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## **Illness reporting and demand for medical care in rural Burkina Faso**

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### **Abstract**

The issue of illness reporting in modelling demand for health care in low- and middle-income countries can be handled according to either of two conceptually-different constructs: (a) considering illness reporting behaviour as endogenous to demand; or (b) considering demand itself as the outcome of a sample selection phenomenon. In this paper, we take the second viewpoint and estimate the demand for medical care with an estimator that uses Heckman-type. Empirical estimates based on household survey data from rural Burkina Faso suggest that there are some implications of illness reporting behaviour for modelling the demand for medical care ( $\rho = -0.73$ ;  $p = 0.04$ ).

**Key words:** demand, health care, sample selection, Heckman, Burkina Faso

## **Introduction**

With less than 0.5 contacts per capita per year, low utilisation of health services remains one of primary policy concerns in low and middle income countries (Raberg & Jeene, 2002; Stierle, Kaddar, Tchicaya & Schmidt-Ehry, 1999). The World Health Organisation estimates that in these settings, excessive morbidity and mortality are largely due to failures on the part of health systems to guarantee adequate access to otherwise available treatment (WHO, 2002).

Substantial efforts have been channelled towards understanding what causes health service utilisation to be so low. In particular, thanks to recent advancements in econometrics and computing, health economists have developed demand models which attempt to predict accurately consumers' behaviour in relation to health and medical care (Ching, 1995; Gertler & van der Gaag, 1990; Hidayat, Thabrany, Dong & Sauerborn, 2004; Hotchkiss, 2001; Hotchkiss, Mock & Seiber, 2002; Sahn, Younder & Genicot, 2003).

Most of these models, however, rely on samples that are truncated, i.e. samples based on the positive reporting of an illness. A few published articles have pointed to the inaccuracy of the estimates that these models may provide. The inaccuracy arises due to selectivity bias that occurs when the characteristics of those who report themselves as ill are significantly different from those who do not (Akin, Guilkey, Hutchinson & McIntosh, 1998; Rous & Hotchkiss, 2003). The alternative approach proposed is a full-information maximum likelihood estimator that is believed to be able to correct for endogenous illness reporting. This approach assumes that the same unobserved variables correlate with both the decision to report an illness and the decision to seek

medical care (Akin et al., 1998; Rous & Hotchkiss, 2003). In other words, illness status has merely an intercept effect on the demand, resulting in a parallel shift upwards or downwards for various illness profiles. Moreover, because those who report illness and subsequently seek care may be different from those who do not report illness at all, illness status may also have a slope effect, meaning that the coefficients in the demand model may differ for various illness profiles. This leads to a problem of sample selection (Greene, 1997; Millimet, 2001). Because sample selection and endogeneity are two different issues that require two different analytical approaches (Greene, 1997; Millimet, 2001), a clear identification and demarcation of the two is necessary to handle empirical data accordingly.

In this article, we report on a study which estimated the demand for medical care using a Heckman-type estimator to correct for bias due to sample selection (Heckman, 1979; Vandeven & Vanpraag, 1981). Our choice was motivated by the underlying assumption that the empirical data at our disposal was subject to sample selection (Greene, 1997; Heckman, 1979). In low and middle income countries in fact, household surveys collect information on health service utilisation conditional upon individuals reporting their illnesses. Empirically, the level of illness reporting in these countries is low, typically in the order of 8-25% (Akin et al., 1998; Gertler & van der Gaag, 1990; Hjortsberg, 2003; Rous & Hotchkiss, 2003; Sahn et al., 2003), suggesting that self-selection may occur at this level. That is to say that those who report illnesses and demand medical care are likely to have significantly different characteristics from those who do not (Pokhrel, 2007). Possible reasons for low illness reporting may include the level of health knowledge and education, income; the perceived disease severity; cultural norms and values, including gender

considerations; and the availability of health services within a community (Akin et al., 1998; Hjortsberg, 2003; Rous & Hotchkiss, 2003; Sahn et al., 2003)

The empirical literature has so far disregarded the potential impact of self-selection due to illness reporting on demand for medical care. Therefore, the aim of our study was to develop a model which could generate unbiased estimates and thus, allow us to formulate more adequate policy leads geared towards overcoming barriers to access and increasing health service utilisation. To illustrate our point, we first developed and estimated an empirical model for the demand for medical care taking into explicit account individual illness reporting. Afterwards, we compared the results of such model with those of an ordinary demand model that did not consider individual illness reporting.

