Crises and Child Health Outcomes: The Impacts of Economic and Drought/smoke Crises on Infant Mortality and Birthweight in Indonesia

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CRISES AND CHILD HEALTH OUTCOMES:
THE IMPACTS OF ECONOMIC AND DROUGHT/SMOKE CRISES ON INFANT MORTALITY AND BIRTHWEIGHT IN INDONESIA

This paper examines the impacts of the recent Asian financial crisis on infant mortality and birthweight in Indonesia. There have been a number of economic and policy studies focusing on impacts of economic crises on finance and production. Although some studies provide evidence of negative impacts of economic crises on real outcomes, little is known about the impact of economic crises on child health outcomes such as changes in nutrition, child health, and mortality. Often, the association between financial and production disturbances and these outcomes are assumed (e.g. an adverse shock to production is thought to be associated with worse child health outcomes.). This paper utilizes data from the Indonesian Family Life Survey (IFLS) to examine impacts of the crises on child health outcomes directly. Specifically, we study the impacts of the crises on birthweight and infant mortality.

Prior to the Asian fianancial crisis, Indonesian rising levels of income, education, and public health programs had been successful in reducing infant mortality and improving the overall health of children. The infant mortality rate was reduced by more than half between the 1960s and the 1990s: from 145 deaths per one thousand live births in 1967 to 46 in 1997 (World Bank, 1999).[1] Table 1a shows neonatal mortality and post-neonatal mortality rates[2] prior to the crisis from the 1997 Demographic Household Survey. Data from this table shows that the rates of neonatal mortality and post neonatal mortality noticeably declined over the period of 15 years from 1982 to 1997. Indonesians experienced higher decline in post-neonatal mortality rate (from 31.0 to 23.9 per 1,000 live births) than in neonatal mortality rate (from 28.4 to 21.8 per 1,000 live births).

The Asian financial crisis struck Indonesia in January 1998.[3] Figure 1 shows that the sustained crisis period lasted more than one year with the peak in Rupiah/USD exchange rate in July 1998. As shown in Figure 2, the food prices in both urban and rural areas increased more than 250 percent at the peak of the crisis. This substantial increase in food prices is argued by Alatas (2002)[4] to be a major source of major impact of the crisis felt by Indonesians, except those that belong to the top of the income distribution. Simulation results from Alatas's study indicates that the drop in food prices between February 1999 and February 2000 accounted for approximately 40 percent of the decline in poverty rate after the crisis. According to Strauss et al (2002), food expenditures (excluding expenditures on tobacco and alcohol) accounted for approximately 50 percent

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[1] Daly, Patricia and Fadia Saadah. "Indonesia: Facing the Challenge to Reduce Maternal Mortality." East Asia and the Pacific Region Watching Brief. World Bank. June 1999. Issue 3.

[2] Neonatal mortality is defined as death before one-month old. 0-28 and 0-30 day periods are both used among researchers. This paper uses 0-30 days. Post-neonatal mortality is defined as death at ages 1 to 11.9 months. Infant mortality is defined as deaths at age 0 to 11.9 months.

[3] Refer to figure 1.

[4] Alatas, Vivi. "What Happen to Indonesia's Poverty? A Micro Simulation Exercise Using Household Surveys. Manuscript. World Bank. Jakarta, Indonesia. March 2002.

of a typical Indonesian's household budget in urban areas and 57 percent in rural areas. The food shares are higher among the poor. [5]

While an increase in food prices could help net food-producers, this large increase in the food prices during the crisis resulted in a sharp reduction of real incomes for most of the Indonesian population, who are net food purchasers. Frankenberg, Thomas, and Beegle in "The Real Costs of Indonesia's Economic Crisis: Preliminary Findings from the IFLS2+," (1999) [6] reported that The proportion of households below the poverty line rose from about 11 percentage points in 1997 to almost 20 percentage points in 1998.[7]

Overall government health expenditures per capita were not sustained at the peak pre-crisis level. Figure 3 shows real government health expenditures from 1980 to 2000. According to Leiberman et al. (2001),[8] the government per capita outlay on heath expenditures fell significantly during the crisis by 2.9 percent and by 6.6 percent in successive years. Then in 1999/2000, the expenditures rebounded. However, during the

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[5] Between 1997 and 1998 there was a significant increase in the food share in both rural and urban households. According to Frankenberg, Thomas, and Beegle (1999), the increase in food shares concentrated among households whose per capita expenditure was below the median of the population. A significant portion of the increase in food share can be attributed to an increase in the share of the expenditures on staples (from 13 percent to 21 percent for urban households and from 31 percent to 39 percent in rural households). As a result of a significant increase in the share of staples, some food shares such as that of meat are reported to decline. According to Strauss et al. (2002), nominal wages also increased during the financial crisis, but the increase was considerably less than the increase in food and nonfood prices. Therefore, real wages for those that rely on market wages also declined.

[6] Frankenberg E., Thomas D., and Beegle K. 1999. "The real costs of Indonesia's Economic Crisis: Preliminary Findings from the Indonesia Family Life Surveys." DRU-2064-NIA/NICHD. Santa Monica, CA: RAND.

[7] Allowed for higher inflation in rural than urban areas as indicated by the price data collected in the IFLS communities. (Frankenberg E., Thomas D., and Beegle K. 1999. "The real costs of Indonesia's Economic Crisis: Preliminary Findings from the Indonesia Family Life Surveys." DRU-2064-NIA/NICHD. Santa Monica, CA: RAND.)

[8] Lieberman, S., M. Juwono, and P. Marzoeki. "Government health expenditures in Indonesia through December 2000: An Update." World Bank East Asia and the Pacific Region Watching Brief. October 15, 2001, Issue 6.

crisis, Indonesia received a 278 percent increase in donor assistance, which contributed to the sustainability of health financing and spending. This donor assistance helped dampen the shock to the government budget, resulting in a five percent decrease in overall outlays[9].

According Frankenberg et al. (1999), prices of both public and private healthcare increased from late 1997 to late 1998, but public healthcare prices increased relatively more than the prices of private healthcare. The median price of BCG immunization for children and tetanus toxoid immunization for pregnant women rose significantly in public facilities,[10] but not in private facilities. There was a decrease in the quality of public healthcare such as a reduction of the quantity of drug given to patients and an increase in the number of referrals to other providers. In terms of supplies, public facilities were found to have been more affected by changes in the availability of drugs and supplies (such as injections and bandages) while private facilities have been more affected by the price increase of these inputs. Overall, public and private facilities that provided vitamin A declined in number. Vitamin A is essential for children under three since it reduces their vulnerability to infectious disease. Those above the median of the income distribution were found to shift away from public healthcare. Visits to the *posyandu* (healthcare post) by children under five dropped from 46.7 percent to 27.7 percent.

