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The impact of health insurance for children under age 6 in Vietnam: A regression discontinuity approach

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ABSTRACT

Accessing health services at an early age is important to future health and life outcomes. Yet, little is currently known on the role of health insurance in facilitating access to care for children. Exploiting a regression discontinuity design made possible through a policy to provide health insurance to pre-school aged children in Vietnam, this paper evaluates the impact of health insurance on the health care utilization outcomes of children at the eligibility threshold of six years. Using three rounds of the Vietnam Household Living Standards Survey, the study finds a positive impact on inpatient and outpatient visits and no significant impact on expenditures per visit at public facilities. We find moderately high use of private outpatient services and no evidence of a switch from private to covered public facilities under insurance. Results suggest that adopting public health insurance programs for children under age 6 may be an important vehicle to improving service utilization in a low- and middle-income country context. Challenges remain in providing adequate protections from the costs and other barriers to care.

1. Introduction

The impact of public health insurance programs has received a great amount of attention due to its importance to universal health coverage, a goal common to many low- and middle-income countries (LMICs) (Lagomarsino et al., 2012). Universal health coverage strives to provide equal access to the health care services that people need without causing financial hardship (World Health Organization, 2010). This goal is particularly important for children for whom there exists a high elasticity of demand for health care (Ching, 1995; Leibowitz et al., 1985; Sauerborn et al., 1994) and the early intervention of health services has been shown to be critical to future health and life outcomes (Chen and Jin, 2010; Kraft et al., 2009). Universal health coverage may be achieved through different means; chief among these is the expansion of public health insurance programs. Little is known about the impact of public health insurance programs on the utilization outcomes for children in LMICs. Improving knowledge on the health seeking behaviors of children under insurance has important implications for policy makers in the design and financing of benefit packages. If benefit packages are aligned with population needs then health finance systems stand the best chance of meeting universal health coverage goals (World Health Organization, 2010).

Through pooling risk across individuals, health insurance lowers the price of health care at the time of purchase leading to increased use of health care (Zweifel and Manning, 2000). The effect on utilization, however, will depend upon the level of benefits provided and individual circumstances, alongside supply-side factors which affect levels of service utilization. Health seeking behaviors of younger children under insurance will conceivably differ to those of older children due to differences in underlying health status and health care needs, among other characteristics. This apparent lack of an adequate comparison group presents an evaluation challenge. Another concern is that the decision to purchase insurance may be linked to decisions around health status and health care usage. In this case, regression estimates of the impact of health insurance on service outcomes will be biased (Cutler and Zeckhauser, 2000).

In this study, we circumvent these problems through a regression discontinuity (RD) design, made possible by a 2005 policy in Vietnam to provide health insurance coverage to children under the age of six. The basic idea of the RD design is that children on either side of the cutoff age are similar and, in a sense, randomly assigned to insurance coverage. Under some plausible assumptions, differences in outcomes for children in the vicinity of the cutoff are attributed to differences in insurance coverage generating an

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unbiased estimate of treatment effects. RD design is generally regarded as having the greatest internal validity of the quasi-experimental estimators and requires relatively mild assumptions (Angrist and Pischke, 2009; Imbens and Lemieux, 2008; Lee and Lemieux, 2010).

Using three cross-sections of the Vietnam Household Living Standard Survey (2006, 2008 and 2010), and exploiting a discontinuity in insurance coverage induced by a policy intervention with a fuzzy RD design, we evaluate the 2005 policy across a range of utilization outcomes at public facilities. We also examine possible substitution effects between public and uncovered private service utilization and the impact on health expenditures. This paper makes a novel contribution to the scant literature on the impact of public health insurance programs for young children in a LMIC context. The following section provides background on child health insurance in LMICs with particular attention to Asia and Vietnam, as well as review of child health insurance program impacts. This is followed by separate descriptions of the data and methodology used in this study before a presentation of the results. The paper concludes with a discussion of results and final remarks for policy.

2. Background

2.1. Public health insurance expansion in LMICs and in Asia

Public insurance schemes typically comprise a collection of co-contributory, contributory and non-contributory schemes targeted to distinct population groups such as formal sector employees and civil servants, informal sector employees and students, and social beneficiary groups (Giedion et al., 2013). In recent years, many LMICs have expanded public health insurance coverage as a means to improving access to formal health care services for their populations, and providing financial protection from the costs of health care, on the path towards universal health coverage (World Health Organization, 2010). For instance, through a series of reforms, China tripled health insurance enrollment across the population from 30% in 2003 to over 90% in 2012 (Xiong et al., 2013); Ghana increased coverage from 35% to 66% from 2007 to 2010 through voluntary contributory insurance reforms (Odeyemi and Nixon, 2013); and most notably, Thailand reached 100% coverage by 2005 through its “30 Baht” reforms of 2001 (Gruber et al., 2014).

2.2. Child health insurance programs in LMICs and in Asia

Targeting specific population groups, such as children, the poor and persons with disabilities in social health protection programs is critical to health equity and has been highlighted as a priority area in the Post-2015 Development Agenda (Groce et al., 2011; Mitra, 2013; Odeyemi and Nixon, 2013; Petrera et al., 2013; Reinbold, 2011; United Nations, 2013). Targeting specific groups may be motivated by other considerations. Several studies have demonstrated that the elasticity of demand for health care may be highest for children and the poor in LMIC settings, implying that user-fees will deter these groups from using formal health care the most unless financial risk protection measures are taken (Asfaw et al., 2004; Gertler et al., 1987; Sauerborn et al., 1994). There also exists evidence, in the context of the United States and Burkina Faso, that medical demand for younger children may be more price sensitive than it is for older children (Leibowitz et al., 1985; Sauerborn et al., 1994).

Some LMICs have expanded health insurance by targeting certain vulnerable populations like children. This is for example the case in Guatemala, Argentina and Paraguay with programs for children and mothers, and Egypt with a program for school age children (World Bank, 2013; Yip and Berman, 2001). The arrival of child health insurance programs in Asia is a recent development. To our knowledge, only four countries have introduced child-targeted health insurance programs (India (state of Karnataka), the Philippines, Taiwan and Vietnam) and little is known on the impacts of these programs. The Indian state of Karnataka developed a state-funded health insurance scheme for school children in 2006 (Government of Karnataka, 2014). The scheme automatically provides health cards to children enrolled in grades 1–10 offering regular health check-ups and free care for various ailments. Prior to 1995, Taiwan had developed various health schemes, which left uninsured mostly children under 14 and adults over 65. When Taiwan created the National Health Insurance in 1995, its social insurance program which consolidated the previously established programs, it started to target coverage of certain benefits to previously uninsured children (e.g. free annual check-ups and immunizations to children) (Cheng, 2003). In the Philippines, as part of an expansion towards universal health insurance coverage, an insurance program was launched for school-aged children (Quimbo et al., 2011).