We used data from rural Burkina Faso, precisely from the Nouna Health District (NHD), a region located in the North-West of the country, about 300 km from the capital Ouagadougou. Burkina Faso rates among the poorest countries in the world. At the time of the study, the per capita GDP was 1,100 US dollars (adjusted for purchasing power parity) (United Nations Development Program 2005). As elsewhere in sub-Saharan Africa, problems of access to care linked to geographical and cultural accessibility and to the poor quality of the services available, are exacerbated by the application of user fees. In particular, health services in the NHD suffer from extreme under-utilisation, with 0.16 contacts per capita per year (Secretariat General du Ministère de la Santé du Burkina Faso 2002). The local health system is organised in two levels: a series of first-line facilities located in the rural areas, Centre de Santé et Promotion Social (CSPS), and a District Hospital, located in the town of Nouna.

### *Factors affecting demand for medical care*

We begin our theoretical discussion with a construct which assumes that the demand for medical care is derived from the demand for health (Grossman, 1972; Jack, 1999). In this construct, it is assumed that individuals generate utility from consumption of commodities. Some of those commodities are purchased directly from the market. Medicines and diagnostic services can fall in this category. However, there are other types of commodities that are produced by the individuals through a process that combines their own time with other inputs which can be purchased in the market. 'Health' is such a commodity. Therefore, the demand for medical care is derived from the demand for health. Given the resources they command (both time and money), individuals decide to consume medical care to the extent to which it maximises their utility.

In order to translate this theoretical basis into our empirical analysis, we further note that individuals choose to seek medical care only once they perceive themselves to be ill (Pokhrel, 2007; Pokhrel & Sauerborn, 2004). It follows that illness awareness is the first step in one's health-restoration process. Moreover, the type and severity of the illness lead individuals to decide whether or not they are eventually willing to purchase medical care. Thus, the recognition of an illness and the decision to seek medical care are individual discrete sequential choices that can be analysed empirically.

Economic theory states that health care costs are one of the significant determinants of demand for medical care. A number of recent studies have empirically confirmed this theoretical postulation (Akin et al., 1998; Ching, 1995; Gertler & Hammer, 1997;

Gertler & van der Gaag, 1990; Pokhrel, Hidayat, Flessa & Sauerborn, 2005; Sahn et al., 2003; Sauerborn, Nougara & Latimer, 1994), although the overall scale of the effect has been shown to be modest, with important differences in costs responsiveness between different socio-economic and demographic groups. Most of the published studies, however, ignored illness-reporting behaviour and therefore, they might have produced biased elasticity estimates.

A number of other variables have been found to influence the demand for medical care. In societies where households allocate resources based on the productivity of the single household members, as it is the case in many African settings, age has appeared to be a significant determinant of the demand for medical care (Dong, Gbangou, De Allegri, Pokhrel & Sauerborn, 2008; Dong, Kouyate, Cairns, Mugisha & Sauerborn, 2003; Mugisha, Kouyate, Gbangou & Sauerborn, 2002; Sauerborn, Berman & Nougara, 1996). Furthermore, within the household, the one who controls resources and/or who is responsible for making household decisions is usually the one who defines who should receive medical care (De Allegri, Kouyate, Becher, Gbangou, Pokhrel, Sanon et al. 2006; De Allegri, Sanon, Bridges & Sauerborn, 2006; Dong, Kouyate, Snow, Mugisha & Sauerborn, 2003). This person is usually the household head and his/her characteristics, such as his/her age, sex, educational background and employment status, may determine the demand for medical care (De Allegri, Sanon & Sauerborn, 2006; Dong et al., 2008; Dong et al., 2003). On the supply side, the poor quality of health services is probably the most documented factor justifying the low-uptake of medical care in low and middle income countries (Acharya & Cleland, 2000; Akin, Griffin, Guilkey & Popkin, 1986; Chawla & Ellis, 2000; Hotchkiss et al., 2002; Mariko, 2003). Community characteristics, such as one's place of residence

(rural or urban) and its developmental status, including available health care services and financing arrangements to pay for them, are other determinants of the demand for medical care (Kroeger, 1983; Mugisha et al., 2002; Pokhrel, 2004). It is not clear, however, if an explicit consideration of illness reporting behaviour would alter the significance of these variables in predicting the demand for medical care.

