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ABSTRACT

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We analyze child mortality in Vietnam focusing on gender aspects. Contrary to several other countries in the region, mortality rates for boys are substantially larger than for girls. A large rural-urban mortality difference exists, but much more so for boys than for girls. A higher education level of the mother reduces mortality risk, but the effect is stronger for girls than for boys.

JEL Classification: C13, C31, C35, C41, I12

Keywords: child mortality, gender differences, hazard rate, frailty model

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1 Introduction

Researchers in demography, public health, and population economics have documented the remarkably low child mortality levels in Vietnam compared to other countries in the same region with similar income levels; see e.g. Nguyen-Dinh and Feeny (1999) and Wagstaff and Nga (2002). Previous studies also reported large persistent differences between child mortality rates in rural versus urban areas. Recent work by Nguyen et al. (2007) and O'Donnell et al. (2009) and others suggest that these differences are related to rural-urban gaps in nutritional status, income, and parental education, in particular education of the mother.

While most researchers have allowed for possible gender differences in mortality rates by including a gender dummy in their empirical models, gender patterns have not been a central research question.[1] Wagstaff and Nga (2002) do not find a significant gender coefficient in a Weibull survival model estimated using the 1993 and 1998 Vietnam Living Standards Survey (VLSS). Dao (2006) and Nguyen-Dinh and Feeny (1999), on the other hand, report higher mortality rates for men at all ages.

The present paper scrutinizes gender differences in Vietnam's child mortality rates using a sample with information on close to 30,000 births. We use econometric models that control for unobserved heterogeneity at the family level and allow for parametric flexibility. We estimate models separately for boys and girls to allow the effects of covariates to be different across genders.

Our empirical results show some remarkable gender patterns. Contrary to several other countries in the region, we find a 30 percent higher overall mortality rate for boys. Secondly, we find a large rural-urban gap in mortality rates, with the gap being substantially larger for boys than for

[1] Nguyen et al. (2007) and O'Donnell et al. (2009) focus on urban-rural differences, but do not consider gender effects.

girls. Third, we find that maternal education has a strong effect on survival, as reported in much of the previous literature, but the effect on daughters appears to be stronger than the effect on sons.

The paper proceeds as follows. Section 2 reviews developments in the Vietnamese society and health care institutions that are relevant for understanding child mortality. Section 3 describes the data, while details of the econometric analysis are provided in section 4. Section 5 presents the empirical results, and section 6 concludes.

2 The Vietnamese context

After the reunification of the country in 1975, Vietnam witnessed a boom in the construction of health infrastructures and the training of health workers. By 1995, the Vietnam primary health care network covered 93% of the communes around the country. As Table 1 shows, in 1995 infant mortality rate was around 32‰, much lower than in Bangladesh (83‰), Indonesia (48‰), and China (37‰). Other health indicators also compared favorably with other countries in South-East Asia. With all health indicators showing strong improvements between 1995 and 2000 in all countries, Vietnam's retained its top ranking.

Under the central planning economy most national health services were provided free of charge or at highly subsidized prices. However, these circumstances changed profoundly upon the introduction of *Doi Moi* (reform) policy, which was introduced in 1986 and reinforced in 1989. While *Doi Moi* increased average living standards, it also increased inequality across social strata; see e.g. Wagstaff and Nga (2002). The early years of the transition to a market-oriented economy entailed a decline of health services in terms of health service delivery, equipment maintenance, and health workers' salaries. User fee charges for health care severely restricted access to health services for low-income households.

Facing the deterioration of the primary health care system and the rapid population growth in the late 1980s, the Vietnamese government formulated strategies to renew the health care system and family planning programs. Up to 2000, more than ten national health care programs and projects were implemented, including the leprosy elimination program (1982), National Tuberculosis Control (1986), National Malaria Control (1991), the Expanded Program of Immunization (1993), Acute Respiratory Infection control and the prevention of iodine deficiency (1996), AIDS prevention, a hospital equipment upgrading program (1996). One of the effects of these programs was a strong increase in vaccination rates. The time line in figure 1 summarizes the main events and programs.

Goodkind (1995) mentions a study that suggests the existence of a boy preference in Vietnam, based on stated preferences. Any possible effect on the male/female birth ratio might have been reinforced by a birth control policy implemented in 1998 that calls for a maximum of "one or two children". However, as shown in the next section, the fraction of male births in our sample is 0.519, very close to the natural male birth probability.

3 Data and descriptive statistics

Most of the studies on child mortality in Vietnam have used the 1988 Vietnam Demographic and Health Survey (VNDHS), the 1990 Vietnam Accessibility of Contraceptives Survey (with data on children born between 1983 and 1990), and the 1993 and 1998 Vietnam Living Standards Survey (VLSS).