In addition to the financial crisis, some rural areas in Indonesia were badly affected by the 1997-1998 drought. The drought crisis was a consequence of climatic

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[9] Lieberman, S., M. Juwono, and P. Marzoeki. "Government health expenditures in Indonesia through December 2000: An Update." World Bank East Asia and the Pacific Region Watching Brief. October 15, 2001, Issue 6.

[10] The prices rose from 500 Rp. in 1997 to 750 Rp. in 1998 for Child immunization and from 500 Rp. in 1997 to 900 Rp. in 1998 for Tenanus Toxoid.

conditions identified with the El Nino Southern Oscillation (ENSO). The impacts of the ENSO on Indonesia included postponement of monsoons, damaging rainless winds, greater rate in infestations, disturbances in fishing patterns, scarcity of drinking water and forest fires. [11] In Kalimantan and Sumatra, coupled with a political decision on land clearing for plantation concessionaires, local populations, and new immigrants, the drought led to erupted fires that lasted for months between late 1997 and early 1998. The fires affected 9.75 million hectares and over 700 million tons of carbon were emitted into the atmosphere causing major health hazard in Indonesia, Singapore, and Malaysia.

An apparent result of the drought crisis was a sharp drop in food production, especially rice (Indonesia's major food crop). Figure 4 shows that the amount of land harvested for rice in late 1997 was smaller than usual. The harvesting cycle in late 1997 and early 1998 was delayed by as much as two months. It has been reported that some of the worst damage was to Maluku, Nusa Tenggara, and parts of Sulawesi and southwestern Sumatra.[12] Farmers in the driest and poorest provinces in Indonesia were reported at times to survive on one meal a day and often being forced to eat food generally reserved for livestock.[13] As a result of the drought crisis, Indonesia became a large food aid recipient in 1998.

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[11] James J. Fox. "The 1997-1998 Drought in Indonesia." Natural Disasters and Policy Responses in Asia: Implications for Food Security. Harvard University Asia Center. August, 1999.

[12] James J. Fox. "The 1997-1998 Drought in Indonesia." Natural Disasters and Policy Responses in Asia: Implications for Food Security. Harvard University Asia Center. August, 1999.

[13] James J. Fox. "The 1997-1998 Drought in Indonesia." Natural Disasters and Policy Responses in Asia: Implications for Food Security. Harvard University Asia Center. August, 1999.

According to Sastry in his study, "Forest Fires, Air Pollution, and Mortality in Southeast Asia,[14] smoky haze caused by a widespread series forest of fires in Indonesia between April and November 1997 had possible short-term and long-term effects on health. Possible problems that may have resulted in mortality include respiratory infections and chronic conditions. Sastry uses levels of micro-particulate matter with diameter less than 10 microns (PM_{10}) and the mean daily visibility in kilometer as measures of air quality to analyze the effects of smoke haze on mortality (non-traumatic, cardiovascular, respiratory deaths) of population of different age groups (all ages, <1, 65-74,and >74 years) in Kuala Lumpur and Kuching, Malaysia, in 1996-1997. Sastry found significant short-term cumulative effects of smoke haze on mortality in all age groups. Sastry claims that the displacement of deaths from the smoke haze was short-term, however, in one segment of the Malasian population -- those age 65-74 in Kuala Lumpur--- there was an upward shift in mortality that lasted at least a few weeks. Sastry suggests that an implication of his results on the short-term effects of the smoke haze in Malaysia is that the effects in Indonesia, where the main fires took place, are likely to have been large.

Evidence from IFLS2+ report suggests that the short-run impacts of the financial and the drought/smoke crises on children's health have been small. Results from physical assessments show no deterioration in children's health status. There were only negligible changes in the measurements of children older than six months in their height-for-age and weight-for-height. Very young children were well protected from the effects

[14] Sastry, Narayan. "Forest Fires, Air Pollution, and Mortality in Southeast Asia." Demography. Volume 39 number1. February 2002.

of the crisis although there is a suggestion that weight-for-height of this group of children may have worsened (Frankenberg, Thomas, and Beegle, 1999).

This paper uses IFLS data to examine the effects the crises on neonatal mortality, post-neonatal mortality, and birthweight. Mortality status and birthweight of children of different cohorts (born during pre-crisis, crisis, and post-crisis periods) are examined. In addition, detailed data on births and times of death of children allow this study to examine mortality hazard rates at different specific ages of children.

Estimated results on mortality outcomes show that the financial crisis had adverse impacts on neonatal mortality in both urban and rual areas. The adverse effects of the financial crisis on post-neonatal mortality risks were larger and more statistically significant for urban infants than for rural infants. Overall, the financial crisis increased infant mortality risks by about 3.2 percentage points in both urban and rural areas, a very large effects.

The drought/smoke crisis adversely affected post-neonatal mortality risks in rural areas. The increase in the post-neonatal mortality risk is about 3.1 percentage points. When community fixed effects are controlled for, the drought/smoke crisis had much larger effects. Overall, the drought/smoke crisis had no statistically significant advese effects on infant mortality in urban areas, while the effects in rural areas were large. Our estimates show that rural infants born during the drought/smoke crisis experienced approximately 4.4 percentage points increase in their infant mortality risks.

Results from hazard models confirm that the financial crisis and the drought crisis had adverse effects on neonatal and post-neonatal mortality. The financial crisis increased the odds of both neonatal and post-neonatal mortality for urban children more than for rural children. As expected, in rural areas, the crisis-noncrisis differential of the mortality risk at specific age (months) is relatively smaller for financial crisis than that of the drought crisis.

Our findings on differential effects on infants born to mothers with different levels of education show that in urban areas infants born to mothers with different levels of education exhibited significantly different trends in mortality risks over time. We also find that even though some of the crises had adverse effects on infant mortality, infants born to mothers with different education levels did not experienced different adverse crisis effects.

Results from the cumulative distribution comparisons of birthweights suggest that the financial crisis also had adverse impacts on birthweight in urban areas. However, under multivariate analyses, the adverse effect seems to disappear. None of the crises affected birthweights in rural areas. The lack of an evidence on the adverse effects maybe due to a selection problem in reported birthweights.