Vietnam's public health insurance journey can be traced back to 1986 when it introduced *Doi Moi*, a system of economic and political reforms that opened the socialist society to market forces (Lieberman and Wagstaff, 2009). In 1992, social health insurance was first introduced for civil servants and formal sector employees in a compulsory co-contributory scheme. The reforms also included a non-contributory scheme for ‘policy beneficiaries’ that included war heroes and other persons of merit in the socialist revolution. In 1994, a voluntary contributory insurance scheme was established initially for school aged children who by 2003 comprised over 99% of enrollees (Ekman et al., 2008). This proportion has reduced over time due to the gradual take up by farmers and the self-employed (71% in 2010, calculation by authors). In 2003, persons living in poor households became eligible to non-contributory health insurance through the Health Care for the Poor Program (non-contributory health insurance has been available to the poor since 1999 but was expanded more formally in effect from 2003 and later in 2006). In 2005, all children under the age of six years were added to the list of non-contributory beneficiaries (Socialist Republic of Vietnam, 2004, 2005). By the year 2010, some 61% of the Vietnamese population was insured (calculation by authors).

Fig. 1 plots insurance coverage for children, aged 0–10 years, over the years 2004–2010. The dramatic uptake for children under the age of six after the 2005 intervention illustrates that the policy

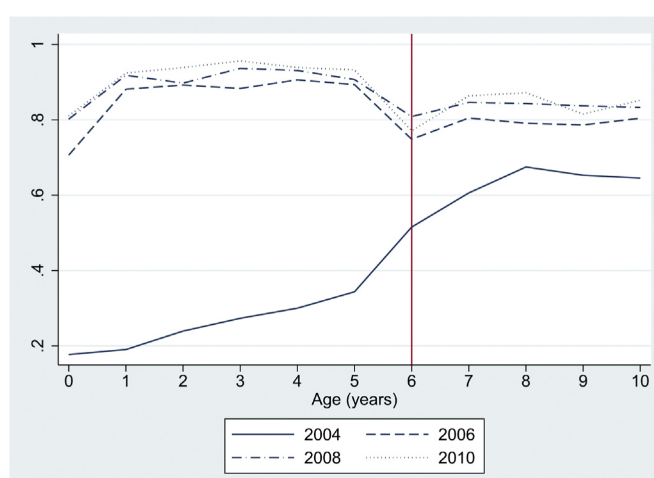

Fig. 1. Insurance coverage 2004–2010, aged 0–10 years.
Source: Vietnam Household Living Standards Survey 2004, 2006, 2008, 2010

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was rolled out effectively. By the year 2006, 83% of the target group under six years of age was issued with a free health insurance card, compared to just 26% in 2004. Prior to the national law on the protection, care and education of children that came into effect in 2005 (Socialist Republic of Vietnam, 2004, 2005), and for children six years of age and above, children were eligible for the voluntary (school) health insurance scheme or the poor household scheme. The rise in coverage for school aged children is most clearly illustrated in 2004, and is predicted in Appendix Fig. 2.

2.3. Child health outcomes and care utilization in Asia

The expansion of health insurance coverage to children in some countries in Asia through targeted or other programs has come in a background of less than optimal health and health care utilization indicators. Many LMICs in Asia have reduced their under-5 mortality rates since the 1970s. As shown in Appendix Table 1, Vietnam seemed to be above par with most Asian nations at 32 deaths per 1000 live births in 2000. The rate is lower than in most Asian countries (38 in Cambodia, 37 in China, 87 in India, 52 in Indonesia, 120 in Laos, 40 in the Philippines) but higher than in Thailand (23). From 1995 to 2003, in Vietnam, 33% of children under 5 suffered from being moderate to severely underweight; 36% and 6% of children under 5 suffered from moderate to severe stunting and wasting, respectively. These anthropometric figures fall in between the values reported for other Asian nations. For instance, for moderate to severely underweight, Vietnam's share at 33% falls below Cambodia (45%) and Laos (40%), but above China (10%), Indonesia (26%), the Philippines (31%), and Thailand (19%).

The picture on health care utilization for children under 5 prior to the 2005 reform is somewhat mixed. Using figures compiled in Appendix Table 1, Vietnam had immunization rates between 96% and 99% in 2004, and had reached immunization levels of high-income countries. During the 1998–2004 period, Vietnamese children under 5 with acute respiratory infections sought treatment from health care providers 71% of the time. Likewise, 71% of children under 5 suspected of having pneumonia were taken to a health care provider for treatment. These levels exceed those of nearby countries, but also highlight the fact that there is limited access to health care for nearly one third (29%) of children. Further, from 1996 to 2004, 39% of Vietnamese children under 5 with diarrhea received ORT with feeding, compared to 59% in Cambodia, 22% in India, 61% in Indonesia, 37% in Laos, and 76% in the Philippines.

Overall, available data for the early 2000s suggests that there remained potential for further improvements in the health outcomes and health care utilization levels of children under age 6 in Vietnam. The numbers point to a need to increase health care utilization among the child population and justify a study of the impact of health insurance expansion on health care utilization for children under age 6.

2.4. Impacts of child health insurance programs in Asia

Few studies have assessed the impact of child health insurance on health, health care utilization and other life outcomes, and especially for children under age 6. Ensuring access to quality health services is particularly important to children and has been found to contribute to positive health and educational outcomes. In the Philippines, through a difference-in-difference design, Kraft et al. (2009) found that a health insurance intervention targeted at poor children resulted in a reduction in the likelihood of wasting and having an infection. It was also found that delaying care by more than two days contributed to wasting and was less likely among the insured. In rural China, insured households were found to have better outcomes in child school enrollment and young child mortality, although the latter effect was removed once controlled for selection bias in a difference-in-difference with matching design (Chen and Jin, 2010).

Interestingly, most studies on health care utilization impact originate from Vietnam. The exception was a study from Indonesia that evaluates the impact of the financial crisis years (1997–2000) on child health care utilization using a difference-in-difference with matching design (Somanathan, 2008). The study found that utilization of public outpatient care declined for all children during the crisis years but by less for insured children with greatest protective effect for children aged 0–5 years.

