# Child Mortality and Health Programs in Vietnam's Transitional Period

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**Abstract**

We analyze changes in child mortality in Vietnam during four decades of profound changes in the country's political, socio-economic, and health care conditions. Our data and econometric models allow for a difference-in-differences approach to estimate the impact of the Population and Family Health Project (PFHP), one of several health programs that have been implemented. We find strong improvements in overall mortality rates, although large regional differences persist. About half of the regional disparities can be attributed to differences in education and fertility-related behaviors. We find no evidence of a direct causal impact of the PFHP program on mortality rates.

**Keywords:** Child mortality; health policies; hazard rate; difference in differences.
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# 1 Introduction

Child mortality has long been a focus of academic researchers and policy makers in developing countries. Studies usually attribute early child mortality to maternal, socio-economic, biological and environmental factors (Cochrane et al., 1982; Caldwell and McDonald, 1982; Mosley and Chen, 1984; Hobcraft et al., 1985; Rutstein, 2000). 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 (Hobcraft et al., 1983; Palloni and Millman, 1986; Boerma and Bicego, 1992; Nath et al., 1994) while others explore the causal relationship between socio-economic factors and child mortality (Trussell and Hammerslough, 1983; Hobcraft et al., 1985; Guilkey and Riphahn, 1998).

Most of the studies for Vietnam have used the 1988 Vietnam Demographic and Health Survey (VNDHS) and the 1990 Vietnam Accessibility of Contraceptives Survey (with data on children born between 1983 and 1990). Swenson et al. (1993) and Nguyen-Dinh and Feeny (1999) found that children born to illiterate mothers were at higher risk of death, but no differences were found between intermediate and higher levels of education; no effects were found for the age at birth of the mother and the sex of the child; and shorter birth spacing was associated with increased mortality risk. Recently, Wagstaff and Nga (2002) confirmed the role of the mothers' education in reducing the risk in a hazard model estimated using the 1993 and 1998 Vietnam Living Standard surveys.

The present paper's contribution to the existing literature is twofold. First, our analysis is based on a sample of close to 30,000 children born between 1960 and 2002. During this relatively long time span Vietnam has witnessed profound changes in terms of its political system, socio-economic conditions, and health services. This allows for a broader, longer-term perspective than in earlier studies. Secondly, we use econometric models that control for unobserved heterogeneity at the community or family level, allow for parametric flexibility, and permit the estimation of program effects using a difference-in-differences approach. We do so by using a Piecewise Constant hazard rate model with unobserved heterogeneity, as well as Linear Probability Models with random or fixed effects. While these econometric tools are now well-established, this paper is the first to apply them to the Vietnamese case.

We find strong improvements in the country-wide overall child mortality rate, from more than 1 percent in 1970 to less than 0.4 percent three decades later. However, large regional differences persist, with mortality rates in rural areas being almost twice as high as in urban areas. About half of the regional disparity can be attributed to differences in education and fertility related behaviors. We find no evidence of a direct causal impact of the PFHP program on the health outcomes.

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

## 2 The Vietnamese context

Health has long received high priority from the Vietnamese government since the independence from the French in the North in 1954 and the reunification of the whole country in 1975. This translated into 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 the 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. Figure 1 graphs the developments of the overall mortality rate in our data since 1960, and shows a strong and persistent decline. It also shows a large urban-rural disparity, with the rural mortality rates being about 75 percent higher. The magnitude of this difference is large. For example, Poel et al. (forthcoming), who analyze a cross-section of Western Sub-Saharan countries, find an average difference of 47 percent.

Vietnam's transition to a more market-oriented economy entailed a decline in health service delivery, equipment maintenance, and health workers' salaries. The widespread coverage of public health services was insufficient to ensure the quality, sufficiency and effectiveness of the network (Chen and Hiebert, 1994). Although the indicators on hospital beds and doctors of Vietnam were among the top in Asia, the public expenditure on health care was among the lowest, less than USD 1 per capita (World Bank, 1992). At the same time Vietnam has experienced a steady increase in maternal cognitive social capital and an advance in access to and quality of health care services. Although it is tempting to interpret this relationship as causal, they might also be explained by an autonomous time trend. These issues will be addressed in our econometric analysis.

Under the central planning economy before 1986, 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 the gap in living standards across social strata.[1] 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. Due to budgetary constraints and high inflation the availability of low cost services as well as their quality rapidly slumped. In 1989, Vietnam introduced user fee charges for health care as an attempt to reduce the budget burden, which severely restricted access to health services for low-income households. Wagstaff and Nga (2002) assert that *Doi Moi* brought much benefit for the better-off but little for the poor, in terms of child mortality and other health indicators.