Conceptual Framework and Literature

The financial crises and the drought/smoke haze crisis represent short-term exogenous shocks to Indonesian households. Although specific biological mechanisms by which smoke haze may affect child health outcomes are not directly estimated,[15] both crises are expected to have negative consequences on child health through resource availability, income, prices, and environment.

According to agricultural household models (Singh, Squire, and Strauss (1986); Hill et al. (1993)), in a country with both agriculture and manufacturing sectors, an adverse agricultural shock not only reduces average resource availability but also affects the distribution of resources between the two sectors, largely through their effects on the relative prices of goods and services households consume. The direction of the effect depends on how the price of the agricultural good, expressed in terms of the manufactured good, responds to the shock.

With an increase in the price of agricultural goods after an adverse shock, agricultural household who are net producers may receive a greater amount of manufactured good for each unit of the agricultural good it does produce although each agricultural household produces less. However, if the price of manufactured goods also significantly increased during the crisis, these net sellers of agricultural goods could suffer from overall price increase.

While an increase in food prices could help net food-producers, households in sectors not benefiting either directly or indirectly from the favorable price movement such as net buyers of rice may face an erosion of their purchasing power as the relative prices of goods change. As noted earlier, a large increase in the food prices and an adverse income shock during the financial crisis resulted in a sharp reduction of real incomes for most of the Indonesian population, who are net food purchasers.

An important dimension of a reduction in purchasing power that is relevant to this paper is through the changes in the costs of raising children. These costs include both direct and indirect (opportunity) costs. When there is a reduction in purchasing power as

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[15] We do not estimate health production function in this paper.

a result of adverse income shocks or a rise in the price of agricultural goods, households are likely to reduce their consumptions such as that of health inputs and family planning services, provided that these goods and services are normal goods to the households. In addition, the effects of the crises on consumption could be worsened if there is an increase in healthcare costs and a reduction in availability of healthcare services, which are widely provided by the government.

It is, however, worth noting that a drought or a financial crisis in a particular year may have only a limited effect on the household's resources in that year if the household can transfer resources from another place or another time through the formal sector (such as bank loans or crop insurance) or the informal sector (such as loans or transfers from family or friends). However, if the shock occurs at the aggregate level, for instance a widespread financial crisis, the ability to transfer resources could be limited.[16] In this case, we expect to observe short-run negative effects of the adverse shock on health outcome such as in child mortality.

A large mainstream body of research on the effects of short-term economic fluctuations on demographic outcomes has used time-series data. The most influential work in this area is the study by Lee (1981) of the impacts of conditions in pre-industrial North and Western Europe using economic indices such as grain prices and weather.[17]

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[16] See Bardhan and Udry (1999) for detailed discussions.

[17] R. Lee, " Short-term variation: vital rates, prices and weather," in E.A. Wrigley and R.S. Schofield (eds.), The Population History of England, 1541-1871 (Cambridge, Mass: Harvard University Press, 1981), P. Galloway, "Basic Patterns of annual variations in fertility, nuptiality, mortality, and prices in pre-industrial Europe," Population Studies, 42 (1988), pp. 275-302, D. Weir, "Life under pressure: France and England, 1670-1680," Journal of Economic History, 44 (1984), pp. 27-47. T. Richards, "Weather, nutrition, and the economy: the analysis of short-run fluctuation in births, deaths, and marriages, France 1740-1909," in T. Bengtsson et al. (eds.), Pre-Industrial Population Change (Stockholm: Almquist and Wiksell, 1984).

The evidence from this study shows that Post-infant mortality is positively associated with real prices. In a similar study, Galloway (1988)[18] found that a ten percent increase in grain prices in pre-industrial Europe leads to a decrease of approximately 1 percent increase in mortality.[19]

A number of authors have tried to study the effects in less developed countries.[20] A major effort to assess the effect of economic fluctuations on demographic outcomes in developing countries is a collective work by the National Research Council.[21] In this study, first marriage, timing of first and second birth, and child mortality in seven countries in Sub-Saharan Africa were examined using data from the Demographic Health Surveys (DHS) that were conducted in each of the seven countries sometime in the period of between 1986 and 1990. In this paper, various economics indicators are used as measures of economic fluctuations. The indicators used are the gross domestic product per capita, the quantity of exports, term of trade, and commodity prices. The authors find

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[18] Galloway, P. "Basic pattern in annual variations in fertility, nuptiality, mortality, and prices in pre-industrial Europe." Population Studies 42(2): 275-302. 1988.

[19] Although he observed that the relationship between prices and mortality in countries that are more developed economically is weaker.

[20] J. Bravo, "Economic Crisis and mortality: short and medium-term changes in Latin America." Paper presented at the Conference on the Peopling of the Americas, Veracruz, Mexico (1992). J. Brovo, "Demographic Consequences of structural adjustment in Chile." Paper presented at the Seminar on Demographic Consequences of Structural Adjustment in Latin America, Ouro Preto, Brazil (1992). K. Hill and A. Palloni, " Demographic responses to economic shocks: The Case of Latin America." Paper present at the Conference on the Conference on the Peopling of the Americas, Veracruz, Mexico (1992). A. Palloni and K. Hill, "The Effects of Structural Adjustments on mortality by age and cause in Latin America." Center for Demography and Ecology. Working Paper 92-22, University of Wisconsin, Madison, Wis. (1992). D. Reher and J.A. Ortega, " Short run economic fluctuations and demographic behaviour: some examples from twentieth century South America." Paper presented at the Seminar on Demographic behaviour: some examples from twentieth century South America." Paper presented at the Seminar on Demographic Consequences of Structural Adjustment in Latin America, Ouro Preto, Brazil (1992).

[21] Hill, K, G. Adansi-pipim, L. Assogba, A. Foster, J. Mukiza-gapere, and C. Paxson. " Demographic Effects of Economic Reversals in Sub-Saharan Africa. National Research Council.National Academy Press. Washington D.C. 1993.

strong evidence of adverse effects of economic reversals on time of first marriage and first births. [22] They, however, found no effect of the economic reversals on child mortality net of trend, except in Ghana (in rural areas) and Nigeria.