With respect to Vietnam, Nguyen and Wang (2013) assessed the impact of the 2005 child health insurance reform in Vietnam using a difference-in-difference design and two waves of the Vietnam Household Living Standards Survey (VHLSS), 2004 and 2006. The authors compared utilization outcomes for children aged 6–7 to the treatment group of children younger than six years (divided into 0–3 years and 4–5 years). They found a positive impact of the policy on public inpatient and outpatient district hospital visits and a reduction in total health expenditures and the likelihood of catastrophic health expenditures among policy beneficiaries in the 4–5 years age range.

All other studies in Vietnam relate to older children and results are mixed. Using two waves of VHLSS panel data (2006 and 2008) and a fixed effects specification, Nguyen (2011) evaluated outpatient service utilization of children aged 6–14 years and found no impact of insurance on the number of visits and a negative impact on expenditures per visit. Other studies examined the voluntary insurance scheme more broadly which, as noted, consists largely of school age children. Using self-collected data from three provinces and an endogenous dummy variable model, Jowett et al. (2003) found that insurance significantly reduces average expenditures. By contrast, using VHLSS 2004–2006 data and a difference-in-difference with matching specification, Nguyen (2012) found no impact of voluntary insurance on out-of-pocket expenditures per visit, although he did find a significant increase in the number of outpatient and inpatient visits.

This paper provides an alternate estimation to the recently published study by Nguyen and Wang (2013). Whilst the adopted methodology of this earlier study sweeps out differences between the two groups that are fixed over time, as the authors note, there remains the problem of time-variant unobservable differences. Furthermore, results are limited to non-poor children only when there exists evidence that medical care demand may differ between poor and richer children (Ching, 1995). The authors use a pseudo panel specification resulting in small sample sizes which may affect the precision of their estimates, particularly those relating to inpatient admissions where the numbers of visits is low. To circumvent these various issues in estimation we apply an alternate fuzzy RD design to a large unrestricted sample of children. The caveat of our approach is that treatment effects are a weighted average at the discontinuity threshold.

Unlike the majority of papers, we disaggregate public and private visits to examine substitution effects. In Vietnam, like in many LMICs, insurance is often effective only at public facilities which consist of hospitals (district, provincial, central) and commune clinics. Private facilities can register with the national insurance agency. However, the proportion of registered private facilities is presently low: up to 2010, only 200 private health care facilities had registered with the national insurance agency versus 7600 public facilities (Nguyen, 2012). Prior to the child health insurance reform, the private sector was a strong service provider for young children (Nguyen et al., 2002). The afore-mentioned evaluation of the reform finds a positive, yet statistically insignificant, effect on private

Table 1
Description of outcome variables.

	0–5 years	6–10 years	Difference
	Mean (S.E)	Mean (S.E)	Mean (S.E)
Inpatient (public)			
Had a visit	0.082 (0.003)	0.044 (0.002)	0.038*** (0.004)
Number of visits	1.363 (0.034)	1.270 (0.037)	0.093* (0.054)
Expenditures per visit	472.667 (39.879)	611.332 (69.076)	−138.665* (74.584)
Outpatient (public)			
Had a visit	0.370 (0.005)	0.228 (0.004)	0.142*** (0.007)
Number of visits	2.843 (0.052)	2.413 (0.053)	0.430*** (0.008)
Expenditures per visit	47.676 (2.032)	52.419 (2.592)	−4.743 (3.303)
Outpatient (private)			
Had a visit	0.178 (0.004)	0.134 (0.004)	0.044*** (0.005)
Number of visits	3.928 (0.098)	3.012 (0.099)	0.916*** (0.143)
Expenditures per visit	58.995 (2.313)	56.425 (2.690)	2.57 (3.543)

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Standard errors in parenthesis. Expenditures are in Vietnamese dong and adjusted to 2006 prices using CPI. Number of visits and expenditures per visit estimates are for children with at least one visit.
Source: authors' calculation based on Vietnam Household Living Standards Survey 2006, 2008, 2010

outpatient service use (Nguyen and Wang, 2013). Similarly, no significant impact of insurance was found on the use of private sector services among children in Indonesia (Somanathan, 2008). These studies suggest that insurance is not associated with any substitution from private to public health care among child populations as is commonly reported in the adult population (Axelson et al., 2009; Pradhan et al., 2003).

Table 2
Description of independent variables.

	0–5 years	6–10 years	Difference
	Mean (S.E)	Mean (S.E)	Mean (S.E)
Insured	0.896 (0.003)	0.818 (0.004)	0.078*** (0.005)
Age (months)	36.147 (0.210)	102.724 (0.186)	−66.577*** (0.282)
Male	0.519 (0.005)	0.514 (0.005)	0.005 (0.007)
Kinh ethnicity	0.751 (0.004)	0.752 (0.005)	−0.001 (0.006)
Urban	0.236 (0.004)	0.212 (0.004)	0.024*** (0.006)
Household size	5.296 (0.019)	5.069 (0.018)	0.227*** (0.026)
Poor household	0.172 (0.004)	0.196 (0.004)	−0.024*** (0.006)
Age of household head	44.288 (0.150)	42.535 (0.133)	1.753*** (0.201)
Male household head	0.801 (0.004)	0.825 (0.004)	−0.025*** (0.006)
Married household head	0.865 (0.003)	0.892 (0.003)	−0.027*** (0.005)
Education of household head			
None	0.104 (0.003)	0.099 (0.003)	0.006 (0.004)
Less than primary	0.202 (0.004)	0.199 (0.004)	0.003 (0.006)
Primary	0.269 (0.005)	0.301 (0.005)	−0.033*** (0.007)
Lower secondary	0.251 (0.004)	0.256 (0.005)	−0.004 (0.006)
Upper secondary	0.125 (0.003)	0.109 (0.003)	0.016*** (0.005)
Above secondary	0.048 (0.002)	0.036 (0.002)	0.013*** (0.003)
Employment sector of household head			
Not employed	0.126 (0.003)	0.085 (0.003)	0.041*** (0.005)
Informal employment	0.656 (0.005)	0.684 (0.005)	−0.028*** (0.007)
Formal employment	0.059 (0.002)	0.050 (0.002)	0.009*** (0.003)
Government	0.160 (0.004)	0.181 (0.004)	−0.022*** (0.006)
Region			
Red River Delta	0.171 (0.004)	0.159 (0.004)	0.012** (0.005)
Northeast	0.146 (0.004)	0.148 (0.004)	−0.002 (0.005)
Northwest	0.086 (0.003)	0.066 (0.003)	0.020*** (0.004)
North Central Coast	0.098 (0.003)	0.112 (0.003)	0.014** (0.004)
South Central Coast	0.084 (0.003)	0.093 (0.003)	−0.010** (0.004)
Central Highlands	0.095 (0.003)	0.108 (0.003)	−0.013*** (0.004)
Southeast	0.121 (0.003)	0.126 (0.004)	−0.005 (0.005)
Mekong Delta	0.199 (0.004)	0.187 (0.004)	0.012*** (0.006)
Remote commune	0.256 (0.005)	0.259 (0.005)	−0.003 (0.007)
Observations	9588	8929	