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 and enhance 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 "Population and Family Health" project (PFHP) was another main part program. Its objectives were broad, and focused on the restoration

[1]The GDP in real terms was doubled during the period 1991-2000, with an average growth rate of approximate 7.5%, compared to 4.4% between 1986 and 1990, and to 3.7% between 1976 and 1985. Poverty incidence was 58.1% in 1993, 37.4% in 1998, and 28.9% in 2002 (GSO, 1998; GSO, 2002).

of the primary health care network, poverty alleviation, fertility reduction, and improving women's and childrens' health. The PFHP started in late 1996 and ended in 2002. Provinces involved in the project were those with relatively poor indicators of fertility, family planning and health indicators; inadequate support on family planning from the central government and international donors; high proportion of rural areas and disadvantaged ethnic minorities, regional spread and adequate management capacity. These criteria generally resulted in the selection of 15 relatively poor provinces out of 53 provinces. The actual costs were USD 108 million which was jointly financed by the Vietnamese government, World Bank, Asian Development Bank, and Kreditanstalt fur Wiederaufbau (KfW). About two thirds of this amount went directly to the project provinces, for the building/renewing of health stations, the supply of essential drugs and medicines and training of health workers at the commune and district levels. As a consequence, yearly, the project added about USD 0.5 to the health expenditure per capita per year in project provinces during 6 years. By comparison to the total health expenditures per capita in Vietnam of around USD 12 (World Bank, 1995), the project accounted for around 4% of the total health expenditures. A program evaluation was done soon afterwards. Most of the evaluation indicators were commended as highly satisfactory, but it was recognized that the true impact of the project was difficult to assess (World Bank, 2004).

The time line in figure 2 summarizes the main events and programs.

# 3 Data

The empirical analysis is based on the Vietnam Demographic and Health Surveys (VNDHS) from 1997 and 2002.[2] Both surveys are nationally-

[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.

representative household surveys. Data were collected at the level of individual ever-married women at 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. While information on birth history, child survival status and so forth is retrospective, that on socio-economic factors refers to conditions at the survey time.

In VNDHSs, survey sites covered 41 provinces throughout the country, 18 of which (formerly 15) were targeted by the PFHP (NCFPFC and PFHP, 2003). 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. Note that all individual observations can be classified as either "PFHP project province" or "PFHP non-project province".

From the complete birth history of sampled women, we have information on 29,900 live births to 10,734 mothers.[3] This child data set contains information on child survival status, time of death (if occurred before the time of the survey) as well as a number of socio-economic characteristics.

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

[3] There are 595 women who do not report any live births.

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').

Table 4 reports descriptive statistics for a number of health outcomes, health inputs, and socio-economic characteristics, by survey year and project and non-project provinces.

The year of birth of the children spans from 1961 to 2002. Nearly 90% of the children was born before 1997 when the project started. It should 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 1. 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.

Explanatory variables in the present study explore the effects of fertility behaviour and socio-economic characteristics on duration of child survival. Most of them follow the Mosley-Chen analytical framework, as far as data availability permits. As shown in table 4 most variables are not statistically different between project and non-project provinces, with a few exceptions. Mother's education level is substantially lower in project provinces, but - perhaps surprisingly - child mortality is lower.[4]

[4]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 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.

Figure 3 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.[5]

# 4 Empirical Methods

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)). 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. So, 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 $\sigma^2$. 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 $\sigma^2$. 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,

[5]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.

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 $\delta_j$ equals 1 if child $j$ dies at $t_j$ and 0, otherwise. Hence, the hazard rate of dying increases by $\sigma^2$ multiplied by the number of deaths. This expression provides a measure of the size of the mother-specific unobserved effects.

To allow for the project evaluation, we include a dummy variable for a project community ($D_P$), a dummy for the year 2002 ($D_Y$), and an interaction terms between both dummies. In the Linear Probability Model (LPM) the coefficient of the interaction term can be directly interpreted as a double difference estimator – conditional on other explanatory variables – of the effect of the project on the average probability (or hazard rate) of child mortality. In the other models it depends on both covariates and the underlying probability function:

$$\begin{aligned}\mu = &\left[\overline{\Pr(D_P=1,D_Y=1|X)}-\overline{\Pr(D_P=1,D_Y=0|X)}\right]\\ &-\left[\overline{\Pr(D_P=0,D_Y=1|X)}-\overline{\Pr(D_P=0,D_T=0|X)}\right]\end{aligned} \tag{3}$$

In those cases the standard error is calculated using the Delta method; see Ai and Norton (2003). The parameter ($\mu$) estimates the net policy impact on the treatment group by controlling both group-specific and time-specific effects on the outcomes. A causal interpretation requires additional assumptions. One is that the assignment of communities as a project community is uncorrelated with *time-varying* unobserved variables represented by $v$, conditional on observables. Note, though, that the being labeled as a project province took place before 1997, and that a possible correlation with dependent variables is controlled for in the fixed effect models.

# 5 Estimation results and discussion

Table 5 presents the estimation results of the hazard rate models with and without unoberved heterogeneity at the family level. As mentioned, we specify a piece-wise linear hazard, with intervals 0-1, 1-6, 6-12, 12-24, 24-60, with $> 60$ months as the reference category. Although the estimate of $\sigma^2$ in model with unobserved heterogeneity is large (0.56) and highly significant, the differences between the two sets of estimates are small. Our discussion below is based on the model with heterogeneity.

*Sibling correlations*

The presence of a mother specific random effect induces a correlation between the risk of death among siblings and family death clustering. Given the estimate of $\sigma^2$, 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. These factors operate through proximate determinants such as household income, amenities, parental education, and genetic factors. The estimate of $\sigma^2$ 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, although a set of covariates on household wealth and amenities is included in their model.[6]

*Time effects*

Table 6 presents the proportional effect of the calendar and survival time

[6]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 $\sigma^2$ drops to 0.465, but all other estimates remain virtually unchanged.