Palloni, Hill, and Aguirre[23] employed distributed lag models and used average GDP to study the effects of short-term economic fluctuations on marriage, marital fertility and mortality during 1920-1990 in Latin America. They found the directions of the response of the number of marriages are not uniform across countries. The net effects (the sum of all lagged coefficients) are positive in all cases, except Guatemala and Mexico, while the magnitudes of the net effect elasticities vary across countries, from 0.01 (in Mexico) to 0.67 (in Chile). Their results from the estimated responses of births show considerable greater heterogeneity. In five out of eleven countries, the coefficients for lags 0 and 1 of the average GDP are positive as expected, but only lag1 in Cuba is significantly different from zero. The estimated net effects are positive in only seven countries, and the magnitude of the net effect elasticity ranges from 0.04 (in Chile) to 0.88 (in Cuba). The effects on infant mortality rate also display considerable heterogeneity in terms of direction of the effects. Results from only five countries are negative, as expected, and the estimated elasticity of the net effects ranges from 0.08 (in El Salvador) to 0.61 (in Panama). The results from pooled sample estimations using pooled sample (except Cuba) give estimated elasticity of the net effects of 0.12, 0.19, and -0.12 on births, marriages, and infant mortality rate respectively. An important finding as

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[22] Their results are strongest for the effects on first birth. The authors found a positive relation to economic variation net of trend in all seven countries studied, except Kenya. Evidence of marriage delays as a result of economic reversals were found in Botswana, Senegal, and Togo (especially in urban areas).

[23] Palloni, A, K. Hill, and G.P. Aguirre. "Economic Swings and Demographic Changes in the History of Latin America. Population Studies, 50 (1996), pp. 105-132.

suggested by the authors is that the effects of the lag0 and lag1 on infant mortality are significant, but their absolute size are small (-0.10, and 0.00), about one-third that of the effects on marriages and births.

Mckenzie (2002), in his study of how Mexican households coped with aggregate shocks of the Peso crisis of 1994-1996, finds lesser role of inter-household transfers in consumption smoothing in presence of aggregate shocks. He finds that the average transfer that household made to non-household members was reduced by 25 percent. Although remittances from friends and family members in the United States to Mexican households increased, on average the households received 19 percents less gifts and donations from other Mexican households.

Another pioneer empirical study in this area is the study of child mortality of children born or conceived during the Dutch Hunger Winter of 1944-45 by Stein *et al*.[24] This study uses vital registration records to compare age-specific mortality rates for different cohorts. The authors found an excess mortality of children born or conceived during the famine crisis. However, due to data limitations, the authors control for only social status (manual/non-manual occupation).

In their study of the effects of the 1974-75 famine in Bangladesh, Razzaque *et al*. (1990)[25] address the above data issues using richer data from Matlab field research in a rural area in Bangladesh. Mortality of children of different cohorts, born during famine, post-famine, and non-famine periods, is examined using various socioeconomic controls

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[24] Stein, Z., M.Susser, G. Saenger and F. Morolla, "Famine and Human Development: The Dutch Hunger Winter of 1944-1945, New York (1975).

[25] Razzaque, A., N. Alam, L. Wai, and A. Foster. "Sustained effects of the 1974-75 famine on infant and child mortality in a rural area of Bangladesh." Population Studies 44: 145-154. 1990.

such as child's gender and household economic status. Using linear logistic regression models, they found excess mortality of children born or conceived during the famine. However, Razzaque and his colleagues found that the effects of the famine on mortality were not uniform. Better off household experienced little effect.

The incidence of the recent Indonesian financial crisis and the 1997/1998 drought crisis and the availability of IFLS data allow this study to address the affects of larger exogenous short-run shocks on demographic outcomes. This paper uses IFLS data to examine the effects the crises on neonatal mortality, post-neonatal mortality, and birthweight. By using IFLS3 and earlier IFLS waves (IFLS1 and IFLS2), mortality data on child cohorts that were born during pre-crisis, crisis, and post-crisis periods can be examined. In addition, we can allow for the time trends that are independent of the crises when studying these adverse effects. Since the IFLS covers respondents in both rural and urban areas of Indonesia, the area-specific impacts of the crisis can also be studied. In addition, detailed data on births and times of death of children allow this study to examine mortality hazard rates at different specific ages of children.

A Theoretical Model

The theoretical model in this paper is an adaptation of the work by Foster (1995).[26] Consider a household j. Abtracting from fertility selection, assume parents care about the health of their surviving children, but not their individual consumption. Denote

[26] Foster, Andrew D. "Price, Credit Markets and Child Growth in Low-Income Rural Areas." The Economic Journal, Volume 105, Issue 430 (May, 1995), 551-570.

t the beginning of period t. Let h_{it} be the health status of the child i at time t and C_t a vector of the household's consumption of goods and services other than those that are inputs in the production of children's health. Single period household's utility function is assumed to additively separable between consumption and child's health. The utility function is also assumed to be increasing and concave in household consumption and child's health status ($v'(C_t) > 0, v'''(C_t) < 0 \; and \; u'(h_{jit}) > 0, u''(h_{jit}) < 0$).

The expected discounted utility of household j at time s is

$$V_s = E_s \sum_{t=s}^{T} \beta^{t-s} \left[v(C_t) + \sum_{i=1}^{I} u(h_{it}) \cdot M_{it} \right],$$

where E_s the expectation conditional on information at time s, β is the discount factor, and I is the total number of children in the household, and the subscript j is dropped for notational simplicity. $u(h_{it})$ is the utility of the household from the health of the child i, h_i. M_{it} is the child i's mortality index, where M_{it} =1 if the child is alive in period t and M_{it} = 0 if the child is not alive in period t. In every period t, the child dies with certainty when his health, h_{it}, is below a heath threshold, h_{it}^*.

The production function of the child's health status in the current period is characterized as a function of the child's health status in the last period, augmented by health inputs in the last period. Let h_{it} be the child's health status at time t given characteristics at the individual level (such as age and gender), household level (such as education of the mother), and community level (such as quality of water supply), z_{it}, and unobserved factors that affect health, δ_{it}. Following Foster (1995), the effects of health inputs, N_{it}, in the last period on the current period health status is assumed to be proportional to a time-varying rate at which health inputs are translated into health status gain. This rate is a function of the child's characteristics and unobserved factors δ_{it}. Parents are assumed to know z_{it} with full certainty. The health status of the child at time $t+1$ can be characterized as

$$h_{i\,t+1} = f_{it}\left(h_{it}; z_{it}, \delta_{it}\right) + k_{it}\left(z_{it}, \delta_{it}\right) N_{it},$$

where $\frac{\partial f_{it}\left(h_{it}; z_{it}, \delta_{it}\right)}{\partial h_{it}} > 0$ and the Mortality index can be characterized as

$$M_{it} = \begin{cases} 1 & if \quad h_{i\,t}\left(h_{it-1}, N_{it-1}, z_{it-1}, \delta_{i-1t}\right) > h_{it}^{*}\left(z_{it}, \varepsilon_{it}\right) \\ 0 & Otherwise \end{cases},$$

where ε_{it} is unobserved factors, unknown to the household, that affect the mortality threshold at time t.