Our empirical analysis is based on more recent versions of the VNDHS: 1997 and 2002.[2] Both surveys are nationally-representative household surveys. Data were collected at the level of individual ever-married women at

[2]The surveys were conducted by The National Committee for Population, Family and Children - NCFPFC (formerly NCFPFP) and the General Statistics Office (GSO), with technical assistance from ORC Macro, a professional service firm in Washington, DC, USA.

reproductive ages (15-49 years old; 5,664 in 1997 and 5,665 in 2002), and at the community level. The main objectives of the surveys were to collect detailed information on fertility, family planning, maternal and child health and survival indicators, as well as personal characteristics of the women such as their age, education, place of residence, working status, and their housing facilities. In VNDHSs, survey sites covered 41 provinces throughout the country. Although the repeated survey in 2002 returned to the same sample points (communities) as in the 1997 survey, the survey is not a panel at the level of individuals of households. However, it does constitute an (unbalanced) panel at the level of communities: we have data on 202 and 203 communities in 1997 and 2002, respectively, with 164 communities being surveyed in both years.

While information on socio-economic factors refers to conditions at the survey time, information on birth history and child survival status is retrospective. Clearly, data derived from retrospective reporting could be subject to (time varying) measurement error. It should also be noted that our data is not a representative sample of the cohorts born between 1961 and 2002. The mothers with the earliest birth years in our data are those who were 49 in 1997, i.e with birth years 1947 or 1948. This implies that the children in our sample from birth year 1961 had very young mothers. This is likely to somewhat overestimate the mortality rates in the earlier years in figure 2. In our econometric analysis, this problem is less of a concern, as we will control for the mother's age at birth when explaining mortality. We will perform a robustness check by reestimating the models using children born to mothers from younger cohorts only.

From the complete birth history of sampled women, we have information on 29,900 live births to 10,734 mothers (there are 595 women who do not report any live births). The year of birth of the children spans from 1961 to 2002. During this relatively long time span Vietnam has witnessed profound changes in terms of its political system, socio-economic conditions, and health services.[3] Table 2 shows the distribution of mothers by number of births and deaths. On average, a mother gives 2.79 live births. Almost half of the mothers have 3 or more children and about 7% have more than five. The 1,721 children who died account for 5.8% of the live births. The number of deaths is incurred by only 13% of the mothers. Only about 2.5% of mothers have experienced the loss of more than one child, but these deaths account for about 37% of total deaths. So, there is clear evidence of clustering of deaths within households. This may be related to observed as well as unobserved mother-specific factors that induce a correlation between survival probabilities of the children of a given mother (sometimes referred to as 'frailty').

Figure 1 graphs the developments of the overall child mortality rate in our data, and shows a strong and persistent decline over the past decades. It also shows that mortality rates for boys have been higher in almost all years. This contrasts the pattern in several other countries in Asia, in particular India and China, in which girl mortality rates boy mortality rates; see for example van der Klaauw and Wang (2011) and www.who.org. Figure 2 plots the ratio of the rural to urban mortality rates, for boys and girls separately. The rural-urban ratios exceed 1 for both genders in virtually all years, in line with earlier findings. According to the present data set mortality rates are about 75 percent higher in rural compared to urban Vietnam. This difference is large. For example, Poel et al. (2009), who analyze a cross-section of Western Sub-Saharan countries, find an average difference of 47 percent. Second, the rural-urban differences for boys appear to be substantially larger for boys than for girls, in some year even 5 to 10 times larger.

[3] We have opted not to use variables regarding housing conditions. These only describe the household situation at the time of the interview. Given Vietnam economic growth during the past decades these conditions are likely to have changed considerably during a mother's child bearing life.

Table 3 reports additional descriptive statistics, while figure 4 shows the kernel-smoothed baseline hazard curve over survival time. In general, children suffer from high risk of mortality early after birth, especially during the first month that represents more than 40% of total deaths. The hazard rate then declines sharply over time and then becomes small and more or less constant after the fifth year.

4 Empirical Methods

Mosley and Chen (1984) proposed an analytical framework for examining determinants of child mortality in developing countries that has been used widely. Their framework postulates that the effects of socio-economic factors on child mortality operate through a common set of biological mechanisms, or proximate determinants which are categorized into five groups: *i*) maternal factors including fertility behaviors, *ii*) environmental health conditions, *iii*) nutritional status, *iv*) injury and *v*) personal illness control. Although the framework is often referenced, only a few studies control for all the proximate determinants due to the limitations of the data. Most studies examine the effect of fertility-related behavior such as birth interval and breastfeeding duration **?** while others explore the causal relationship between socio-economic factors and child mortality **?**. The explanatory variables in the present study are motivated by the Mosley-Chen framework as far as data availability permits.[4]

Our key econometric tool for investigating the determinants of child mortality is a proportional hazard model with a semi-parametric piecewise-constant hazard rate for the intervals 0-1, 1-6, 6-12, 12-24, 24-60, and > 60 months (c.f. Greene (2003)).[5]

[4]Given the absence of income information in the VNDHSs, we experimented with dummy variables for television and motorbike ownership as proxies for household wealth. The dummies were insignificant in preliminary analyses and dropped henceforth.