## **Methodology**

### ***Data***

We used data from the Nouna Health District Household Survey (NHDHS), a survey routinely conducted in a sub-portion of the NHD under demographic surveillance and approved by the Ethics Committee of the Faculty of Medicine of the University of Heidelberg, Germany and the Nouna Ethics Committee, Burkina Faso (Würthwein R, Gbangou A, Kouyaté B, Mugisha F, Ye Y, Becher H et al. 2001).

Details of the survey are described elsewhere (De Allegri, Pokhrel, Becher, Dong, Mansmann, Kouyate et al. 2008). In brief, households were selected following a two-stage cluster sampling procedure, with each household holding the same probability of being selected. First, the area was divided into 33 clusters of approximately equal size and then 30 households were randomly selected within each cluster. A total of 990 households, of which 606 in the rural area and 384 in the town of Nouna, were selected for inclusion in the survey.

The NHDHS collects information, at least once a year, on each member of the household individually. The survey comprises information on demographic

characteristics, socio-economic conditions, self-reported morbidity, and health care seeking behaviour. We used data from the round of the survey conducted in May and June 2004. The survey was purposely conducted at the end of the dry season/very beginning of the rainy season. This is a time in the year when travel is limited and most household members can be found at home.

### ***Statistical methods***

In our analysis, we needed to control for a potential selection bias for the following reason. The NHDHS records information on medical care use only from those individuals who report having experienced an illness during the recall period (four weeks prior to the survey date). Thus, excluding individuals who do not report an illness from the analysis may generate biased estimates because individuals are likely to report an illness in a non-random way, as a result of variations which the survey cannot capture (i.e. due to unobserved heterogeneity). In other words, the sample of individuals who report an illness is likely to include only individuals who have a relatively high probability of illness perception, a phenomenon known as self- or sample-selection (Heckman, 1979; Vandeven & Vanpraag, 1981). As a result, a probit model of medical care choice which does not control for self-selection may yield estimates of demand for medical care that will be upward or downward biased, depending upon the direction of the sample selection bias. As the proportion of those who demand medical care given illness is much higher (83%) than the proportion of those who report themselves as ill (9%), we expect *a priori* an upward bias in an ordinary probit.



Let us assume that that  $\mathbf{CARE}_i^*$  denotes a latent variable that measures the propensity to seek care among ill individuals in the sample. Given the vector  $\mathbf{X}$ , which denotes individuals' demographic, socio-economic and other factors affecting their decision to choose medical care, we can write

$$\mathbf{CARE}_i^* = \mathbf{X}_i\boldsymbol{\beta} + u_{1i} \quad (1)$$

such that we observe only the binary outcome

$$\begin{aligned} \mathbf{CARE}_i &= 1 && \text{if } \mathbf{CARE}_i^* > 0; && \text{and} \\ \mathbf{CARE}_i &= 0 && \text{if } \mathbf{CARE}_i^* \leq 0 \end{aligned}$$

In equation (1),  $\boldsymbol{\beta}$  represents a set of parameters to be estimated and  $u_{1i}$  the error term.

However, we only observe  $\mathbf{CARE}_i$  for observations  $i$  if individuals report their illnesses. Let  $\mathbf{ILL}_i^*$  denote the propensity to report an illness given a set of underlying variables (the vector  $\mathbf{Z}$ ) with the following relationship

$$\mathbf{ILL}_i^* = \mathbf{Z}_i\boldsymbol{\lambda} + u_{2i} \quad (2)$$

where,  $\mathbf{ILL}_i = 1$  if  $\mathbf{ILL}_i^* > 0$ ; and  $\mathbf{ILL}_i = 0$  if  $\mathbf{ILL}_i^* \leq 0$ ;  $\boldsymbol{\lambda}$  represents a set of parameters to be estimated;  $u_{2i}$  the error term. Finally, let's assume that  $u_{1i}$  and  $u_{2i}$  are jointly normally distributed with correlation coefficient,  $\boldsymbol{\rho}$ . This allows maximum likelihood estimation of the log likelihood function

$$L = \sum_{i \in S}^{y_i=1} \ln[\Phi_2(\mathbf{X}_i\beta, Z_i\lambda, \rho)] + \sum_{i \in S}^{y_i=0} \ln[\Phi_2(-\mathbf{X}_i\beta, Z_i\lambda, \rho)] + \sum_{i \notin S} \ln[1 - \Phi_1(Z_i\lambda)] \quad (3)$$

where,  $S$  is the set of observations for which medical care is observed;  $\Phi_1(\cdot)$  is the standard cumulative normal, and  $\Phi_2(\cdot)$  is the cumulative bivariate normal distribution function. It is important to note that we will need identifying assumptions in the selection equation to identify the bivariate probit with sample selection (Greene, 1997).