Assume no inter-temporal borrowing and lending, the household's budget constraints at time t is

$$p_t^C C_t + p_t^N \sum_{i=1}^{I} N_{it} = y_t$$

By the Law Iterated Expectation, the expected discounted utility of household j at time s can be written as

$$V_s = E_s \sum_{t=s}^{T} \beta^{t-s} \left[v(C_t) + \sum_{i=1}^{I} \left(u(h_{it}) .\ prob(M_{it} = 1 \mid t-1)\right) \right],$$ [27]

[27] Proof:

Let function g_{it} characterizes mortality risk of the child i in period t;

$$g_{it}(h_{it}, z_{it}, \varepsilon_{it}) = prob(M_{it} = 0) = prob\left(h_{it}(h_{it-1}, N_{it-1}, z_{it-1}, \delta_{i-1t}) < h_{it}^{*}(z_{it}, \varepsilon_{it})\right),$$
$$where\ g_1' < 0\ and\ g_1'' > 0$$

Suppose there is a temporary adverse income shock at period 0, either C_0 or N_{i0} or both will be reduced in order to satisfy the budget constraints at time 0. If children's health inputs are considered normal goods, N_{i0} will be reduced as a result of a reduction in income. Supposed N_{i0} is reduced by dN_{i0}, then the health status of the child i in period 1 will be reduced by $k_{i0}(z_{i0}, \delta_{i0})\, dN_{i0}$. Furthermore, this temporary adverse income shock will have spillover effects on future health status of the child in every period after period 0 as the health status at any period t depends on the health status in period $t-1$ in the health production function. Specifically, a decrease in N_{i0} by dN_{i0} will reduce the health status of the child i in period t by

$$\frac{\partial h_{it}}{\partial N_{i0}}.dN_{i0} = \left(\prod_{s=1}^{t} \frac{\partial f_{is}(h_{is}; z_{is}, \delta_{is})}{\partial h_{is}}\right).\, k_{i0}(z_{i0}, \delta_{i0})\, dN_{i0}\,.$$

$$\begin{aligned} V_s &= E_s \sum_{t=s}^{T} \beta^{t-s} \left[v(C_t) + \sum_{i=1}^{I} u(h_{it}) \cdot M_{it} \right] \\ &= E_s \sum_{t=s}^{T} \beta^{t-s} v(C_t) + \sum_{t=s}^{T} \beta^{t-s} \sum_{i=1}^{I} E_s\left[u(h_{it}) \cdot M_{it}\right] \end{aligned}$$

By the Law of Iterated Expectation,

The effect of a temporary adverse income shock on mortality risk

An adverse shock to income in period 0 that results in a change of health inputs by dN_{i0} will have adverse effects on the mortality risk through a reduction of health input by

$$\frac{\partial\ prob(M_{it}=0)}{\partial y_0}=\frac{\partial g_{it}}{\partial h_{it}}\frac{\partial h_{it}}{\partial N_{i0}}\frac{\partial N_{i0}}{\partial y_0}=\frac{\partial g_{it}}{\partial h_{it}}\left(\prod_{s=1}^{t}\frac{\partial f_{is}(h_{is};z_{is},\delta_{is})}{\partial h_{is}}\right).\ k_{i0}(z_{i0},\delta_{i0})\,dN_{i0}$$

Measuring Mortality Rates

There are two principal estimation methods to calculate mortality rates: direct methods and indirect methods. (Spiegelman, 1955;[28] Pressat, 1978[29]) Direct methods calculate mortality directly using data on the date of birth of children, survival status, and the dates of death or ages at death of deceased children.

The direct methods require data that are usually only obtained in specifically designed surveys with birth/pregnancy histories or from vital statistics systems. There

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$$\begin{aligned}V_s&=E_s\sum_{t=s}^{T}\beta^{t-s}v(C_t)+\sum_{t=s}^{T}\beta^{t-s}\sum_{i=1}^{I}\ E_s\left[E_{t-1,h_{it}}\left(u(h_{it})\cdot M_{it}\right)\right]\\&=E_s\sum_{t=s}^{T}\beta^{t-s}v(C_t)+\sum_{t=s}^{T}\beta^{t-s}\sum_{i=1}^{I}\ E_s\left[u(h_{it})\cdot prob(M_{it}=1\mid t-1)\right]\\&=E_s\sum_{t=s}^{T}\beta^{t-s}\left[v(C_t)+\sum_{i=1}^{I}\ \left(u(h_{it})\ .\ prob(M_{it}=1\mid t-1)\right)\right].\end{aligned}$$

[28] Spiegelman, M. Introduction to Demography. 1955. The Society of Actuaries. Chicago. Illinois.

[29] Pressat, R. *Statistical Demography*. Translated and Adapted by Damein A. Courtney. 1978. Methuen & Co Ltd.

are three variants of direct estimation methods: a vital statistics approach, a synthetic life table approach, and a true cohort live table approach.

"A vital statistics approach" is an approach in which the number of deaths to children under age 12 months in a particular period is divided by the number of births in the same period. Under "a synthetic cohort life table approach," mortality probabilities for small age segments based on real cohort mortality experience (e.g. 0, 1-2, 3-5, 6-11 months) are first calculated. Then these component death probabilities are combined into the mortality rates, taking into account exposure to mortality risk of each age cohort. Specifically, component death probability for each small age segment is calculated by dividing the number of deaths to live-born children during specified age range and specified time period by number of surviving children at beginning of specified age range during the specified time period. Births and death incidences within each specific age and specific time ranges are weighted according to exposure to mortality risk, which is indicated by birthdate and mortality risk of interest (e.g. neonatal mortality). This approach allows full use of the most recent data and is also specific for time periods. However, this approach requires intensive computation and the number obtained using this method could be influenced by arbitrary length of age segments used.[30]

Under "a true cohort life table approach," the number of deaths to children under age 12 months of a specific cohort of births are divided by the number of births in that cohort. This method gives true probabilities of death but requires that all children in the