[5]For children who are younger than one month we take age equal to half a month. The results are very similar to those from a model that uses $\Pr(0 < t < 1)$. For children

Children born to the same mothers are likely to share unobserved mother-specific biological and genetic factors, and unobserved household socio-economic conditions. We therefore allow for a potential correlation between the hazards of sibling by including a mother-specific random effect ('frailty') parameter. The hazard function for child i of mother m with I_m children is specified as:

$$h(t_i|v_m) = v_m h_0(t)\exp(\beta' \mathbf{X_i}), \quad i = 1, \ldots, I_m. \tag{1}$$

Following Vaupel et al. (1979), we assume v_m to follow a Gamma distribution with unit mean and a finite variance σ^2. Abbring and van den Berg (2007) show that in a large class of hazard models with proportional unobserved heterogeneity, the distribution of unobserved heterogeneity converges to a Gamma distribution. Gutierrez (2002) provide details on estimation. The null hypothesis that $v = 0$ can be tested using a boundary-value likelihood-ratio test based on a mixture of $\overline{\chi}^2_{01}$ distributions; see Gutierrez et al. (2001). Clayton (1978) and Guo and Rodriguez (1992) noted an interesting interpretation of σ^2. The model implies that the ratio of the hazard of a child at t_1 given a number of child deaths in the family, to the hypothetical hazard of the same child given no deaths in the family, is given by:

$$\frac{h_i(t_1|t_2, \ldots, t_{I_m}; \delta_2, \ldots, \delta_{I_m})}{h_i(t_1|t_2, \ldots, t_{I_m}; 0, \ldots, 0)} = 1 + \sigma^2 \sum_{j=1}^{I_m} \delta_j \tag{2}$$

where δ_j equals 1 if child j died at t_j and 0, otherwise. Hence, the hazard rate of dying increases by σ^2 multiplied by the number of deaths. This expression provides a measure of the size of the mother-specific unobserved effects. [6]

who die within one month after birth, the original dataset gives the time at death in days which is converted into months by dividing the time by 30. For living children, the variable survival time represents (censored) age in months at the time of the survey. For children who died before the survey, the (uncensored) survival time is defined as age at the time of death.

[6] Van der Klaauw and Wang (2011) also allow for child-specific frailty, in addition to

We estimate these models on the total sample, as well as separately for boys and girls.

5 Estimation results

Table 4 presents the estimation results of the hazard rate models with and without unobserved heterogeneity for the total sample of boys and girls. First of all, we find a very strong and highly significant gender effect. The coefficient implies that for a boys the probability of dying before a particular age is 30 percent larger than for girls. While excess girl mortality has been found for a number of Asian countries, in particular India and China, excess male mortality rates (at all ages, including the youngest) is found in almost any society around the globe; see e.g. www.who.org. The estimated gender coefficient in the Weibull model estimated by Wagstaff and Nga (2002) using the 1993 and 1998 VLSS also suggests excess boy mortality, but their coefficient is insignificant. This could be related to their smaller sample size or to the fact that they use older data (and older cohorts). Dao (2006) and Nguyen-Dinh and Feeny (1999), on the other hand, report higher mortality rates for men at all ages, as reported here. A similar result for another country in the Asian Pacific region, the Philippines, is reported by Guilkey and Riphahn (1998).

In subsection 5.2 we investigate this gender difference in more detail. We now first discuss the effects of the other coefficients.

5.1 The determinants of child mortality

Sibling correlations

A mother specific random effect reflects a correlation between the risk of death among siblings and family death clustering. Given the estimate of

frailty at the family level. However, their empirical results do not show any significant and substantial child-specific frailty.

σ^2, each death of a child in a family increases the risk of the index child by 56%, other things equal. This large effect may come from various factors like genetic factors, illness control behavior, parental competence, and household economic status. The estimate of σ^2 varies widely across studies. While Guo (1993) estimates 0.14, using the Institute for Nutrition in Central America and Panama Survey data, Omariba et al. (2007) using the 1998 Kenya Demographic and Health Survey find a value of about 0.98.[7]

Although the estimate of σ^2 in the model with unobserved heterogeneity is large (0.58) and highly significant, the differences between the two sets of other coefficients are small. Below we discuss the results for the model with unobserved heterogeneity.

Time effects

Table 5 presents the proportional effect of the calendar and survival time on the hazard rate. The result in the first column, first row reflects the raw data, in which about 40 percent of deaths occur during the first months after birth; cf. Lawn et al. (2005). Figure 2 already showed a strong declining time trend of mortality as a function of calendar time. The estimated coefficients on calendar time in table 4 indicate that this strong decline remains after controlling for observed explanatory variables. This indicates a substantial role for unobserved factors that are changing over time and not included in the model, like household income, technical advances in health care services, and learning effects.

Birth order and fertility behavior

Birth order has a strong effect on child mortality. Estimated coefficients

[7]From table 2, it is evident that there are only a few families with very high mortality. If we exclude all families with five or more deaths, the estimated σ^2 drops to 0.465, while all other estimates remain virtually unchanged.

of the first order birth and high order birth variables are statistically significant and positive and imply a U-shaped effect of birth order on child mortality. Given these estimation results, the risk for the first birth and higher order birth are 36% higher than that for the other order births, respectively. These findings are similar to those found in previous studies in Vietnam; see Swenson et al. (1993), Swenson et al. (1995), and Nguyen-Dinh and Feeny (1999). The high risk for the first births is caused by generally more difficult deliveries at first births (*dystocia*) and the lack of experience with child-bearing and rearing. The lower risk for the later births may result from learning effects among mothers. The increased risk for higher order births might be related to maternal depletion and sibling competition with resources. Moreover, higher birth orders are more likely to occur in poor families with limited resources Horton (1988).