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Standard errors in parenthesis.
Source: authors' calculation based on Vietnam Household Living Standards Survey 2006, 2008, 2010

3. Description of the data

Data for this study is drawn from three cross-sections of the VHLSS: 2006, 2008, and 2010. We limit the sample to children aged 0–10 years, yielding sample sizes of 6,042, 5,870, and 6605 for the respective three years ($N = 18{,}517$). The survey contains a rotating panel component with fifty percent of 2006 clusters retained for the 2008 round. Developed since 1992–93 from the World Bank living standards measurement survey (LSMS), the VHLSS collects a vast array of socio-economic and health information at the individual, household and commune level (e.g. General Statistics Office of Vietnam, 2010).

We draw heavily on the health care module which collects information on individual inpatient and outpatient visits and expenditures in the last twelve months prior to survey. Table 1 presents descriptive statistics of the outcome variables for children under age six, and aged six to ten years. Children less than six years of age were more likely to contact public health services and recorded a higher number of average visits when accessing services compared to older children. They were roughly twice as likely to record a public inpatient and outpatient visit. Approximately half the proportion of younger children with a public outpatient visit recorded a private outpatient facility visit, at a level slightly higher than older children. Expenditures per visit were one quarter lower at public inpatient facilities and similar at outpatient facilities among young children recording at least one visit compared to older children.

We draw upon other survey modules to control for a range of individual, household head, and commune level variables. Descriptive statistics of the variables are presented in Table 2 for the two age groups. Approximately 90% of younger children were insured compared to 82% of older children. With the exception of age, the distribution of the remaining variables was similar across the samples. Both age groups were evenly divided between male and female, and approximately one-quarter belonged to an ethnicity other than *Kinh* and lived in urban areas. Household size was around five and approximately one fifth of the households were classified as poor by local commune authorities. The vast majority of household heads were married males with a lower secondary school education or less and self-employed. One quarter of children surveyed were classified as living in remote communes.

4. Methodology

We adopt a RD design where assignment to treatment is determined (at least partly) by a cutoff of another 'forcing' variable (Angrist and Pischke, 2009; Imbens and Lemieux, 2008; Jacob et al., 2012; Lee and Lemieux, 2010). First introduced in 1960, RD design has recently taken a prominent place in the evaluation literature. Applications abound and include, for example, the use of test score, poverty index, and geographical boundary cutoffs to determine the effects of an intervention on outcomes (Lee and Lemieux, 2010). The design exploits the discontinuity in the forcing variable around the cutoff as a source of exogenous variation to identify a causal impact of the treatment at the threshold. A discontinuous jump in the outcome variable in the vicinity of the cutoff is attributed to the change in the level of treatment.

We apply a fuzzy RD (FRD) design where the probability of treatment increases by less than one at the cutoff. In our case, some children below the age of six are not formally enrolled in insurance and some children equal to or above the age of six are. For the analysis, we code all children under six as insured since under the implementation of the law, uninsured children under the age of six are entitled to fee exemptions at covered facilities upon the presentation of identification information (Socialist Republic of

Vietnam, 2005, 2008). Since the probability of treatment jumps by less than one at the threshold, the corresponding jump in outcomes can no longer be interpreted as the average treatment effect under the FRD design. Instead, it is recovered as the ratio of jumps in the estimated outcome and probability of treatment at the limit towards the cut-off point from both directions (Lee and Lemieux, 2010):

$$\tau = \frac{\lim_{c\to 6^-} E[Y|A=c] - \lim_{c\to 6^+} E[Y|A=c]}{\lim_{c\to 6^-} P[D=1|A=c] - \lim_{c\to 6^+} P[D=1|A=c]}$$

Estimation of the FRD can be performed with parametric or nonparametric approaches. We elect a parametric approach over a larger range of data primarily due to the low number of inpatient observations, but we use a non-parametric approach in sensitivity analyses in Section 5.1. Drawing strength upon observations further away from the cutoff, parametric approaches provide a more precise estimate of the treatment effect. Potential for incorrect estimation of the functional form increases with distance from the cutoff, leading to a bias in the treatment effect. Alternatively, a narrow interval near the cutoff associated with non-parametric estimations will reduce sample size and the precision of estimates (standard errors), especially when the population under study is limited to start with, as in this study of small children. Unless the underlying function is exactly linear there will also be bias associated with nonparametric procedures. Without knowing the true functional form, it is impossible to know which bias is the lesser. For further discussion on the merits of alternate estimators refer Lee and Lemieux (2010).

Following Hahn et al. (2001) who first drew the link between instrumental variables (IV) and *fuzzy* RD, average treatment effects under a parametric design are estimated by two-stage-least-squares with an indicator variable for the treatment group as the excluded instrument:

$$D_i = \gamma + \beta_1 T_i + \beta_2 (A_i - c) + \beta_3 (A_i - c)\cdot T_i + \beta_k \sum_k X_{ki} + \varepsilon_i \qquad (1)$$

$$Y_i = \alpha + \beta_4 D_i + \beta_5 (A_i - c) + \beta_6 (A_i - c)\cdot T_i + \beta'_k \sum_k X_{ki} + \varepsilon'_i \qquad (2)$$

where D is a binary indicator of insurance status for individual i, Y is the outcome variable, T is a binary indicator of the treatment group (age < 6 years), A denotes age (in months), c is the cutoff (72 months), X is a vector of independent variables, ε and ε' are random error terms. The coefficient, β_4, is identified as the local average treatment effect (LATE) of children who receive insurance (but would not otherwise receive it) at the age cutoff, in the limit. It is analogous to the general discontinuity gap ratio estimand (τ) in the exactly identified case and may be viewed as a weighted LATE with weights reflecting the *ex-ante* likelihood that the individual is near the cutoff (Lee and Lemieux, 2010). We use for Y three variables: a dummy variable indicating that the person had a visit, the number of visits, and the expenditures per visit.

In equations (1) and (2), with the third term of the right hand side, age is centered at the cutoff point to enable any shift at the cutoff point to be interpreted as a shift in the intercept, and is interacted with the treatment group dummy to allow for differences in slopes on either side of the cutoff. This is recommended in the case of nonlinear relationship between the outcome and rating variable or data far away from cutoff (Jacob et al., 2012).