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 already showed a strong declining time trend of mortality as a function of calendar time. The estimated coefficients on calendar time in table 5 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 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% ($= e^{0.31} - 1$) and 11% ($= e^{0.10} - 1$) higher than that for the other order births, respectively. These findings are similar to those found in previous studies in Vietnam (Swenson et al., 1993; Swenson et al., 1995; Nguyen-Dinh and Feeny, 1999). The high risk for the first births is caused by more difficult delivery at first births (*dystocia*) and the lack of experiences of child-bearing and rearing. The lower risk for the later births may result from the learning effect 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 Lanca (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 among the most important determinants 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% ($= e^{0.58} - 1$) higher than that for others. The effects of short birth interval are also likely to be related to maternal depletion and sibling competition. Given low nutritional status and wealth, giving births soon after the previous birth may significantly increase the physical and mental burden with child-care and resource allocation, resulting in higher risk of child mortality. 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 is likely to be attributable to profound changes in socio-economic context during the transition in Vietnam initiated from 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. This changes might have induced maternal depletion and sibling competition.

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. However, in a model that does allow for a two-directional causal relationship (child mortality risk is influenced by the timing and spacing of births and birth-spacing and fertility are, in turn, a function of realized mortality) Bhalotra and van Soest (2008) still find strong effects, using data for India.

*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 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.

*Project versus non-project provinces*

Perhaps surprisingly, the mortality risk for a child in a project province is smaller than for a non-project counterpart, by about 19%. This is related to a non-random selection of project provinces: Provinces in which child mortality is already relatively low were more likely to be selected for the program. The estimated effect of the PFHP program on child mortality is negative (as expected), but is not statistically significant. Thus we find that the difference between the decline in the mortality risk in a project province and a non-project province is negligible, both before and after controlling for observed and unobserved heterogeneity. [7]

[7] Using modal values of the covariates in the model and following equation 3, the point estimate of the treatment effect for an under-one-month child in the Mekong Delta is $-0.001$ with the standard error of 0.002. Modal values are: First birth = 0; High birth order = 0; Short preceding birth interval = 0; Age at first birth = 22; Age at birth is less

One can think of several explanations for the lack of a significant program effect. First, project provinces might be on a flatter (and on a higher level) part of the health production function, so that the marginal benefits of the health program are smaller and less easily detected. Secondly, as shown by the timeline in figure 2, numerous health-related programs have been implemented simultaneously, and some of them have been implemented in both PFHP project and PFHP non-project provinces. This will obviously bias the program effect coefficient toward zero. Thirdly, one third of the PFHP funds were spent on non-project provinces. Fourthly, the PFPH was relatively small, with per capita investments of less than a dollar. Hence, the program might simply have been too small to lead to detectable improvements in child survival and health in general. In addition, it is conceivable that the government, or other agencies, would invest more in non-project provinces from other sources, as a way of compensation.

While we do not find a separate effect of this specific program, the trend variable unambiguously points at substantial improvement in Vietnam's child survival rates in recent decades, which is consistent with a substantial combined effect of the various programs and policies.

As a specification and robustness check, we have also estimated linear probability models (LPM) with community specific fixed effects or mother specific fixed effects. Separate models were estimated for explaining neonatal, infant, and child mortality. The LPM models allow for a flexible specification of unobserved heterogeneity, as the fixed effects model does not require the assumption of independence between errors terms and explanatory variables (as does the hazard rate model with mother specific effects). Thus, comparing the LPM fixed and random effects models provides an indication of the impact of this assumption on the hazard rate model's estimation re-

than 18 = 0; Age at birth is greater than 35 = 0; Mother has no education = 0; Mother has high education = 0; Rural areas = 1; Vietnamese ethnicity = 1; Year of birth of the child = 1992 if born before 1997 and 2002 if born after 1997 sults. We find, however, that the difference between the random effects and fixed effects LPM models are small. This suggests that any bias that could result from the orthogonality assumption in the hazard rate model is small.

## 6 Conclusion

We have analyzed child mortality in Vietnam during four decades characterized by profound changes in the country's political system, socio-economic characteristics, and the implementation of numerous health care programs. Our econometric models mimic the strong clustering of deaths within families, and show significant effects of education and fertility behaviors. Applying a difference-in-differences approach, we do not find a separate contribution to this decline by the Population and Family Health Project, a health care program implemented in a selection of provinces between 1997 and 2002. Most of the decline in mortality rates remains unexplained, as witnessed by the large estimated calendar time effects.

Along with the strong overall decline in mortality rates we find large and persistent regional differences, with children in the Mekong River Delta, Northern Central, Central Coast, and Central Highlands facing an about 40 percent higher mortality risk than those living elsewhere. The combination of overall improvements and persistent inequalities was also found by O'Donnell et al. (2009) in a recent study on the nutritional status of children in Vietnam.

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.

## References

Ai, C. and Norton, E. C. (2003). Interaction terms in logit and probit models, *Economics Letters* **80**: 123–129.

Anh, D. (1995). Child survival in vietnam: An investigation of quality of life. paper presented at the 1995 American Population Association Annual Meeting, San Francisco.