Mother's age and birth intervals

The literature on the effects of mother's age on child mortality shows mixed results. Swenson et al. (1993) and Nguyen-Dinh and Feeny (1999), using a smaller sample for Vietnam (one fifth of the current sample size) for the birth years 1979-1988 find that mother's age at birth does not contribute to increased mortality risk of the children. However, others find signifcant effects for other Asian countries such as India (Nath et al., 1994) and Sri Lanka (Trussell and Hammerslough, 1983). The present results imply that a one year increase in the mother's age at first birth increases the risk of death for her children by about 3%. Makepeace and Pal (2006) use this variable and literacy as proxies for income and wealth of a household and also find the risk for the first birth is significantly reduced with the age at first birth. The coefficient for older mothers (over 35 years old) is not significant.

A short preceding birth interval is an important determinant of child mortality. The estimated coefficient is statistically significant and positive, and substantial in magnitude. Other things equal, the mortality risk for non-first birth with less than 24 months birth spacing is as much as 78.5% higher than that for others. The effects of short birth interval are also likely to be related to maternal depletion and sibling competition. While the results are consistent with most of the existing literature, they differ from findings in previous studies on Vietnam that document no significant relationship between child spacing and mortality (Swenson et al., 1993; Nguyen-Dinh and Feeny, 1999). The difference might be related to the profound changes in the socio-economic context during the transition in Vietnam initiated from the early 1990s. The shift from the centrally planning economy to the market economy increased working pressures on women and reduced subsidies related to childbearing and childrearing. These changes might have contributed to maternal depletion and sibling competition.[8]

Socioeconomic factors

It has been well-documented that better maternal education is associated with lower child mortality risk due to better feeding practices, illness controls and child-care (Caldwell and McDonald, 1982; Cochrane et al., 1982; Hobcraft et al., 1984; Majumder and Islam, 1993; Wagstaff and Nga, 2002). Our findings confirm this for Vietnam. The estimated coefficients are positive for illiterate mothers and negative for well-educated mothers, and highly significant. That is, the risk is highest for children born to mothers without education, moderate for mothers with elementary and secondary education, and lowest for mothers with higher education. Taking the second group of mothers as the reference group, the risk suffered by children born to the first group is 52% higher while that of those born to the third group is about 45% lower, other things equal. These estimation results somewhat differ

[8] It should be noted, though, that in the absence of a simultaneous model that explains both child mortality and birth spacing, the present results (like those in most of the existing literature) are likely to overestimate the causal effects of birth spacing on child mortality.

from findings in previous studies on Vietnam. Swenson et al. (1993),Swenson et al. (1995),Anh (1995) and Nguyen-Dinh and Feeny (1999) find that basic education plays a dominant role in reducing child mortality rate and there is no significant difference in early child mortality between the two higher education levels.

Ethnicity has a significant effect on child mortality risk. The probability that a Vietnamese child dies is 32% smaller than the probability for a child from an ethnic minority. Most of the ethnic minorities live in rural, remote and disadvantaged areas, and face less favorable economic conditions than their Vietnamese peers, even conditional on residence and education.

Regional effects

Rural children suffer a significantly higher risk of mortality than urban ones. The rural-urban difference in the risk is 37%. With a difference of 77% in the raw data this indicates that half of the gap is attributable to differences in observable characteristics like education level and fertility behaviors.

Estimated coefficients of regional variables are all negative, showing the lower risk for the children in these regions relative to that in the reference region (the Mekong River Delta). The results point at the presence of unobserved region-specific effects, and might be related to differences in natural environment and socio-economic development levels.

5.2 A closer look at gender patterns

To better understand the gender difference in child mortality we have estimated the model separately for boys and girls; see table 6. We present the no-frailty results as the coefficients on the explanatory variables are very similar between the frailty and no-frailty versions of the model. For convenience we repeat the results for the total sample from table 4. A likelihood ratio test shows that the sets of estimates obtained for the two genders are statistically different ($p < 0.001$).

Besides the overall higher mortality rate for boys, there are two covariates with notably different effects for boys and girls. First, a high education level of the mother has a strong negative effect on the mortality hazard rate of girls ($p = 0.056$), but its effect is insignificant for boys. The protective effect of high maternal education might be stronger for girls if the bond between mothers and daughters is tighter than between mothers and sons. Some evidence of the bonding difference is suggested by the shorter breast-feeding duration for boys; see Pham (2009). Second, the dummy for rural areas is highly significant for boys, but insignificant for girls. The study by Huong et al. (2006) on the causes of premature mortality in rural Vietnam might help to understand this pattern. Based on verbal autopsy information their results indicate that traffic accidents and drowning are leading causes of death of male children and adults in the rural sites studied. The rural areas are characterized by many ponds, lakes, canals and rivers. If boys have a large radius of action, farther away from home, possibly with less parental supervision, a rural area might pose more threats to boys than to girls. Moreover, several studies report that boys are more susceptible to diarrhea and respiratory infection; see e.g. Adair et al. (1993), Cebu Study Team (1992), and Pham (2009). These diseases are more prevalent in rural areas, and are sometimes a precursor to death.[9] Finally, some of the larger boy mortality could have resulted from war-related violence; see Merli (1998).