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[30] Studies done during the World Fertility Survey have shown the difference to be negligible when using monthly segments and using 0, 1-2, 3-5, 6-11, 12-23, 24-35, 36-47, 48-59 months segments. (Demographic Health Survey).

cohort must have been fully exposed to mortality risk.[31] Although this method does not take into account the most recent experience because of exposure requirement, it is the method we chose in this paper since this method gives true probability of death while full exposure problem is mitigated by using IFLS3, which includes information on children born in the crisis. Most of these children had had full exposure to neonatal and post-neonatal mortality risk by the time of the interview in 2000. Moreover, the problem is less severe because information of children born in the IFLS survey year (who had less than one year mortality exposure) can be added using data from subsequent IFLS waves so that all of these observations are fully exposed to infant mortality.[32]

An "indirect method" is used to calculate death rates when age-specific death rates for the community are not available, but the total number of deaths is known. By this method, the number of actual age-specific deaths in the community is multiplied by a constant number of the age-specific death rates[33] in a standard population that are usually derived from European experience, which are referred to as the "*standard mortality schedule.*" The result yields the adjusted death rates by the indirect methods. (Spiegelman, 1955; Pressat, 1978).[34]

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[31] This requirement of full exposure becomes more limiting the higher the age segment of interest. For example, to calculate under-five mortality rates, only information on children born at least five years before the survey can be used.

[32] Please refer to data section for detailed discussions. Later in this paper, monthly hazard rate estimations are performed to avoid the problem of full exposure to infant mortality risk.

[33] The number of deaths between ages x and x+n among residents in a community during a year divided by the average number of persons between ages x and x+n living in that community during the year, multipled by 1,000. (Spiegelman, 1955)

Under indirect methods, mortality rates are calculated using only the number of children ever born, the number of living children to women, and ages of women. The indirect methods can utilize data that are commonly collected in censuses and many general surveys.[35] Although these indirect methods can utilize data that do not contain detailed information on date of birth of children and the dates of death or ages at death of deceased children, these methods are based on an implicit assumption that the births of a cohort of a women represents all children born in a time period.[36] Another problem with indirect methods is that the indirect methods estimate the probability of dying based on experience that can extend over many years, resulting in an average over that period. The methods are subject to error when there are changes in fertility and mortality trends, which occured during 1960s -1990s in Indonesia.

Note that both direct and indirect methods could suffer from reporting error due to the omission of deceased children. Estimation of infant mortality using direct methods also depends on the correct reporting of age at death as under or over one year. The heaping of deaths at age 12 months is also common, and to the extent that it causes a transfer of deaths across the one-year boundary, infant mortality rates may be somewhat underestimated.

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[35] "Because many late fetal deaths and neonatal deaths may be attributed to the same underlying conditions, it has been proposed to combine the two to compute a "perinatal rate," where the deaths are divided by either live births alone, or the sum of live births and fetal deaths. Since there is no generally accepted convention for computing this rate, the terms entering into it should be defined wherever it is used." (Spiegelman, 1955). One can conclude that although the indirect methods can be used to extract other mortality rates from census data, perinatal rates cannot be computed because census data give only birth history, where only live births are recorded.

[36] Documentation by the Demographic Health Survey reports that recent and on-going work shows that this assumption may not be valid: births to women 20 to 24 (and in some cases to women 25 to 29) have more elements of high risk of mortality than do all children born within the last five years of a survey.

Data

This paper uses data from the Indonesia Family Life Survey (IFLS). IFLS is a continuing longitudinal socioeconomic and health survey that includes more than 30,000 individuals living in 7,200 households. The sample covers 321 communities in 13 provinces in Indonesia and represents about 83 percent of the Indonesia population in 1993[37]. The first wave of IFLS was fielded in 1993 (IFLS1). The same households were revisited in 1997 (IFLS2) and again in 2000 (IFLS3). This paper the data from these three IFLS waves. A 25 percent sub-sample of households was re-interviewed in 1998 (IFLS2+), but the data are not used in this paper.

In IFLS surveys, special attention is paid to the measurement of health, work, migration, marriage, child bearing, life history data on education, and economic status of individuals and households. In each wave of IFLS, the individual and household surveys are complemented by an extremely comprehensive community and facility survey. There is also considerable attention placed on minimizing sample attrition in IFLS. Targeted households and individuals who "split-off" from original households were followed if they moved to new locations within 13 provinces of the survey areas. In each re-survey, about 95 percent of targeted households have been re-contacted. The split-off households added just under 1,000 households to the sample in 1997 and about 2,600 households in 2000.

The survey periods of each IFLS wave is shown in Figure 1. Data on pregnancies, birth outcomes, infant mortality, and age of death are obtained from retrospective data on pregnancy histories. Complete pregnancy histories were given by women in the household who were between 15 and 49 years of age. These women were asked about information of each pregnancy in detail. This information includes pre-natal care, pregnancy outcome, birth information, post-natal care, and survival status of the child. For prenatal-care, complete information on frequency and type of pre-natal care of each trimester of pregnancy was obtained. For birth information, the women were asked detailed information about their pregnancy outcomes. In the case where there was a miscarriage, the length of time before the pregnancy ended was reported. If the pregnancy resulted in a live birth, information was obtained retrospectively on length of pregnancy, birth date, place of birth, healthcare provider at birth, whether the child was weighed at birth, and birthweight (if the child was weighed). For the child's survival status, mothers were asked if their children were alive at the interview date. If the child died before the interview date, a complete history of the child's death was obtained. This information includes how old the child was when the child died (in days, weeks, months, or years).

Data are obtained from IFLS1, IFLS2, and IFLS3, but are restricted to include only children that were born between 1988 and 2000. This is because we want to restrict the recall to five years for each wave in order to minimize recall error. This is similar to strategies used by demographers.[38]

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[37] Frankenberg, E., Hamilton, P., Polich, S., Suriastini, W., and D. Thomas. User's Guide for the Indonesia Family Life Survey. DRU2238/2-NIA-NICHD. March 2000.

[38] For instance, such as in the Demographic Health Surveys report infant and child mortality based on events in the previous four or five years before the survey.

As IFLS2 and IFLS3 follow the respondents of the original IFLS1, birth data are organized as follows:

If the respondent (mother) was previously interviewed and reported some children in earlier IFLS wave(s), the mother reported only children born since the last reported child in the new IFLS wave.