### ***Dependent variable***

We constructed a model with the exclusive aim of assessing demand for formal medical care, meaning the biomedical care provided at the government health facilities in the area. Thus, our dependent variable includes individuals seeking care at a rural CSPA and individuals seeking care at the urban District hospital (which serves as a secondary care centre, but also as a primary care centre for urban residents).

### ***Explanatory variables***

We estimated a parsimonious model that includes several explanatory variables. Table 1 lists them and provides their distribution in the sample. Medical care costs and income variables deserve some explanations. As the household survey records the *ex post* medical care costs, we followed the standard method to create medical care costs variables (Ching, 1995; Gertler, Locay & Sanderson, 1987; Pokhrel, 2007; Pokhrel et al., 2005). This method considers medical care costs as a function of biological factors (i.e. age and sex), type and severity of illnesses, and market structure variables such as the population of the community [see Appendix A]. The self-selection term used in

this equation was generated after running a logistic model of provider choice (Dubin & McFadden, 1984; Lee, 1983). The medical care costs include out-of-pocket expenditure individuals paid for seeking care, including travel costs. Time costs (both travel time and waiting time) are not included in this measure as the survey does not record them consistently. As recommended by many authors (Deaton, 1998; Hjortsberg, 2003; Pokhrel et al., 2005; Rous & Hotchkiss, 2003), we used household six-month expenditure as a proxy of household socio-economic status.

### ***Instrumental variables***

The identification of the probit with sample selection depends on the identification of adequate instruments supposed to have no independent effect on the demand for medical care. As suggested by earlier studies in this area (Akin et al., 1998; Rous & Hotchkiss, 2003), we used housing and sanitation variables for this purpose. It is likely that these variables are correlated with wealth and thus with the ability to pay for medical services, but not with the choice to seek care *per se*. To verify whether our assumption was correct, and thus whether we were using the right instruments, we ran a probit model for medical care including, among other variables, housing and sanitation variables as regressors. We found no evidence of a significant independent correlation between the instruments and the demand for medical care.

### ***Tests for sample selection and equality of parameters***

In order to establish whether sample selection was important, we tested the hypothesis that the correlation coefficient,  $\rho$ , given by the sample selection model was indeed zero (StataCorp, 2001). If the correlation was zero, the sample selection model would be equivalent to the combination of the probit models for the demand for medical care

and for illness reporting. To confirm whether this was the case, we used the likelihood ratio (LR) test (StataCorp, 2001). Finally, to test whether the parameters estimated by the sample selection model were both consistent and efficient, we applied Hausman specification test (Hausman, 1978; StataCorp, 2001). Since one of the models did not contain information about illness reporting, we could simply not compare the height of the coefficients as they were subject to different normalization processes. Thus, we applied a procedure known as differential normalization, i.e. we forced one particular parameter to be 1 in both models and then performed the test for equality. We forced the parameter of the variable “married” to be 1 in both models and performed Hausman specification test to assess whether the sample selection model provided better parameter estimates than the ordinary probit.

## **Results**

Table 1 shows the descriptive statistics for the variables considered in this study. We interviewed 7668 individuals. The mean age of the sample was 23 years (range: 0-98) - about 46% were below the age of 16 years and 48% between 16-60 years. About 52% were male, 81% did not attend school, 39% were Marka, 49% indicated agriculture as their main occupation, 37% were currently married, and about 33% lived in the town of Nouna. Six-month household expenditure, as a proxy of socio-economic status, was 319,622 CFA (488 Euro).

There were a total of 667 episodes (or about 9%) of acute illness reported in the one month prior to the survey date. Thirty-seven individuals (less than 1%) reported more than one illness. Treatment was sought in 555 (83.2%) episodes. The majority (68.1%) of those seeking treatment, resorted to self-care, followed by those (24.7%)

who paid a visit to a formal medical facility. The average medical care costs for the single illness episode treated at a medical facility was 3957 CFA (6.04 Euro).