Bhalotra, S. and van Soest, A. (2008). Birth-spacing, fertility and neonatal mortality in India: Dynamics, frailty, and fecundity, *Journal of Econometrics* **143**: 274–290.

Boerma, J. T. and Bicego, G. T. (1992). Preceding birth intervals and child survival: Searching pathways of influence, *Studies in Family Planning* **23**(4).

Caldwell, J. and McDonald, P. (1982). Influence of maternal education on infant and child mortality: Levels and Causes, *Health Policy and Education* **2**: 251–267.

Chen, L. and Hiebert, L. (1994). From socialism to private markets: Vietnam's health in rapid transition, *Technical Report 94.11*, Working paper series, Harvard Center for Population and Development Studies, Cambridge, MA.

Clayton, D. (1978). A model for association in bivariate life tables and its application in epidemiological studies of familial tendency in chronic disease incidence, *Biometrika* **65**(Supplement: Child survival: Strategies for research): 141–151.

Cochrane, S., Leslie, J. and O'Hara, D. (1982). Parental education and child health: Intracountry evidence, *Health policy and Education* **2**: 213–250.

Greene, W. (2003). *Econometrics analysis*, fifth edn, Prentice Hall, Pearson Education, Inc., Upper Saddle River, New Jersey, 07458.

GSO (1998). *Vietnam Living Standard Survey 1997-1998*, Hanoi, Vietnam.

GSO (2002). *Vietnam Household Living Standard Survey 2002*, Hanoi, Vietnam.

Guilkey, D. K. and Riphahn, R. T. (1998). The determinants of child mortality in the philippines: estimation of a structural model, *Journal of Development Economics* **56**: 281–305.

Guo, G. (1993). Use of sibling data to estimate family mortality effects in guatemala, *Demography* **30**(1): 15–32.

Guo, G. and Rodriguez, G. (1992). Estimating a multivariate proportional hazards model for clustered data using the em algorithm, with an application to child survival in guatemala, *Journal of the American Statistical Association* **87**(420): 969–976.

Gutierrez, R. G. (2002). Parametric frailty and shared frailty survival models, *Stata Journal* **2**(1): 22–44(23).

Gutierrez, R. G., Carter, S. and Drukker, D. M. (2001). On boundary-value likelihood-ratio tests, *Stata Statical Bulletin* **10**(60): 15–18.

Hobcraft, J., McDonald, J. W. and Rutstein, S. (1983). Child-spacing effects on infant and early child mortality, *Population Index* **49**(4): 585–618.

Hobcraft, J., McDonald, J. W. and Rutstein, S. (1984). Socio-economic factors in infant and child mortality: A cross-national comparison, *Population Studies* **38**(2): 193–223.

Hobcraft, J., McDonald, J. W. and Rutstein, S. (1985). Demographic determinants of infant and early child mortality: A comparative analysis, *Population Studies* **39**(3): 363–385.

Horton, S. (1988). Birth order and child nutritional status: Evidence from the philippines, *Economic Development and Cultural Change* **36**(2): 341–354.

Lawn, J. E., Cousens, S. and Zupan, J. (2005). 4 million neonatal deaths: When? Where? Why?, *Lancet* **365**: 891–900.

Majumder, A. K. and Islam, S. M. S. (1993). Socioeconomic and environmental determinants of child survival in Bangladesh, *Journal of Biosocial Science* **25**: 311–318.

Makepeace, G. and Pal, S. (2006). Effects of birth interval on child mortality: Evidence from a sequential analysis, *World Health & Population* pp. 1–14.

Mosley, W. and Chen, L. (1984). An analytical framework for the study of child survival in developing countries, *Population and Development Review* **10**(Supplement: Child survival: Strategies for research): 24–45.

Nath, D. C., Land, K. C. and Singh, K. K. (1994). Birth spacing, breastfeeding, and early child mortality in a traditional Indian society: A hazards model analysis, *Social Biology* **41**(3-4): 168–180.

NCFPFC and PFHP (2003). *Vietnam Demographic and Health Survey 2002*, Hanoi, Vietnam.

Nguyen-Dinh, H. and Feeny, D. H. (1999). Are parental characteristics important for child survival? The case of Vietnam, *Pacific Economic Review* **4**(1): 1–29.

O'Donnell, O., Nicolas, A. L. and van Doorslaer, E. (2009). Growing richer and taller: Explaining change in the distribution of child nutritional status during vietnam's economic boom, *Journal of Development Economics* **88**: 45–58.

Omariba, D. W. R., Beaujot, R. and Rajulton, F. (2007). Determinants of infant and child mortality in kenya: An analysis controlling for frailty effects, *Population Research and Policy Review* **26**(3): 299–321.

Palloni, A. and Millman, S. (1986). Effects of inter-birth intervals and breast-feeding on infant and early childhood mortality, *Population Studies* **40**(2): 215–236.

Poel, E., O'Donnell, O. and van Doorslaer, E. (forthcoming). What explains the rural-urban gap in infant mortality - household or community characteristics?, *Demography* .

Rutstein, S. (2000). Factors associated with trends in infant and child mortality in developing countries during 1990s, *Bulletin of the World Health Organization* **78**: 1256–1270.