If the respondent was interviewed in the previous wave, but reported no children then, the mother reported all children born after the last interview in the new interview, which amounts to a complete history

If the respondent was a new respondent, either because she turned 15 or older in the new survey or because she was a new household member, a complete history was taken.

As a result, birth data of children born in any particular year could be from any of the IFLS waves. Information on children born during the financial crisis period (1998-1999) can be obtained only from the IFLS3 wave, except for a few children born in January 1998 when IFLS2 was still taking place.

In addition, observations across IFLS waves for panel respondents with preprinted roster are carefully compared to avoid duplication, which occurred occasionally. According to discussions of the quality of retrospective data on longitudinal surveys,[39] retrospective data are of better quality if the length between the real event and the interview date is minimized. Therefore, in the case of duplication, only the observation

[39] Beckett, M., J. DaVanzo, N. Sastry; C.Panis, and C. Peterson. "The quality of retrospective data: An examination of long-term recall in a developing country." Journal of Human Resources. Summer 2001.

This paper on reporting error provides insights into the quality of retrospective reports, particularly as it pertains to short-term recall. Studies were reviewed which analyzed the quality of retrospective reports in from the earlier IFLS wave is included. An exception is made for 1997 and 1998 births. Since these children were not at least one year old by IFLS2 interview date, we cannot extract uncensored mortality information from observations that were reported in IFLS2.[40] In this case, data from IFLS3 are used where duplication occurs. Refer to Appendix I for details.

To identify neonatal mortality and post-neonatal mortality, only live births were included. As previously discussed, this paper uses a "true cohort-live table approach" to calculate neonatal and post-neonatal mortality rates. This approach requires that the children in our sample had full exposure to mortality risks, which, in our case, is one year old.[41] Instead of dropping observations that did not have one-year full exposure to mortality risks by the interview date in IFLS1 (1993) and IFLS2 (1997), an effort was made to recover these observations. For those who were younger than one-year old by the interview date in these earlier IFLS waves, information on their survival is obtained from the household roster of the subsequent IFLS using identification numbers that are consistent throughout all waves of IFLS.[42] This household roster provides information on whether the child was still living and the time of death if the child had died. In the case where we could not track a particular child from IFLS1 in IFLS2, we also looked for him/her in IFLS3. With a high re-survey rate of about 95 percent, we were able to

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the Malaysian Family Life Surveys (MFLS), fielded in Peninsular Malaysia in 1976 and 1988, and conclude that many of the data quality problems found previously are present in the MFLS.

[40] Some respondents of the IFLS1997 were interviewed in January and February 1998.

[41] This full-exposure restriction is not our concern when studying monthly hazard rates later in the paper.

[42] When the child did not live in the same household in subsequent IFLS survey, survival and date of death (if died) information was obtained from the new household.

recover most of the observations. As a result, we are able to add 487 children born in 1993/1992 and 511 children born in 1996/1997 to our sample.

In addition to data on pregnancy and survival histories, data from individual, household, and community characteristics surveys are used to allow this study to control for other socioeconomic characteristics such as child's gender, mother's education, urban/rural areas of residence, and province of residence.

The residence of mothers we use in this paper is the residence of mother at the time of the survey. We are aware of a general practice of taking the residence of females at the time when they were 12 years old or residence before the first marriage in feritity and marriage studies, but in this study of the effects on mortality, the residence of mothers after their marriages are thought to affect healthcare provided to their children more than pre-marital residence in terms of type and quality of the healthcare. We take the residence of the mother at the time of the survey to be a proxy for the post-marital residence.

The data are linked using individual and household identifications, which are consistent throughout different surveys of each IFLS and different waves of the IFLS.

Method of Analysis

Testing Different Measures of Mortality Rates

The first goal of this paper is to compare mortality rates from IFLS data to 1997 Indonesian Demographic Household Survey (DHS) data. We use a "true cohort live table" approach and compared the rates to mortality rates published in the Indonesian DHS 1997 publication, that uses “a synthetic cohort life table” approach. We then compare mortality rates using DHS and IFLS data, but using the “true cohort live table” approach with both datasets.

Table 1a, taken from the Indonesian DHS publication, shows neonatal and post-neonatal mortality rates that are calculated using the “synthetic cohort life table” approach. From the same dataset, table 1b shows computed neonatal mortality and post-neonatal mortality rates from the 1997 DHS using the “true-cohort-live table”. Data are divided into three different periods by birth date (1982-1987, 1987-1992, 1992-1997) to match DHS publication numbers. Mortality rates are calculated with and without individual sampling weights provided in DHS.

Table 1b shows that across different time periods, weighted and unweighted mortality rates are similar, and hence suggesting that the DHS observations give a good representation of the Indonesian population. When comparing results from different estimation approaches, weighted mortality rates are used since the DHS’s rates are weighted. The comparison results show that different estimation approaches yield similar rates in all three periods for both neonatal and post-neonatal mortality rates. Regardless of methods used, we can conclude that Indonesia experienced a decrease in both neonatal and post-neonatal mortality rates during the 10 years period prior to the economic and drought/smoke haze crises. The decline is not as sharp for neonatal mortality as it is for post-neonatal mortality.

When comparing the mortality rates of full DHS samples and the mortality rates of DHS observations that are only from the 13 provinces surveyed in the IFLS, both neonatal and post-neonatal mortality rates drop just by a small magnitude for both weighted and unweighted rates. The mortality rates are smaller in the 13 IFLS provinces across all periods. This corresponds to the provinces not surveyed being from the poorer, eastern provinces.

Table 1c presents neonatal and post-neonatal mortality rates from IFLS data using the "true-cohort-live-table" approach. Mortality rates are calculated for the corresponding DHS periods. When using both IFLS1 and IFLS2, IFLS data exhibits higher post-neonatal mortality rates than DHS rates for 1982-1987 and 1987-1992. The IFLS post-neonatal mortality rate is, however, lower for 1992-1997. The differences in the neonatal mortality rates between the two datasets are smaller than the differences in the post-neonatal rates for 1982-1987 and 1987-1992, but larger for 1992-1997.

When comparing mortality rates using only IFLS for consistency check, we find that including data from a more recent wave of IFLS gives higher rates for both neonatal and post-neonatal mortality. The data suggest that when longer retrospective is used, mortality rates are higher.