We have reestimated the models using more recent cohorts only, dropping from the sample children born before 1986. The estimated coefficients only

[9] The Vietnam DHS data in principle also allows for an analysis of morbidity. A difficulty - besides a smaller sample size - is a lack of instrumental variables to control for the endogeneity of health inputs, like vaccinations and the use of contraceptives. As a consequence, questions about the effectiveness of programs like PFHP remain difficult to answer. The best approach with respect to future programs is to have them preceded by small-scale pilots in which program communities or provinces are randomly selected. Such a procedure allows for a sound scientific evaluation of program effectiveness, and thus for an equitable allocation of resources such that the living conditions of the most vulnerable children are improved.

change slightly, suggesting that possible biases resulting from recollection error and sample non-representitiveness are small.

6 Conclusion

Scholars in various disciplines have documented the relatively low child mortality levels in Vietnam, as well as its strong decrease during the past decades. Large urban-rural differences have been reported as well. Focusing on gender differences in mortality patterns, we find that Vietnam conforms to the usual pattern found almost anywhere else in the world (contrary to some other Asian countries): excess boy mortality. We also found two other gender patterns that have not been the focus of research in the literature so far. First, the rural-urban mortality gap is much larger for boys than for girls. Second, mother's education is more protective for girls than for boys. We have suggested that gender specific patterns of social interactions within the family could be a part of the explanation. Additional data, preferably with information on the intrahousehold allocation of time and other resources, is a prerequisite for moving from suggested explanations to firm answers.

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Figure 1: **Time line of relevant events during the transitional period**

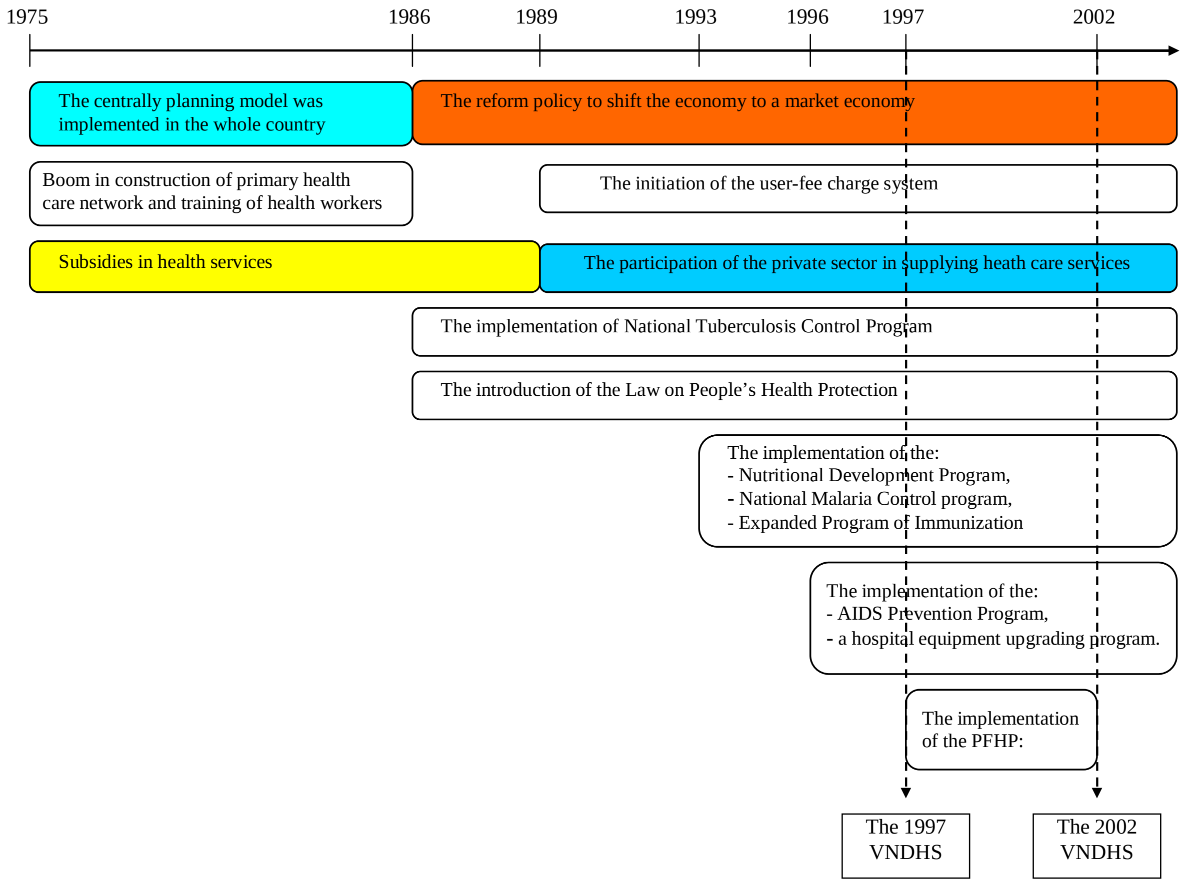
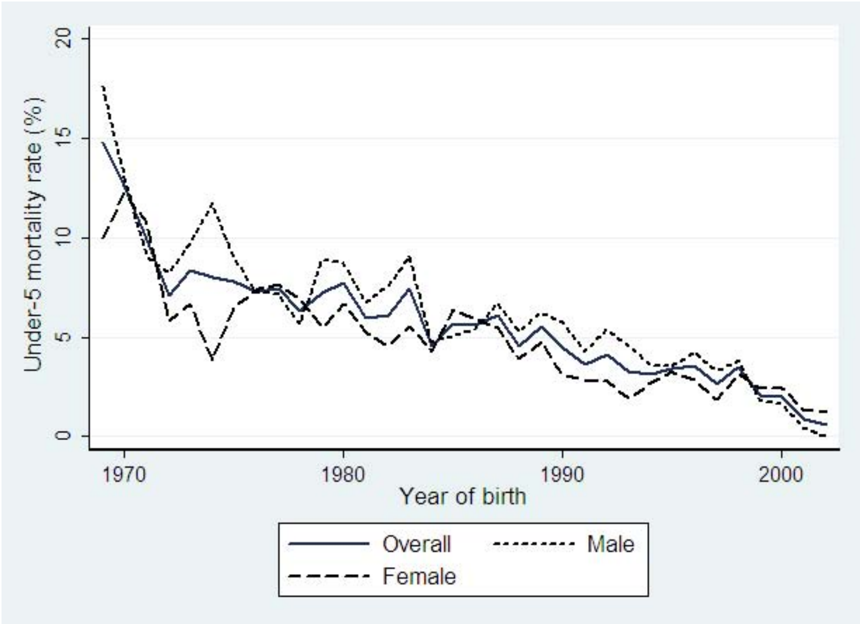
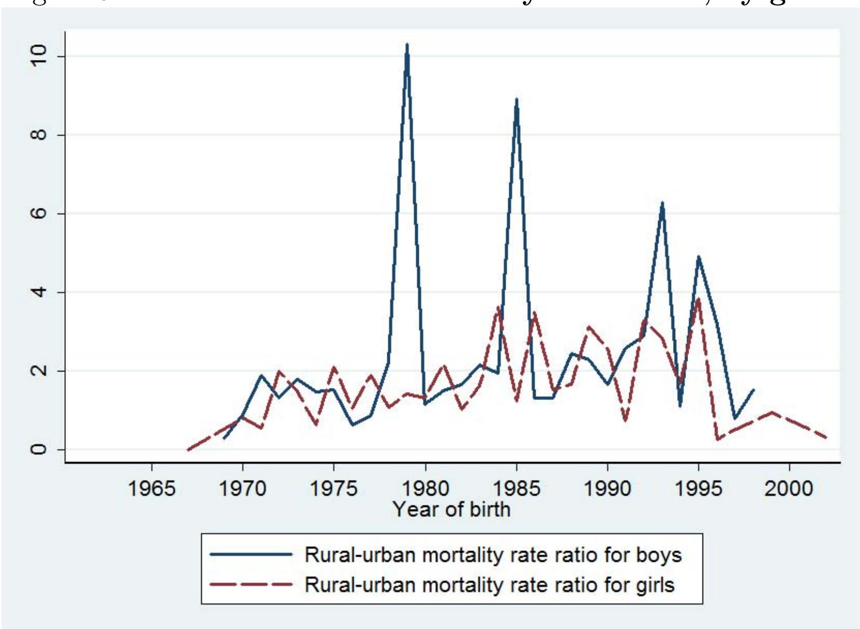