Swenson, I. E., Thang, N. M., San, P. B., Nhan, V. N. and Man, V. D. (1993). Factors influencing infant mortality in Vietnam, *Journal of Biosocial Science* **25**: 285–302.

Swenson, I. E., Thang, N. M., San, P. B., Nhan, V. N. and Man, V. D. (1995). Early childhood survivorship in Vietnam, *Journal of Tropical Medicine and Hygiene* **98**: 204–208.

Trussell, J. and Hammerslough, C. (1983). A hazards-model analysis of the covariates of infant and child mortality in Sri Lanka, *Demography* **20**: 1–26.

Vaupel, J. W., Manton, K. G. and Stallard, E. (1979). The impact of heterogeneity in individual frailty on the dynamics of mortality, *Demography* **16**(3): 439–454.

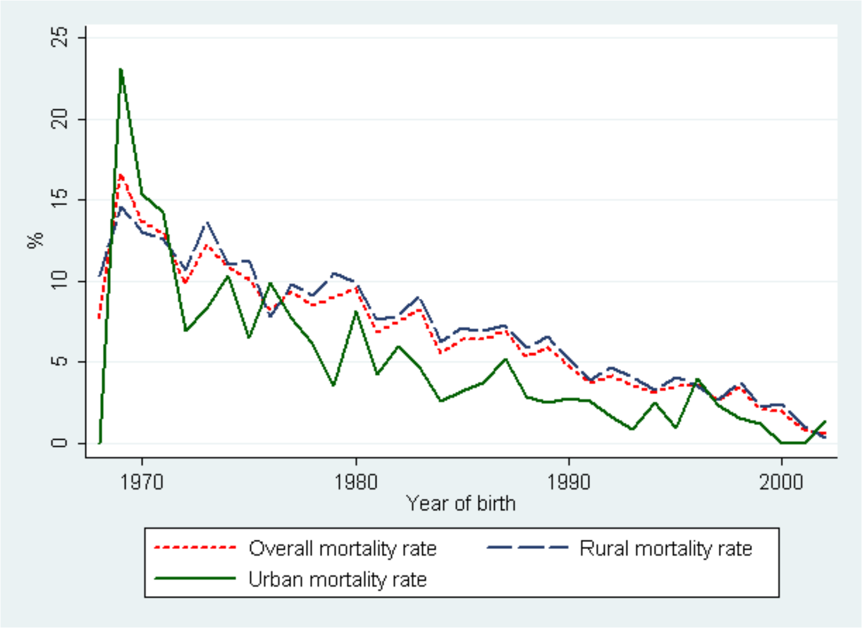
Wagstaff, A. and Nga, N. N. (2002). Poverty and child survival prospects of Vietnamese children under *doi moi*, *Policy research working paper*, World Bank.

World Bank (1992). Awakening the market, Viet Nam's economic transition, *World Bank Discussion Paper 157*.

World Bank (1995). Socialist Republic of Vietnam: Population and Family Health Project, *Staff appraisal report 14966-VN*.

World Bank (2004). Implementation completion report (IDA-28070) on a credit in the amount of SDR 33.6 million (US$ 44.9 million equivalent) to the Socialist Republic of Vietnam for Population & Family Health Project, *Report 28400*, Human Development Sector Unit, East Asia and Pacific Region.