The first stage equation in these models is estimated using ordinary least squares (OLS) regression. For the second stage, to the extent that levels of health care utilization, or the effects of insurance, differ with the age of children linear estimation of the treatment effect will be biased. Descriptive plots of outcomes against age suggest that the true functional form may be nonlinear. To limit bias associated with incorrect specification of the functional form, for the second stage, we apply nonlinear models appropriate to each of the outcome variable: Probit for the dummy variable indicating that the person had a visit, Poisson for the number of visits and Tobit for expenditures per visit. As recommended for parametric estimation of RD (Lee and Lemieux, 2010), we tested the functional form with the inclusion of polynomials on the forcing variable and determined a low-order polynomial with interaction with the treatment group indicator as the best approximation, as specified in equation (2).

Covariates were included in the models to reduce sampling variability and improve precision of the treatment effect estimates, as well as to address bias associated with observations that were not close to the cutoff (Imbens and Lemieux, 2008). Nguyen and Knowles (2010) analyze the demand for insurance by school aged and adolescent children (6–20 years) in Vietnam and found a strong socio-economic gradient at both the household and commune level. We thus include a collection of individual, household-head and geographical variables suspected as jointly correlated with the demand for insurance and health care (refer Table 2). Standard errors are estimated as robust 2SLS standard errors, as suggested by Imbens and Lemieux (2008), and all calculations were performed using Stata (Version 13).

Table 3
Probability model for being insured, aged 0–10 years.

	Coef. (S.E)
Treatment age (<72 months)	0.202*** (0.009)
Age (months)[a]	−0.001*** (0.000)
Age[a] × treatment age	0.001*** (0.000)
Male	−0.003 (0.004)
Kinh ethnicity	−0.021*** (0.007)
Urban	0.031*** (0.005)
Household size	0.001 (0.001)
Poor household	0.079*** (0.005)
Age of household head	0.001*** (0.000)
Male household head	−0.004 (0.007)
Married household head	0.014* (0.008)
Education of household head (reference category: none)	
Less than primary	−0.011 (0.007)
Primary	0.012* (0.007)
Lower secondary	0.015* (0.008)
Upper secondary	0.023** (0.009)
Above secondary	0.038*** (0.011)
Employment sector of household head (reference category: not employed)	
Informal employment	−0.010 (0.007)
Formal employment	0.023** (0.010)
Government	−0.012 (0.009)
Region (reference category: Red River Delta)	
Northeast	−0.004 (0.007)
Northwest	0.003 (0.009)
North Central Coast	−0.023*** (0.008)
South Central Coast	0.019** (0.008)
Central Highlands	−0.024*** (0.009)
Southeast	−0.025*** (0.008)
Mekong Delta	−0.053*** (0.007)
Remote commune	0.018*** (0.005)
Year (reference category: 2006)	
2008	0.024*** (0.005)
2010	0.024*** (0.005)
Constant	0.747*** (0.019)
Observations	18,517

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Ordinary least squares are used. Standard errors in parenthesis.
[a] Centered at 72 months.
Source: authors' calculation based on Vietnam Household Living Standards Survey 2006, 2008, 2010

Table 4
IV Regressions on public health care utilization outcomes, aged 0–10 years.

	Inpatient			Outpatient		
	Had a visit (1)	Number of visits (2)	Expenditures per visit (3)	Had a visit (1)	Number of visits (2)	Expenditures per visit (3)
Insurance[a]	0.068* (0.040)	1.133** (0.480)	28.421 (85.620)	0.217*** (0.065)	0.748*** (0.285)	4.325 (8.172)
Age (months)[b]	−0.000 (0.000)	−0.000 (0.003)	0.036 (0.346)	0.002*** (0.000)	0.006*** (0.001)	0.155*** (0.032)
Age[b] × treatment dummy	0.001*** (0.000)	0.014*** (0.004)	1.621*** (0.441)	0.000 (0.000)	0.004** (0.002)	−0.068 (0.043)
Male	0.013*** (0.004)	0.256*** (0.078)	24.194*** (8.146)	0.017** (0.007)	0.139*** (0.037)	2.636*** (0.810)
Kinh ethnicity	−0.011* (0.006)	−0.189 (0.126)	−8.368 (13.934)	−0.023** (0.011)	−0.043 (0.063)	1.854 (1.372)
Urban	−0.010* (0.005)	−0.280*** (0.108)	−15.396 (11.953)	−0.019** (0.009)	−0.103** (0.052)	0.403 (1.154)
Household size	−0.006*** (0.001)	−0.098*** (0.025)	−12.349*** (2.961)	−0.021*** (0.002)	−0.117*** (0.014)	−2.415*** (0.281)
Poor household	−0.000 (0.006)	−0.117 (0.109)	2.903 (12.923)	0.014 (0.011)	0.000 (0.056)	0.020 (1.276)
Age of household head	−0.000 (0.000)	−0.004 (0.004)	−0.983** (0.419)	−0.000 (0.000)	−0.001 (0.002)	−0.032 (0.041)
Male household head	−0.009 (0.006)	−0.324** (0.133)	−3.869 (13.782)	−0.019* (0.011)	−0.210*** (0.066)	−1.324 (1.455)
Married household head	−0.003 (0.008)	−0.033 (0.182)	−20.366 (17.479)	−0.009 (0.014)	0.009 (0.086)	−0.636 (1.801)
Education of household head (reference category: none)						
Less than primary	0.006 (0.007)	0.407** (0.160)	13.240 (15.351)	0.029** (0.013)	0.080 (0.078)	6.058*** (1.604)
Primary	0.010 (0.007)	0.421*** (0.141)	20.421 (15.114)	0.035*** (0.013)	0.082 (0.076)	6.868*** (1.598)
Lower secondary	0.003 (0.008)	0.315* (0.161)	6.795 (16.646)	0.039*** (0.014)	0.137* (0.082)	7.672*** (1.741)
Upper secondary	0.002 (0.009)	0.163 (0.180)	22.894 (19.852)	0.043*** (0.017)	0.113 (0.093)	9.446*** (2.070)
Above secondary	−0.013 (0.013)	−0.206 (0.259)	−13.484 (29.269)	0.040* (0.023)	0.143 (0.131)	16.039*** (3.188)
Employment sector of household head (reference category: not employed)						
Informal employment	0.006 (0.008)	0.049 (0.161)	−1.546 (18.009)	0.016 (0.013)	−0.033 (0.073)	−1.324 (1.633)
Formal employment	0.020* (0.012)	0.675*** (0.231)	19.218 (26.736)	0.023 (0.021)	0.019 (0.112)	−0.768 (2.801)
Government	0.003 (0.009)	0.145 (0.190)	−14.508 (20.969)	0.035** (0.015)	0.125 (0.088)	1.189 (1.881)
Region (reference category: Red River Delta)						
Northeast	0.005 (0.007)	0.097 (0.155)	−7.640 (16.020)	−0.005 (0.013)	−0.056 (0.067)	−8.250*** (1.666)
Northwest	0.003 (0.009)	0.028 (0.189)	20.005 (21.246)	−0.066*** (0.017)	−0.244*** (0.093)	−8.411*** (2.290)
North Central Coast	−0.006 (0.008)	−0.149 (0.163)	−23.598 (17.596)	−0.067*** (0.014)	−0.285*** (0.079)	−12.744*** (1.734)
South Central Coast	0.003 (0.008)	0.089 (0.184)	−17.794 (16.382)	0.016 (0.014)	0.222*** (0.073)	−6.473*** (1.692)
Central Highlands	−0.009 (0.008)	−0.107 (0.192)	−21.124 (18.024)	0.066*** (0.014)	0.443*** (0.080)	−0.312 (1.768)
Southeast	−0.001 (0.007)	−0.046 (0.160)	−19.708 (16.076)	0.052*** (0.013)	0.471*** (0.078)	5.184*** (1.729)
Mekong Delta	−0.003 (0.007)	0.137 (0.171)	−18.038 (15.177)	0.071*** (0.012)	0.735*** (0.060)	2.043 (1.445)
Remote commune	−0.009* (0.005)	−0.060 (0.118)	−22.466** (10.951)	−0.017* (0.009)	−0.113** (0.047)	−0.421 (1.092)
Year (reference category: 2006)						
2008	−0.007 (0.005)	−0.149 (0.097)	3.425 (10.202)	−0.036*** (0.008)	−0.182*** (0.049)	5.694*** (1.012)
2010	−0.001 (0.005)	0.064 (0.099)	−5.793 (10.114)	0.017** (0.008)	0.051 (0.046)	2.514** (1.051)
Observations	18,517	18,517	18,517	18,517	18,517	18,517