Figure 2: **The time trend of child mortality rate**



Source: VNDHS, 1997 and 2002

Figure 3: **Rural-urban child mortality rate ratios; by gender**



Source: VNDHS, 1997 and 2002

Figure 4: **Variation of hazard rate over survival time**

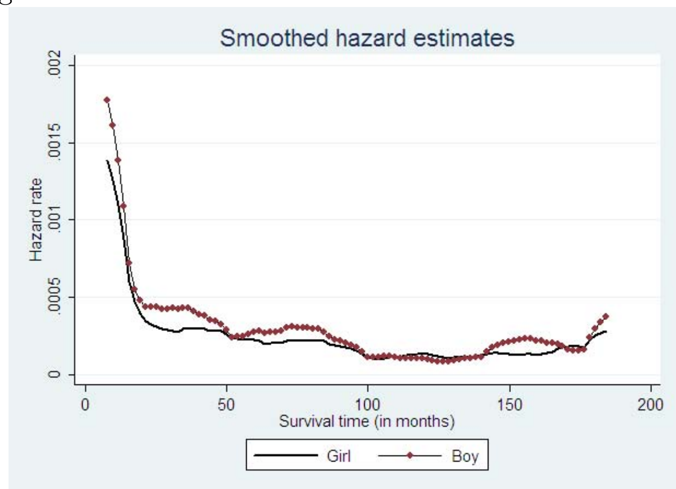


Table 1: **Vietnamese context in comparative heath indicators in 1995 and 2000**

Indicators	Vietnam		Indonesia		Bangladesh		India		China	
	1995	2000	1995	2000	1995	2000	1995	2000	1995	2000
Population (millions)	73	78	193	206	126	139	932	1,016	1,205	1,263
GNI per capita (current USD)	250	390	1,010	590	310	360	380	450	530	930
Adult literacy rate (% age 15 and above)	92.1	90.3	83.7	83.4	38	49	53	61	80.8	90.9
Immunization coverage (% of children aged 12-23 months)	95	97	63	72	79	76	72	52	80	85
Infant mortality rate (per 1,000 live births)	32	23	48	36	83	66	74	66	37	30
Under-5 mortality rate (per 1,000 live births)	44	30	66	48	120	90	102	89	44	37
Life expectancy at birth (years)	67	69	64	66	58	61	61	63	69	70
Total fertility rate (births per woman)	2.7	1.9	2.7	2.4	3.8	3.3	3.4	3.1	1.9	

Source: World Development Indicators database, World Bank, 2006 and CIA World Factbook, 2002.