Figure 1: **The time trend of mortality rate**




Source: VNDHS, 1997 and 2002

Figure 2: **Time line of relevant events during the transitional period**

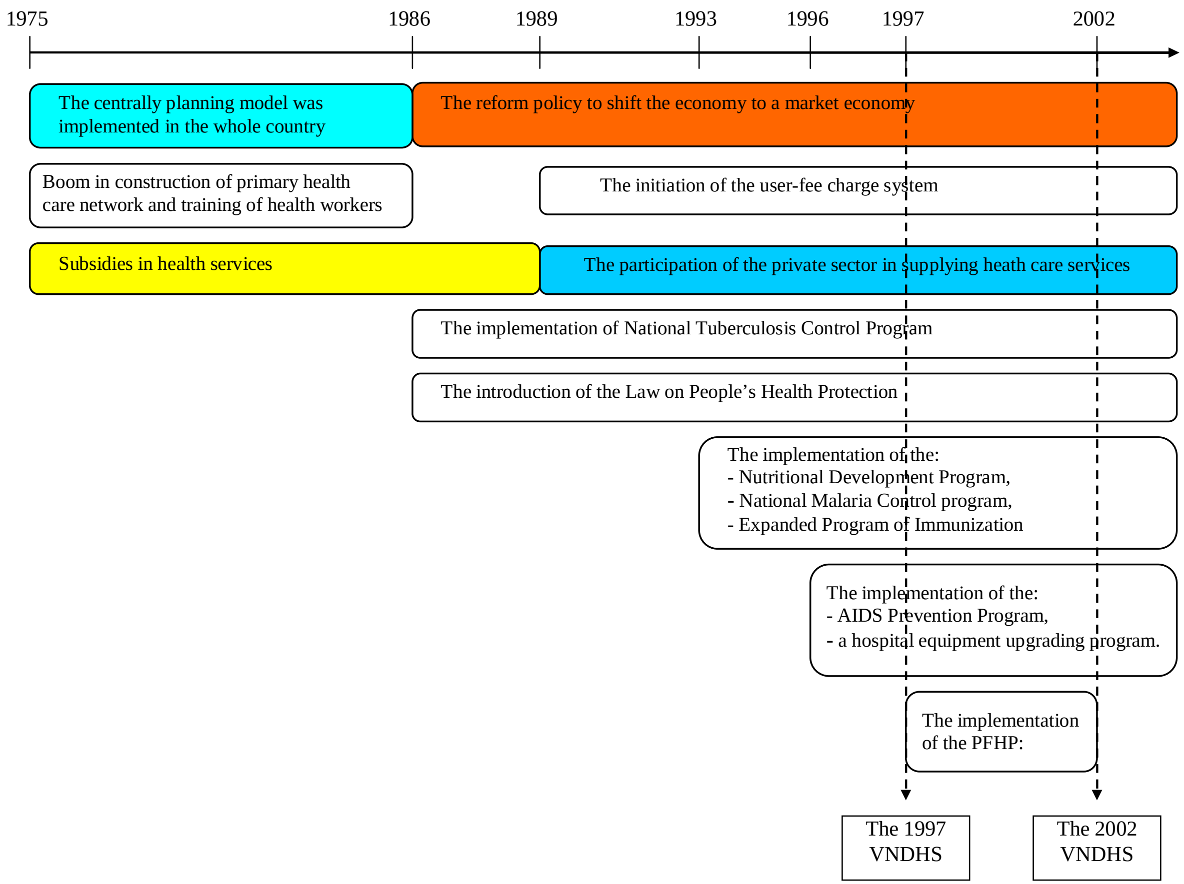


Figure 3: **Variation of hazard rate over survival time**

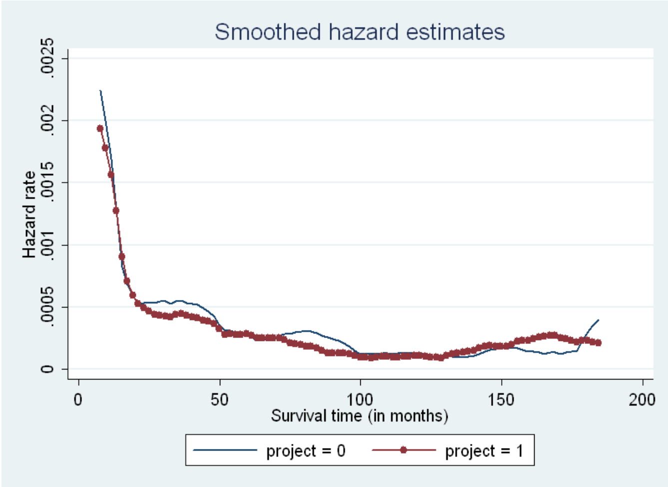


Table 1: **Vietnamese context in comparative heath indicators in 1995 and 2000**

| 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* | Mean | Std.Dev | Min | Max |
|---|---|---|---|---|
| Age at death or at time of survey (months)* | 135.246 | 84.696 | 0.03 | 438 |
| Child mortality | 0.058 | 0.233 | 0 | 1 |
| Birth characteristics | | | | |
| First birth | 0.361 | 0.48 | 0 | 1 |
| High birth order (> 4) | 0.098 | 0.297 | 0 | 1 |
| Short preceding birth interval | 0.166 | 0.372 | 0 | 1 |
| Fertility behaviour | | | | |
| Age at first birth* | 21.60 | 3.33 | 12 | 43 |
| Age at birth is less than 18 | 0.021 | 0.144 | 0 | 1 |
| Age at birth is greater than 35 | 0.042 | 0.200 | 0 | 1 |
| Socio-economic characteristics | | | | |
| Mother has no education | 0.082 | 0.274 | 0 | 1 |
| Mother has primary or secondary education** | 0.899 | 0.301 | 0 | 1 |
| Mother has high education | 0.019 | 0.136 | 0 | 1 |
| Rural areas | 0.816 | 0.388 | 0 | 1 |
| Vietnamese ethnicity | 0.837 | 0.370 | 0 | 1 |
| Regions | | | | |
| Northern Uplands | 0.197 | 0.398 | 0 | 1 |
| Red River Delta | 0.171 | 0.376 | 0 | 1 |
| Northern Central | 0.152 | 0.359 | 0 | 1 |
| Central Coast | 0.106 | 0.308 | 0 | 1 |
| Central Highlands | 0.042 | 0.201 | 0 | 1 |
| Southeast | 0.105 | 0.306 | 0 | 1 |
| Mekong River Delta** | 0.228 | 0.419 | 0 | 1 |
| Calendar time trend | | | | |
| Born before 1976** | 0.048 | 0.214 | 0 | 1 |
| Born between 1976 and 1986 | 0.382 | 0.486 | 0 | 1 |
| Born between 1986 and 1989 | 0.146 | 0.353 | 0 | 1 |
| Born between 1989 and 1993 | 0.200 | 0.400 | 0 | 1 |
| Born between 1993 and 1996 | 0.122 | 0.328 | 0 | 1 |
| Born after 1996 | 0.101 | 0.302 | 0 | 1 |
| Treatment variables | | | | |
| Project province | 0.352 | 0.478 | 0 | 1 |
| Born in project provinces after 1996 (DinD) | 0.037 | 0.189 | 0 | 1 |

Number of observations: 29,000.