<insert Table 1 about here>

The estimated coefficients of the demand models and their standard errors are given in Table 2. Two model coefficients are presented - one without any correction for sample selection due to illness reporting behaviour and the other with the correction. The ordinary probit (without the correction) gave model coefficients which were on average 15% higher than those given by probit with sample selection, but the range was as wide as -93% to +77%. For instance, the coefficient of 60+ year age group was 0.543 (significant at the 5% level) in the ordinary probit, but it was insignificant and 77% lower (0.125) in the probit model correcting for sample selection. Likewise, the coefficient of medical care costs, although significant in both models, appeared to be 29% higher in the ordinary probit. Further, the significance of the effect of certain variables differed across the two models. Particularly, one's sex and whether or not they attended primary schools were less important in the model correcting for selection bias than in the ordinary probit.

<insert Table 2 about here>

The correlation coefficient ( $\rho = -0.73$ ) was found to be significant at the 5% level, confirming that sample selection bias did occur in our sample. This indicated that controlling for the probability of reporting an illness was critical to determine the

effects of background variables including socio-economic characteristics on the decision to seek medical care. Moreover, the likelihood ratio (LR) test confirmed that the decision to report the illness was not statistically independent of the decision to seek medical care ( $\text{Chi}^2=105.05$ ;  $p=0.00$ ). Finally, the Hausman specification test suggested that the parameters were both consistent and efficient in the model correcting for sample selection ( $\text{Chi}^2=0.73$ ;  $p=0.99$ ).

A few variables turned out to be significant predictors of the demand for medical care. As expected, the coefficient for the medical care costs had a negative sign indicating that an increase in medical care costs induced a reduction in the demand for medical care. As this variable was measured on a logarithmic scale, the magnitude of the effect (-0.167) cannot be interpreted directly. Figure 1 presents the point elasticity estimates derived from the model coefficients (setting the values of all other variables at their means). The figures suggest that a 10% increase in the medical care costs led on average to a 3.4% fall in its demand (Figure 1). The responsiveness of medical care costs varied by 25% between the poorest and the richest groups (-0.37 and -0.28 respectively). The predicted probability of medical care use given illness varied between 12-27% across income groups, the corresponding average and actual figures being 18% and 25%.

<insert Figure 1 about here>

The results also suggest that income, Bwaba ethnic origin, and residence in the town of Nouna were other significant predictors of the demand for medical care.

Individuals in the age group 16-60 years were more likely to seek medical care than children and the elderly.

The variables that were significantly associated with the likelihood of illness reporting were: age, household size, whether the household had running water, and whether water pots were always covered [Appendix B]. Elderly people were more likely to report an illness than children. Individuals living in large households (more than 5 members) were more likely to report their illnesses. Having running water available for drinking and having water pots always covered significantly reduced the likelihood of reporting an illness.

## **Discussion**

We estimated a health care demand function using data from rural Burkina Faso with a focus on examining how profound the implications of sample selection bias due to individual illness reporting behaviour were on such estimation. Two important concerns emerged as a result of our analysis. First, ignoring the possibility of sample selection bias due to illness reporting in a demand analysis can lead to wrong policy leads. This is in line with the conclusion appeared in a recent paper in this journal (Pokhrel, 2007). For instance, variables otherwise thought to be important from a policy perspective, such as sex and schooling in our case, might not predict health care demand once the bias due to illness-reporting is corrected for. Second, the effects of the predictors can be, as it was in our case, overestimated if selection bias is not taken into account. For instance, the 29% difference in the impact of the costs variable on the demand for medical care in rural Burkina Faso suggested that costs may be less of a deterrent to utilisation than what indicated by analyses not taking into account

selection bias, moderating the arguments in favour of removing user fees altogether and/or introducing heavy subsidies to reduce the costs of medical services for entire communities.

Our estimates revealed that although the costs of care was a significant predictor of demand, its responsiveness was moderate, with a 3.4% drop for every 10% increase in medical care costs - a close tally with several prior studies (Gertler & Hammer, 1997; Pokhrel et al., 2005). Our findings, however, contradicted decade-old findings from the same area (Sauerborn et al., 1994) in that the magnitude of our elasticity estimates was half of what had been reported earlier. Two differences in the analysis might explain this discrepancy. First, our estimates referred to point elasticity (at sample means) computed using a sample selection model as compared to arc elasticity computed using an ordinary logit in the prior study. Second, our sample considered all those who suffered from an *acute* illness in the one month prior to the survey date while the earlier study considered only those who reported being *severely* ill in the recall period.