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Standard errors in parenthesis. Expenditures are in Vietnamese dong and adjusted to 2006 prices.
Results in columns (1), (2) and (3) are for probit, poisson and tobit regressions. Coefficients are marginal effects.
[a] Instrumented by treatment age dummy (<72 months).
[b] Centered at 72 months.
Source: authors' calculations based on Vietnam Household Living Standards Survey 2006, 2008, 2010

5. Results

The drop in the level of insurance coverage at age six induced by the 2005 policy in Vietnam is clearly depicted in Fig. 1, and remained when we plotted the expected probability of being insured as a function of age (Appendix Fig. 1). First-stage regression results are presented in Table 3. The treatment group indicator (age < 6 years) and instrument is strongly associated with the probability of being insured (20.2%), as needed for any IV specification. Other notable positive predictors include living in a poor household, urban area, remote commune, with an older, married, better educated household head, and being interviewed in more recent years. Variables that negatively affect the probability of being insured include higher age, belonging to the majority *Kinh* ethnic group, having a household head with less than primary education, and residing in central and southern regions.

Results for the second stage IV regressions on public facility outcomes are shown in Table 4. Insurance increases the probability of visiting an inpatient and outpatient service, as well as the number of visits. Insurance increases the probability of an inpatient visit by 6.8% and an outpatient visit by 21.7%; the average number of inpatient and outpatient visits increases by 1.13 and 0.75, respectively. The impact of insurance on expenditures is positive for both service types. However, the impact is not statistically significant. Higher predicted utilization levels among younger age children and discontinuity at the cutoff are illustrated in Appendix Figs. 3–5.

To examine possible substitution effects, we present second stage IV regressions for private outpatient utilization in Table 5. Insurance is positively associated with the probability of a visit, number of visits, and expenditures per visit. Weighted at the age threshold of six years, children experienced a 16.2% increase in the probability of a visit, an average 1.32 increase in the number of visits, and a 17,188 Vietnamese dong increase in expenditures per visit. As shown in Appendix Fig. 6, predicted levels of private utilization are lower than at public facilities.

5.1. Sensitivity testing

We run a number of sensitivity tests. All results from sensitivity-testing are available upon request.

As depicted in Appendix Figs. 3–6, models are not predicting observed utilization levels for children less than one year of age particularly well. Our causal treatment effect estimates are, in effect, a weighted average of observations closer to the eligibility threshold. It is possible that observations far from the threshold are influencing results. We test the sensitivity of the functional form by trimming observations far from the threshold. In turn, we exclude from the sample children less than one year of age, observations near the highest and lowest values of the age range (maximum and

Table 5
IV Regressions on private outpatient utilization outcomes, aged 0–10 years.

	Had a visit (1)	Number of visits (2)	Expenditures per visit (3)
Insurance[a]	0.162***	1.316***	17.188***
	(0.050)	(0.331)	(6.177)
Age (months)[b]	0.001***	0.006**	0.100***
	(0.000)	(0.003)	(0.025)
Age[b] × treatment dummy	−0.001***	−0.007	−0.131***
	(0.000)	(0.004)	(0.033)
Male	0.005	0.205***	0.996
	(0.005)	(0.073)	(0.610)
Kinh ethnicity	0.068***	0.802***	9.027***
	(0.010)	(0.128)	(1.185)
Urban	0.039***	0.421***	4.130***
	(0.007)	(0.085)	(0.767)
Household size	−0.010***	−0.150***	−1.009***
	(0.002)	(0.026)	(0.216)
Poor household	−0.031***	−0.217	−3.611***
	(0.009)	(0.153)	(1.133)
Age of household head	−0.000	0.001	−0.025
	(0.000)	(0.004)	(0.030)
Male household head	−0.002	−0.221*	−0.435
	(0.008)	(0.122)	(0.995)
Married household head	0.011	0.203	1.255
	(0.011)	(0.153)	(1.289)
Education of household head (reference category: none)			
Less than primary	−0.009	0.038	−0.440
	(0.011)	(0.157)	(1.218)
Primary	−0.004	0.184	0.141
	(0.011)	(0.165)	(1.228)
Lower secondary	−0.010	−0.007	0.088
	(0.012)	(0.167)	(1.295)
Upper secondary	−0.001	0.224	1.235
	(0.013)	(0.197)	(1.510)
Above secondary	0.021	0.288	5.042**
	(0.018)	(0.227)	(2.181)
Employment sector of household head (reference category: not employed)			
Informal employment	0.015	−0.018	0.399
	(0.010)	(0.124)	(1.156)
Formal employment	−0.017	−0.215	−1.943
	(0.016)	(0.205)	(1.992)
Government	0.012	0.076	0.412
	(0.012)	(0.151)	(1.398)
Region (reference category: Red River Delta)			
Northeast	−0.082***	−1.079***	−8.606***
	(0.012)	(0.152)	(1.579)
Northwest	−0.145***	−2.140***	−15.660***
	(0.021)	(0.313)	(2.825)
North Central Coast	−0.087***	−1.180***	−10.058***
	(0.012)	(0.158)	(1.630)
South Central Coast	0.065***	0.486***	7.386***
	(0.010)	(0.111)	(1.265)
Central Highlands	0.104***	1.000***	11.379***
	(0.011)	(0.135)	(1.341)
Southeast	0.115***	1.230***	12.002***
	(0.009)	(0.114)	(1.138)
Mekong Delta	0.149***	1.736***	13.660***
	(0.009)	(0.121)	(1.047)
Commune remote	−0.033***	−0.339***	−2.842***
	(0.008)	(0.091)	(0.916)
Year (reference category: 2006)			
2008	−0.021***	−0.104	−1.892**
	(0.007)	(0.103)	(0.789)
2010	−0.007	−0.057	−0.203
	(0.006)	(0.100)	(0.757)
Observations	18,517	18,517	18,517