Table 4: **Descriptive statistics of variables by projects and by years**

| Variable* | Children born before 1997 | | Children born from 1997 | | Differences between | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | NP | P | NP | P / | a | b | c | d | e |
| Age at death or time of survey (months) | 146.161 | 148.288 | 31.725 | 31.631 | −116.657*** | −114.436*** | −0.094 | 2.127** | −2.221 |
| Child mortality | 0.065 | 0.056 | 0.022 | 0.018 | −0.038*** | −0.042*** | −0.004 | −0.009*** | 0.005 |
| Post-neonatal death | 0.037 | 0.035 | 0.007 | 0.009 | −0.026*** | −0.031*** | 0.003 | −0.002 | 0.005 |
| Neonatal death | 0.029 | 0.021 | 0.016 | 0.009 | −0.012*** | −0.042*** | −0.004 | −0.009*** | 0.005 |
| Infant death | 0.037 | 0.035 | 0.007 | 0.009 | −0.026*** | −0.031*** | 0.003 | −0.002 | 0.005 |
| Post-neonatal and first birth death | 0.033 | 0.033 | 0.007 | 0.002 | −0.030*** | −0.026*** | −0.005 | 0.000 | −0.004 |
| Mortality aged > 12 months | 0.027 | 0.024 | 0.003 | 0.006 | −0.018*** | −0.024*** | 0.003 | −0.003 | 0.006 |
| Mortality aged > 12 months and first birth | 0.023 | 0.023 | 0.005 | 0.003 | −0.020** | −0.018*** | −0.002 | 0.000 | -0.003 |
| Mortality aged > 60 months | 0.012 | 0.011 | 0.000 | 0.000 | −0.011 | −0.012* | 0.000 | −0.001 | 0.001 |
| Mortality aged > 60 months and first birth | 0.011 | 0.012 | 0.000 | 0.000 | −0.012 | −0.011* | 0.000 | 0.001 | −0.001 |
| Child is the first birth | 0.357 | 0.358 | 0.397 | 0.393 | 0.035** | 0.040*** | −0.004 | 0.002 | −0.005 |
| Birth order is greater than 4 | 0.102 | 0.101 | 0.065 | 0.073 | −0.028*** | −0.037*** | 0.008 | −0.001 | 0.009 |
| Age of the mother at first birth | 21.494 | 21.615 | 22.215 | 22.004 | 0.389*** | 0.721*** | −0.211 | 0.121*** | −0.332** |
| Age at birth is less than 18 | 0.023 | 0.021 | 0.011 | 0.019 | −0.002 | −0.012*** | 0.008* | −0.002 | 0.010* |
| Age at birth is greater than 35 | 0.036 | 0.039 | 0.082 | 0.084 | 0.044*** | 0.046*** | 0.002 | 0.004 | −0.002 |
| Short preceding birth interval | 0.175 | 0.171 | 0.090 | 0.114 | −0.057*** | −0.084*** | 0.024** | −0.003 | 0.027* |
| Mother has no education | 0.065 | 0.110 | 0.056 | 0.152 | 0.042*** | −0.009 | 0.096*** | 0.045*** | 0.052*** |
| Mother has primary or secondary education | 0.917 | 0.874 | 0.901 | 0.828 | −0.046*** | −0.015** | −0.073*** | −0.042*** | −0.031** |
| Mother has high education | 0.018 | 0.016 | 0.043 | 0.020 | 0.004 | 0.025*** | −0.023*** | −0.002 | −0.021*** |
| Mother is living in rural areas | 0.825 | 0.795 | 0.823 | 0.825 | 0.030** | −0.002 | 0.001 | −0.030*** | 0.031** |
| Ethnicity is Vietnamese | 0.850 | 0.817 | 0.856 | 0.754 | −0.064*** | 0.006 | −0.103*** | −0.033*** | −0.070*** |
| Mother is living in Northern Uplands | 0.218 | 0.161 | 0.195 | 0.172 | 0.010 | −0.023** | −0.023 | −0.056*** | 0.033** |
| Mother is living in Red River Delta | 0.121 | 0.266 | 0.123 | 0.220 | −0.046*** | 0.002 | 0.097*** | 0.145*** | −0.048*** |
| Mother is living in Northern Central | 0.125 | 0.205 | 0.119 | 0.188 | −0.017 | −0.006 | 0.069*** | 0.080*** | −0.011 |
| Mother is living in Central Coast | 0.159 | 0.000 | 0.208 | 0.000 | 0.000 | 0.049*** | −0.208*** | −0.159*** | −0.049*** |
| Mother is living in Central Highlands | 0.000 | 0.115 | 0.000 | 0.156 | 0.041*** | 0.000 | 0.156*** | 0.115*** | 0.041*** |
| Mother is living in Southeast | 0.161 | 0.000 | 0.169 | 0.000 | 0.000 | 0.008 | −0.169*** | −0.161*** | −0.008 |
| Mother is living in Mekong River Delta | 0.216 | 0.253 | 0.186 | 0.264 | 0.011 | −0.030*** | 0.078*** | 0.037*** | 0.041** |
| Year of birth of the child | 1986.28 | 1986.41 | 1999.04 | 1999.11 | 12.700*** | 12.750*** | 0.073 | 0.130 | −0.060 |

*, ** and *** represent significance level at 1%, 5% and 10%, respectively.

