and Health Surveys (DHS) applied by Macro International in at least 44 developing countries.

4. In the 1997 LSMS sample of sick adults, about one in five individuals who reported having looked for a consultation chose this type of provider, but the most common case was the use of the advice of a pharmacist (eight out of 10 cases). Very few individuals report the use of traditional healers.

5. The concentration index is a generalization of the Gini coefficient in the sense that the cumulative variable may differ from the ordering variable. The specific formula is the following:

\[ 2C = 1 - 2 \int_0^y L(y, S^k) dy. \]

6. Note that I do not attempt to control for the endogeneity or simultaneity involved in the estimation of the coefficient for income, and in that sense, the estimated effect cannot be regarded as a causal relationship. The reason to omit this correction with an appropriate instrumentalization of the variable is that it is always difficult to obtain valid instruments with a cross-sectional sample. I consider that such an attempt would be better achieved with a longitudinal database.

7. The information comes from the Instituto Nacional de Inmigración y Estadística (INIE) and is based on district illiteracy rates, the percentage of households in the district that lack proper ceiling and proper connections to treated water sewage systems.

8. The assumptions of the model with random effects are the following:

\[ E[\varepsilon_0] = E[\mu_j] = 0, \quad E[\varepsilon_0^2] = \sigma^2, \quad E[\mu_j^2] = \sigma^2. \]

\[ E[\varepsilon_0 \mu_j] = 0 \forall i, j, k \text{ and } E[\mu_j \mu_k] = 0 \forall j \neq k \]

9. This result is actually consistent with a stylized fact for self-reported morbidity measures. They do not show a clear decreasing pattern with respect to socioeconomic status, and sometimes present a slightly increasing one. For a discussion on that evidence, see Murrugarra and Valdivia (2000) or Sapelli and Vial (1998).

10. The numbers reported in Table 3 correspond to the formula in endnote 5 but multiplied by 100. Consequently, the range here goes from –100 to 100.

11. This pattern is consistent with the findings reported by Petera and Cordero (2001). Although their estimates include utilization patterns by all individuals and those in Table 2 are restricted to adults (age 16 or older).

12. Similarly, a regression that omits the interaction terms implies similar estimated income coefficients for both samples, urban and rural.

13. About 80% of the districts in Peru had levels below 0.9 in 1996.

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**Biography**

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