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Standard errors in parenthesis. Expenditures are in Vietnamese dong and adjusted to 2006 prices.
Results in columns (1), (2) and (3) are for probit, poisson and tobit regressions. Coefficients are marginal effects.
[a] Instrumented by treatment age dummy (<72 months).
[b] Centered at 72 months.
Source: authors' calculations based on Vietnam Household Living Standards Survey 2006, 2008, 2010

minimum 1%, 5%, and 10%). Across the tests, estimates for expenditures per visit at public facilities become negative yet remain statistically insignificant. For utilization outcomes, estimates remain positive and are smaller in magnitude. However, they are no longer statistically significant. We furthermore applied non-parametric local linear regressions to smaller data bandwidths (36 months, 18 months, 9 months) and returned similar results to the trimmed parametric estimators above (Nichols, 2011). More limited statistical power is in line with what we would expect from smaller sample size estimations.

In addition, we run a placebo test: we explicitly test for a treatment effect before the reform with data for 2004. As shown below in Table 6, pre-reform, the treatment effect is negative for all outcomes, and often statistically significant, which indicates lower utilization outcomes for children in the treatment group (under age six). This is further shown in Appendix Table 2: the first stage regression coefficient on the treatment dummy is −0.201 and is significant at the one percent level, and there is a jump in the predicted probability of being insured at age six (Appendix Fig. 2). These negative pre-reform results with 2004 data seem logical given that a lot of children then accessed insurance through the voluntary contributory insurance scheme for school-aged children from age six. We further calculate the difference of the effects with post- and pre-treatment data (Table 6). The differences are positive, and statistically significant for selected outcomes: this result suggests that estimates with post-treatment data only in Tables 4 and 5 may be underestimates of the policy impact.

One obvious concern with our RD design is that something other than insurance status may be contributing to the jump in outcomes at the age threshold. It is plausible that young children are more prone to getting sick and, in turn, access health care at a higher rate than older children thus putting a downward bias on our results. Since we disaggregate age in months in the analysis, it is unlikely that factors will change sharply approaching the threshold. We confirm this by running local regressions of possible confounders, including measures of morbidity (days bedridden and days unable to perform regular activities) and household head characteristics (gender, education level and employment sector), on age (months) and a dummy variable indicating treatment age and found little or no statistically significant association on the interaction term. We found marginal significance on one morbidity measure (number of days unable to perform regular activities) for the largest bandwidth of 36 months.

Equally, it is important that no discontinuity in the average outcome exists at other values of the forcing variable. It is conceivable that the commencement of primary school at age 6 in Vietnam is associated with an outpatient health check-up or vaccination thus putting a downward bias on our results. However, birthdays are distributed throughout the school year so the rate of health care utilization should not change discontinuously with age in months. We confirm this in a reduced form regression of outcome variables on age (months) for the sample aged 6 years. We also ran local linear regressions centered at the age midpoints of the young and older child samples and found no evidence of discontinuity.

A further threat to validity is the extent to which individuals can manipulate the forcing variable around the eligibility threshold. Whilst this is not directly testable, the self-report of relatively more individuals with an age just below the age cutoff is suggestive of a violation of the no-manipulation assumption. We test this using a technique developed by McCrary (2008) which involves running local linear regressions on the number of individuals by age (months) on either side of the cutoff and testing whether the difference in the two intercepts is statistically different from zero. As expected, the t-test fails to reject continuity in the forcing variable.

Table 6
Difference in treatment effects pre- and post-reform, aged 0–10 years.

	Pre-reform		Post-reform		Difference	
	ATT	(S.E)	ATT	(S.E)	ATT	(S.E)
Inpatient (public)						
Had a visit	−0.152	(0.095)	0.068*	(0.040)	0.220**	(0.102)
Number of visits	−1.461	(1.504)	1.133**	(0.480)	2.594	(1.889)
Expenditures per visit	−324.459**	(163.811)	25.183	(95.855)	349.642	(220.111)
Outpatient (public)						
Had a visit	−0.230***	(0.086)	0.217***	(0.065)	0.447***	(0.102)
Number of visits	−4.136	(5.904)	0.748***	(0.285)	4.884	(7.388)
Expenditures per visit	−21.691*	(11.965)	4.955	(9.108)	26.646*	(13.346)
Outpatient (private)						
Had a visit	−0.167**	(0.080)	0.162***	(0.050)	0.329***	(0.097)
Number of visits	−2.702*	(0.080)	1.316***	(0.331)	4.188***	(0.169)
Expenditures per visit	−21.181***	(8.015)	19.394***	(7.122)	40.575***	(9.975)
Observations	6932		18,517			

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Standard errors are bootstrapped at 100 repetitions and in parenthesis.
Coefficients are marginal effects. Expenditures are in Vietnamese dong and adjusted to 2004 prices.
Pre-reform (2004) models are the same as post-reform (2006–2010) models with the exception of the count models which included additional polynomials on the rating variable and its interaction.
Source: authors' calculations based on Vietnam Household Living Standards Survey 2004, 2006, 2008, 2010

Finally, we test for efficiency bias resulting from the presence of panel observations in our pooled cross-sections (for 2006 and 2008) by restricting regressions to individual cross-sections and determine that serial correlation is not a significant problem.