The elasticity pattern we observed across different income groups, however, closely resembled the one reported in the earlier study. Even after a decade, the poorest people are the ones who are most likely to be deterred from using medical care because of its high costs. The fact that the elasticity pattern across income groups has remained the same can be explained in relation to the fact that no substantial changes in the socio-economic conditions of the region nor in the organisation of the local health care system have taken place in the decade between the two studies. The Kossi Province continues to be an area of prevailing poverty, but in spite of this, no



effective exemption mechanism have so far been set in place to allow the poorest to access medical services free of charge. Given that the epidemiological trend has also not changed substantially over a decade (with infectious diseases remaining the major cause of illness), the arguments put forward in the prior study on the negative externality attributable to infectious diseases and on the need to look for alternative financing mechanisms are as relevant today as they were one decade ago (Sauerborn et al., 1994). The cost elasticity pattern identified in our study calls for the introduction of targeted interventions to remove the financial barrier to access to care, through the introduction of either exemptions or heavy subsidies, only for those for whom the costs of medical care really constitutes an important deterrent to utilisation.

Our results also supported earlier reports claiming the existence of an age bias, but no sex bias in the intra-household allocation of resources for health care (Sauerborn et al., 1996). Individuals in the age group 16-60 years were in fact more likely than others to seek medical care, while no statistical difference was observed between men and women. This pattern is likely to be due to the fact that unlike South Asian societies, African societies do not tend to value men more than women (Pokhrel, Snow, Dong, Hidayat, Flessa & Sauerborn, 2005). In an attempt to protect a household socio-economic stability, however, African societies do tend to discriminate between individuals in their productive age as compared to children and the elderly (Sauerborn et al., 1996; Waddington & Enyimayew, 1989; Waddington & Enyimayew, 1990). In line with the argument outlined in the previous paragraph in favour of targeted interventions, governments should also ensure that limited resources are channelled towards facilitating access for those who currently lack it most. Waiving consultation fees for children under five represents the first important

step made by the government of Burkina, as by many other West African countries, in this direction. To ensure adequate access to care, however, additional efforts are needed to ensure that children and the elderly are exempted from all medical costs, not from consultation alone.

Our findings suggesting that the Bwaba were more likely to seek medical care were also consistent with earlier evidence from the study area. Both an earlier study on knowledge and attitudes towards HIV screening and one on the decision to participate in a community health insurance initiative (De Allegri et al., 2008; Sarker, Milkowski, Slinger, Gondos, Sanou, Kouyate et al. 2005) identified the Bwaba as a group holding better health knowledge and displaying a higher propensity to participate in health care related activities. However, we are unable to derive concrete policy recommendations on the basis of the results of our study. Unrevealing the reasons behind ethnic differences to understand how best to inform policy is beyond the reach of a quantitative analysis like ours and requires complementary qualitative inquiry to explore how social dynamics within specific communities influence health related behaviours, including the decision to seek medical care.

The fact that people living in Nouna town were more likely than others to seek medical services was also consistent with prior research from the study area (Mugisha et al., 2002) and can be explained intuitively in relation to their proximity to a better equipped health facility, i.e. the District hospital. Not only do people in Nouna town have to travel only a very limited distance to seek care, but they are also motivated to seek medical care given the wider range of services available at the local hospital. Since no gate-keeping mechanism is set in place for people paying out-of-

pocket fees, residents of the Nouna town can easily bypass the CSPPS and seek care directly at the district hospital. In addition, starting in 2004, explicit efforts coordinated by a joint initiative of the Nouna Health Research Centre, the University of Heidelberg, and the Burkinabè Ministry of Health have been channelled towards improving the quality of care of hospital services. Thus, our findings may have partially reflected the additional motivation to seek medical care induced by such evident improvements in service delivery.

## **Conclusion**

This paper shows that not correcting for sample selection bias in estimating health care demand can lead to an upward bias in model coefficients. Empirical analysis from rural Burkina Faso showed that this bias could be on average as high as 15% and that the range could be overwhelmingly large. This upward bias may lead to identifying false predictors of demand. After correcting for selection bias, the costs of medical services, household socio-economic status, ethnicity, and location of residence (rural/urban) were identified as significant predictors of demand for medical care in rural Burkina Faso. This calls for the development of adequate health policies which, through the introduction of targeted exemptions or subsidies, can facilitate access to care for those currently lacking it.

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