Table 2: **Distribution of mothers by number of births and deaths**

	Number of deaths								
Number of births	0	1	2	3	4	5	6	7	Total
1	2,242	25	0	0	0	0	0	0	2,267
2	3,397	105	4	0	0	0	0	0	3,506
3	1,890	254	14	1	0	0	0	0	2,159
4	1,019	272	32	2	0	0	0	0	1,325
5	454	207	47	9	0	0	0	0	717
6	225	117	46	8	0	1	0	0	397
7	79	45	35	5	2	1	0	0	167
8	38	39	20	6	2	0	0	1	106
9	16	17	8	7	4	0	0	0	52
10	8	6	5	4	3	1	0	0	27
11	2	4	3	0	1	0	0	0	10
12	0	0	0	0	0	0	1	0	1
Total	9,370	1,091	214	42	12	3	1	1	10,734

Table 3: **Definition and descriptive statistics of variables**

Variable	Observation	Mean	Std.Dev	Min	Max
Age at death or censor (in months)	29,900	135.246	84.696	0.03	438
Child mortality	29,900	0.058	0.233	0	1
Birth characteristics					
First birth	29,900	0.361	0.48	0	1
High birth order (> 4)	29,900	0.098	0.297	0	1
Short preceding birth interval	29,900	0.166	0.372	0	1
Child is male	29,900	0.519	0.500	0	1
Fertility behaviour					
Age at first birth*	29,900	21.60	3.33	12	43
Age at birth is less than 18	29,900	0.021	0.144	0	1
Age at birth is greater than 35	29,900	0.042	0.200	0	1
Socio-economic characteristics					
Mother has no education	29,900	0.082	0.274	0	1
Mother has primary or secondary education*	29,900	0.899	0.301	0	1
Mother has high education	29,900	0.019	0.136	0	1
Rural areas	29,900	0.816	0.388	0	1
Vietnamese ethnicity	29,900	0.837	0.370	0	1
Regions					
Northern Uplands	29,900	0.197	0.398	0	1
Red River Delta	29,900	0.171	0.376	0	1
Northern Central	29,900	0.152	0.359	0	1
Central Coast	29,900	0.106	0.308	0	1
Central Highlands	29,900	0.042	0.201	0	1
Southeast	29,900	0.105	0.306	0	1
Mekong River Delta*	29,900	0.228	0.419	0	1
Calendar time trend					
Born before 1976*	29,900	0.048	0.214	0	1
Born between 1976 and 1986	29,900	0.382	0.486	0	1
Born between 1986 and 1989	29,900	0.146	0.353	0	1
Born between 1989 and 1993	29,900	0.200	0.400	0	1
Born between 1993 and 1996	29,900	0.122	0.328	0	1
Born after 1996	29,900	0.101	0.302	0	1

* Reference group

Table 4: **The no-frailty and frailty piecewise constant exponential model**

Variable	No-frailty model		Frailty model	
	Coefficient	s.e.	Coefficient	s.e.
Survival time intervals				
Month 0 - 1	5.653^{***}	0.073	5.618^{***}	0.073
Month 1 - 6	2.751^{***}	0.101	2.724^{***}	0.101
Month 6 - 12	2.205^{***}	0.106	2.182^{***}	0.106
Month 12 - 24	1.507^{***}	0.107	1.489^{***}	0.107
Month 24 - 60	1.232^{***}	0.086	1.219^{***}	0.086
Birth characteristics				
First birth	0.304^{***}	0.061	0.313^{***}	0.062
High birth order (> 4)	0.183^{**}	0.083	0.098	0.086
Short preceding birth interval	0.625^{***}	0.062	0.582^{***}	0.064
Child is male	0.262^{***}	0.049	0.268^{***}	0.050
Fertility behaviour				
Age at first birth	-0.031^{***}	0.009	-0.033^{***}	0.009
Age at birth is less than 18	0.330^{***}	0.128	0.355^{***}	0.133
Age at birth is greater than 35	0.238^{*}	0.144	0.277^{*}	0.147
Socio-economic characteristics				
Mother has no education	0.368^{***}	0.075	0.385^{***}	0.087
Mother has high education	-0.600^{**}	0.306	-0.601^{*}	0.312
Rural areas	0.318^{***}	0.077	0.342^{***}	0.081
Vietnamese ethnicity	-0.389^{***}	0.067	-0.385^{***}	0.074
Regions				
Northern Uplands	-0.170^{**}	0.076	-0.162^{**}	0.083
Red River Delta	-0.306^{***}	0.086	-0.327^{***}	0.091
Northern Central	-0.038	0.079	-0.050	0.086
Central Coast	0.001	0.084	-0.026	0.094
Central Highlands	-0.167	0.131	-0.168	0.145
Southeast	-0.293^{***}	0.100	-0.299^{***}	0.107
Calendar time trend				
Born between 1976 and 1986	-0.187^{**}	0.089	-0.186^{**}	0.094
Born between 1986 and 1989	-0.281^{***}	0.104	-0.297^{***}	0.110
Born between 1989 and 1993	-0.649^{***}	0.106	-0.656^{***}	0.112
Born between 1993 and 1996	-0.683^{***}	0.123	-0.700^{***}	0.129
Born after 1996	-1.039^{***}	0.153	-1.054^{***}	0.157
Constant	-8.596^{***}	0.226	-8.524^{***}	0.242
σ^2			0.584^{***}	0.095
$\Pr > \overline{\chi}^2_{01}(\sigma^2)$			0.000	
Number of observation	165,395		165,395	
Number of children	29,900		29,900	
Log-likelihood	$-10,545$		$-10,512$	
AIC	21,146		21,083	