6. Discussion

To the extent insurance uptake and health care utilization are chosen by households, estimates on the impact of insurance on outcomes are biased. To address problems associated with adverse selection we apply a RD design that exploits discontinuity of insurance coverage at an age threshold as a source of exogenous variation to identify the causal impact of insurance. Due to the existence of competing insurance programs, namely a voluntary health insurance program targeted to school aged children, discontinuity in coverage at the cutoff of six years was not dichotomous. We apply a fuzzy RD (FRD) which takes account factors other than the threshold rule which affect program participation. Following Hahn et al. (2001) who made the link between instrumental variables (IV) and FRD, we estimate treatment effects using two-stage least-squares. As in a standard IV framework, treatment effects are a local average treatment effect for individuals affected by the treatment. Being local, external validity of results is limited. The advantage is that they are robust to any confounding factors as long as they occurred in a continuous way around the age threshold. RD design is thus generally regarded as having a relatively high degree of internal validity compared to other non-experimental estimators.

We find that insurance has a positive impact on the probability of having at least one public inpatient visit and, to a greater extent, at least one outpatient visit, as well as on the number of visits. This result holds up to a number of sensitivity checks which overall add validity to our findings, although the statistical significance is lost while working with trimmed samples closer to the age cutoff or while using a non-parametric local linear regression of smaller bandwidths. This positive result on health care utilization is consistent with results of a difference-in-difference evaluation of the 2005 child health insurance reform (Nguyen and Wang, 2013) but stands in contrast with a study of older children (6–20 years) in Vietnam using a difference-in-difference specification, which finds no impact of insurance on visits (Nguyen, 2011). This relatively larger impact on utilization found for younger versus older children in the Vietnam health insurance literature is consistent with findings in studies beyond Vietnam, such as an assessment of the health card program in Indonesia (Somanathan, 2008), and those relating to the price elasticity of demand for health care for young children in the United States and Burkina Faso (Leibowitz et al., 1985; Sauerborn et al., 1994). To the extent that improved access to health care for younger children leads to improved health and life outcomes, this is an important finding for policy makers in Vietnam and in other LMICs.

At the same time, we find no significant impact of insurance on out of pocket health expenditures. This result is consistent with the result of Nguyen and Wang (2013) for children aged 0–3, but different from the negative and significant impact on expenditures they find for children aged 4–5. This difference in results for children aged 4–5 could come from the difference in sample and methodology: Nguyen and Wang (2013) use a sample of non-poor children only with two waves of the VHLSS (2004 and 2006) and with a difference in difference estimation, while this paper's estimates are for all children near the age cutoff of 6 using RD and based on 2006, 2008 and 2010 VHLSS data. Our finding that insurance has little or no impact on expenditures is consistent with other studies on health insurance in Vietnam (Nguyen, 2012; Palmer, 2014; Sepehri et al., 2011; Wagstaff, 2007) and in other LMICs (Escobar et al., 2010).

For the case of Vietnam, there are several plausible explanations for the lack of an impact. Waiting times and bureaucratic procedures for the insured at public facilities are substantial (Chau, 2013; Lam, 2012; Nghe, 2012; Tran, 2009). In addition, procedures for the insured in seeking care at higher level facilities are cumbersome and involve obtaining a written referral at each tier. Instead, patients (or their guardians) may elect to consult without using their insurance card. In our 2010 sample, around one quarter of insured children under six did not use their health insurance card when consulting at public facilities.

Whilst the list of reimbursable items under insurance is extensive, not all available procedures are covered and expenditure caps exist for covered items. The drugs offered are generic and hospitals routinely suffer drug shortages with insured patients having to purchase medication from private pharmacies at their own expense (Somanathan et al., 2013). Our measure of health expenditures included 'bonus' payments to doctors so it is likely that informal payments are contributing to the health expenditures of the insured. The measure was an aggregated total so we were unable to examine the extent to which informal payments were contributing to overall health expenditures. A culture of informal 'envelop' payments at health care facilities is commonly reported in Vietnam (e.g. Somanathan et al., 2013).

On the supply side, there exist incentives for doctors to supply services for the insured under a fee-for-service financing arrangement. Under a series of recent reforms designed to give hospitals more autonomy, fee-for-service financing was associated with increased hospital admissions, higher out-of-pocket spending, more laboratory tests and imaging per case for the insured in Vietnam (Wagstaff and Bales, 2012). Currently, Vietnam is transitioning towards a capped payment system for each individual enrolled with the provider. By imposing financial risks on providers, capitation has led to a reduction in the level of service provision for the insured at district hospitals (Nguyen et al., 2013).

It is possibly due to these barriers that a notable proportion of children under the age of six reported at least one visit at a private outpatient facility, which by and large are not registered with the national insurance agency. Like Nguyen and Wang (2013) and a study of a child health card in Indonesia (Somanathan, 2008), we find no evidence of a substitution from private to public care under insurance. It may also be that private services are simply more conveniently located or a kind of wealth or income effect associated with insurance is pushing households to consume more services, including more private services. The private sector is growing rapidly in Vietnam, the true extent of which is unknown since many providers have not registered with the Government (Lieberman and Wagstaff, 2009). In one province alone, there was almost twice the number of private than public providers (Tran et al., 2005). Our results suggest that the private sector remains an important source of health care for the insured in Vietnam. Consistent with other countries in Asia and Africa, continued efforts are required in developing the relationship between the private sector and the public health care system (Bloom et al., 2012).

7. Conclusion

This study evaluates a 2005 reform to provide fully subsidized health insurance to children under the age of six in Vietnam. Our review of the literature prior to the reform suggests that there was room for improvement in the health and health care utilization of children in Vietnam. Exploiting a discontinuity in insurance coverage induced by the 2005 reform, we investigate whether insurance might have played a role in improving access to care for children as well as whether it had any significant impact in reducing expenditures. Results overall indicate that the policy has been successful in improving access to outpatient and inpatient care. However, we could find no evidence of a substitution from uncovered private outpatient facility use or that the reform had any significant impact on health expenditures and thus on providing financial protection. Results support recent Government efforts to reduce cost and other barriers to entry into the public health care system and deepen the registration of private providers into the health insurance system. Further research is needed to identify if the reform led to improved health outcomes or improved access to specific care services. Although more research is needed on young children, broadly, the study results suggest that adopting public health insurance schemes for young children may be an effective way to boost health care utilization in a LMIC context.

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Appendix A. Supplementary data

Supplementary data related to this article can be found at http://dx.doi.org/10.1016/j.socscimed.2014.08.012.

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