P: project provinces and NP: non-project provinces. The numbers of observations in the four groups of children from the left to the right are 17,456, 9,418, 1,914 and 1,112, respectively.

a: after and before 1997 in project provinces ($P_{\text{after}} - P_{\text{before}}$);

b: after and before 1997 in non-project provinces ($NP_{\text{after}} - NP_{\text{before}}$);

c: project and non-project province after 1997 ($P_{\text{after}} - NP_{\text{after}}$);

d: project and non-project province before 1997 ($P_{\text{before}} - NP_{\text{before}}$);

and e: double difference $[(P_{\text{after}} - P_{\text{before}}) - (NP_{\text{after}} - NP_{\text{before}})]$.

Table 5: **Piecewise constant hazard model, without and with mother specific effects**

| Variable | No-frailty model Coefficient | s.e. | Frailty model Coefficient | s.e. |
|---|---|---|---|---|
| Survival time intervals | | | | |
| Month 0 - 1 | 5.653*** | 0.073 | 5.620*** | 0.073 |
| Month 1 - 6 | 2.751*** | 0.101 | 2.726*** | 0.101 |
| Month 6 - 12 | 2.205*** | 0.106 | 2.183*** | 0.106 |
| Month 12 - 24 | 1.507*** | 0.107 | 1.489*** | 0.107 |
| Month 24 - 60 | 1.232*** | 0.086 | 1.220*** | 0.086 |
| Birth characteristics | | | | |
| First birth | 0.301*** | 0.061 | 0.310*** | 0.062 |
| High birth order ($> 4$) | 0.186** | 0.083 | 0.100 | 0.086 |
| Short preceding birth interval | 0.622*** | 0.062 | 0.580*** | 0.063 |
| Fertility behaviour | | | | |
| Age at first birth | −0.030*** | 0.009 | −0.032*** | 0.009 |
| Age at birth is less than 18 | 0.331** | 0.128 | 0.358*** | 0.132 |
| Age at birth is greater than 35 | 0.231 | 0.144 | 0.264 | 0.147 |
| Socio-economic characteristics | | | | |
| Mother has no education | 0.404*** | 0.076 | 0.419*** | 0.087 |
| Mother has high education | −0.606** | 0.306 | −0.607* | 0.312 |
| Rural areas | 0.293*** | 0.077 | 0.316*** | 0.081 |
| Vietnamese ethnicity | −0.394*** | 0.067 | −0.390*** | 0.074 |
| Regions | | | | |
| Northern Uplands | −0.184** | 0.076 | −0.181** | 0.083 |
| Red River Delta | −0.273*** | 0.086 | −0.293*** | 0.092 |
| Northern Central | −0.026 | 0.079 | −0.039 | 0.086 |
| Central Coast | −0.089 | 0.087 | −0.112 | 0.097 |
| Central Highlands | −0.032 | 0.136 | −0.038 | 0.150 |
| Southeast | −0.388*** | 0.102 | −0.392*** | 0.109 |
| Calendar time trend | | | | |
| Born between 1976 and 1986 | −0.193*** | 0.089 | −0.195*** | 0.093 |
| Born between 1986 and 1989 | −0.290*** | 0.104 | −0.306*** | 0.110 |
| Born between 1989 and 1993 | −0.653*** | 0.106 | −0.663*** | 0.112 |
| Born between 1993 and 1996 | −0.688*** | 0.123 | −0.707*** | 0.129 |
| Born after 1996 | −0.957*** | 0.176 | −0.981*** | 0.180 |
| Treatment variables | | | | |
| Project province | −0.218*** | 0.059 | −0.209*** | 0.065 |
| Born in project provinces after 1996 (DinD) | −0.230 | 0.276 | −0.211 | 0.279 |
| Constant | −8.350*** | 0.225 | −8.281*** | 0.242 |
| $\sigma^2$ | | | 0.565*** | 0.094 |
| $\Pr > \overline{\chi}^2_{01}$ | | | 0.000 | |
| Number of observation | 165,395 | | 165,395 | |
| Number of children | 29,900 | | 29,900 | |
| Log-likelihood | $-10,552$ | | $-10,521$ | |
| AIC | 21,161 | | 21,102 | |
| $\Pr > \chi^2$ | 0.0000 | | 0.0000 | |

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

Table 6:
**Proportional effect of calendar time and survival time on hazard rate**

| Survival time | | Calendar time | |
|---|---|---|---|
| Child aged less than 1 month | 18.08 | Child born before 1976 | 1 |
| Aged between 1 and 6 months | 1 | Born between 1976 and 1986 | 0.823 |
| Aged between 6 and 12 months | 0.581 | Born between 1986 and 1989 | 0.736 |
| Aged between 12 and 24 months | 0.291 | Born between 1989 and 1993 | 0.515 |
| Aged between 24 and 60 months | 0.222 | Born between 1993 and 1996 | 0.493 |
| Aged greater than 60 months | 0.066 | Born after 1996 | 0.375 |