Note: s.e.: standard error; and *, ** and *** indicate significant level at 10%, 5% and 1%, respectively.

Table 5: **Proportional effect of calendar time and survival time on hazard rate**

Calendar time		Survival time	
Child born before 1976	1	Child aged within 1 month	18.067
Born between 1976 and 1986	0.831	Aged between 1 and 6 months	1
Born between 1986 and 1989	0.743	Aged between 6 and 12 months	0.581
Born between 1989 and 1993	0.519	Aged between 12 and 24 months	0.291
Born between 1993 and 1996	0.496	Aged between 24 and 60 months	0.222
Born after 1996	0.349	Aged greater than 60 months	0.066

Table 6: **Estimation results by gender**

Variable	Total sample		Boys		Girls	
	Coefficient	s.e.	Coefficient	s.e.	Coefficient	s.e.
Survival time intervals						
Month 0 - 1	5.653^{***}	0.073	5.680^{***}	0.098	5.544^{***}	0.111
Month 1 - 6	2.751^{***}	0.101	2.744^{***}	0.135	2.710^{***}	0.154
Month 6 - 12	2.205^{***}	0.106	2.068^{***}	0.148	2.321^{***}	0.154
Month 12 - 24	1.507^{***}	0.107	1.490^{***}	0.144	1.496^{***}	0.162
Month 24 - 60	1.232^{***}	0.086	1.260^{***}	0.115	1.170^{***}	0.131
Birth characteristics						
First birth	0.304^{***}	0.061	0.297^{***}	0.082	0.332^{***}	0.095
High birth order (> 4)	0.183	0.083	0.218^{**}	0.110	0.018	0.137
Short preceding birth interval	0.625^{***}	0.062	0.545^{***}	0.084	0.689^{***}	0.096
Child is male	0.262^{***}	0.049	-	-	-	-
Fertility behaviour						
Age at first birth	-0.031^{***}	0.009	-0.037^{***}	0.012	-0.026^{*}	0.014
Age at birth is less than 18	0.330^{***}	0.128	0.326^{*}	0.178	0.349^{*}	0.199
Age at birth is greater than 35	0.238^{*}	0.144	0.004	0.208	0.578^{***}	0.208
Socio-economic characteristics						
Mother has no education	0.368^{***}	0.075	0.358^{***}	0.110	0.404^{***}	0.120
Mother has high education	-0.600^{**}	0.306	-0.310	0.368	-1.103^{*}	0.585
Rural areas	0.318^{***}	0.077	0.447^{***}	0.108	0.185	0.116
Vietnamese ethnicity	-0.389^{***}	0.067	-0.403^{***}	0.096	-0.386^{***}	0.106
Regions						
Northern Uplands	-0.170^{**}	0.076	-0.193^{*}	0.108	-0.128	0.118
Red River Delta	-0.306^{***}	0.086	-0.341^{***}	0.120	-0.296^{**}	0.132
Northern Central	-0.038	0.079	-0.030	0.111	-0.064	0.125
Central Coast	0.001	0.084	-0.026	0.121	0.012	0.134
Central Highlands	-0.167	0.131	0.018	0.178	-0.445^{*}	0.230
Southeast	-0.293^{***}	0.100	-0.250^{*}	0.141	-0.358^{**}	0.154
Calendar time trend						
Born between 1976 and 1986	-0.187^{**}	0.089	-0.269^{**}	0.122	-0.075	0.141
Born between 1986 and 1989	-0.281^{***}	0.104	-0.314^{**}	0.143	-0.267	0.167
Born between 1989 and 1993	-0.649^{***}	0.106	-0.552^{***}	0.143	-0.835^{***}	0.176
Born between 1993 and 1996	-0.683^{***}	0.123	-0.717^{***}	0.168	-0.673^{***}	0.196
Born after 1996	-1.039^{***}	0.153	-1.216^{***}	0.213	-0.848^{***}	0.230
Constant	-8.596^{***}	0.226	-8.222^{***}	0.316	-8.591^{***}	0.349
σ^2	0.584^{***}	0.095	0.689^{***}	0.156	0.195^{**}	0.159
Number of observation	165,395		85,466		79,929	
Number of children	29,900		8,751		8,135	
Log-likelihood	$-10,545$		$-6,002$		$-4,510$	
AIC	21,146		12,060		9,077	