Birthweight Reporting

When studying the effect of the crises on health outcome such as birthweight, a crucial concern regarding the data is whether there is any selection problem in reported outcomes. Children born at home are likely to be from rural, poorer households, and less educated mothers. These children are likely to have lower birthweight, which are likely to be unreported. If this selection problem is present in the birthweight distribution, then reported numbers are biased. Table 2 shows that in Indonesia while almost all births that took place in hospitals or community health center/deliver posts have reported birthweight, only 51.6 percent of births that took place at home have reported birthweight. This proportion of reported birthweight is close to the proportion of reported birthweight from office or house of traditional midwives. Note that when a woman gave birth at home or at family members' house, often a traditional midwife was called in to assist in the delivery process. Therefore, some proportion of babies that were delivered at home could be weighed if the midwives were well trained and well-equipped with measuring devices.

Table 3 shows proportion of birth location by different time periods. On average, only approximately 15 percent of deliveries took place in public or private hospitals. Health centers, village delivery posts, clinics of physician, and clinics of formally trained midwives serve as major delivery facilities in Indonesia (26.7 percent). The proportion of births that took place in these formal facilities increased during 1988-2000 from 21.0 percent to 26.7 percent. On average, birth deliveries that took place at home account for more than half of all deliveries in Indonesia. Although the proportion of deliveries that took place at home decreased substantially over time, the proportion remained relatively high in 1998-2000 (49.3 percent). From these results, we expect to encounter problems of reporting birthweight. This is due to a higher proportion of births that took place at home. Unfortunately, we do not have any plausible instrument to correct for this selection problem in reported birthweight. It is difficult to find a factor that affects delivery location (and whether birthweight was reported), but does not affect birthweight itself.

Mortality and Birthweight Statistics

The financial crisis periods are divided into two parts: crisis 1 and crisis 2 (refer to Figure 1). Crisis1 marks the period when the Rupiah rapidly devalued, the exchange rates were extremely volatile, and food prices accelerated (January 1998 to September 1998). Crisis 2 marks the period when the exchange rates began to settle and food prices came down. (October 1998 to June 1999). From earlier discussions, the 1997/1998 drought crisis and the smoke haze crisis covers the period from May 1997 until early 1998 in some areas. According to the timing of these different crises, the sample periods are divided into four sub-periods: *non-crisis* (January 1988 to April 1997 and July 1999 to December 2000), *crisis 1* (January 1998 to September 1998), *crisis 2* (October 1998-June 1999), and *drought 97* (May 1997 to December 1997). By this period grouping, *crisis1* also covers some of the drought/smoke haze crisis in some areas of Indonesia. [43]

In addition, since crisis 1 and crisis 2 each lasted nine months, we can interpret children who were born in crisis 2 as those who were conceived during crisis 1. Those that were born during crisis1 in rural areas that were affected by the drought could also be roughly interpreted as those who were conceived during the drought period.

An overview of mortality rates during our periods of interests is shown in Figure 5. Figure 5 shows mortality rates between 1988-1999. One can observe that the post-neonatal mortality rate gradually declined until 1996. Then the rate went up in 1997, further up in 1998, and started to decline in 1999. Neonatal mortality rates were more volatile with a downward trend until 1996. Neonatal mortality rate slightly increased in 1997, but the rate

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[43] Although harvesting cycle in early 1998 was delayed by as much as two months, most of the effects of the effects of the drought crisis on productions had already been felt in the second half of 1997.

was still lower than those of the year prior to 1996. The rates heightened in 1998, and, similar to post-neonatal mortality rate, neonatal mortality rate started to decline in 1999.

The mortality status at the end of the first year of children born in crisis and non-crisis periods is shown in Table 5. As we expect from Figure 5, mortality was at a relatively high level up to 1992 and started to decline in 1993. The non-crisis periods are divided into two periods: "1988-1992" and "1993-1997, 1999."[44]

Table 5 shows that both neonatal mortality and post-neonatal mortality rates are higher in rural areas than in urban areas in all periods. Infant mortality significantly increased from the pre-crisis period (32.5 per 1,000 live births) during both the financial crisis (46.2 per 1,000 live birth) and the drought/smoke haze crisis (76.1 per 1,000 live births).

Neonatal mortality and post-neonatal mortality rates of children born during the financial crisis are higher than those of children born during non-crisis period of 1993-1997 and 1999, but lower than those of the children born during 1988-1992. Overall, for those born during the crisis, about 23 infants out of 1000 died within one month of birth as compared to the pre-crisis rate of about 13 per 1000 live births. The percentage increase of neonatal mortality is higher in urban areas (from10.2 to 20.5 per 1000 live births) than in rural areas (from 15.7 to 25.4 per 1000 live births). For overall sample, those born during the financial crisis also exhibit higher post-neonatal mortality than in the pre-crisis period. However, the percentage increase in the post-neonatal mortality is smaller than the percentage increase in neonatal mortality rate. When looking at rural and urban samples separately, the data show that urban children experienced a large increase in the post-neonatal mortality rate (from 10.9 to 16.4 per 1000 live births) while rural children experienced only a small increase in the rate (from 25.8 to 29.0 per 1000 live births).

The drought/smoke haze crisis seems to have had strong negative impacts on both neonatal mortality and post-neonatal mortality in rural areas where the largest negative impacts are expected. Compared to the pre/post-crisis period of 1993-1997 and 1999, Neonatal mortality increased from 15.7 to 29.0 (per 1,000 live births), while post-neonatal mortality increased from 25.8 to 47.1 (per 1,000 live births). One can observe that the post-neonatal mortality rate increased during the drought/smoke haze crisis so much that it surpassed the post-neonatal mortality rate of 1988-1992 period (38.7 per 1,000 live births). This evidence suggests that the drought crisis had stronger adverse effects on mortality than the financial crisis in rural areas. However, it is worthwhile to recognize that the mortality rates shown in Table.4 are calculated from small samples, especially the mortality rates for the crisis periods. As a result, these rates (per 1,000 live births) are sensitive to the number of death incidences among these small samples.

Birthweight statistics are shown in Table 6. Overall, the financial crisis had a small negative effect on birthweight. The proportion of low-birth-weight children (children with birthweight less than 2.5 kilograms) increased from 8.1 percent during non-crisis 1993-1997 to 8.7 percent during the financial crisis. The data also show some reduction in the mean birthweight with the biggest decline among urban children (from 3.17 to 3.11 kg.). The drought/smoke haze crisis appears to have a negligible adverse effect on the mean birthweight, but the percentage of low-birth-weight children is lower

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[44] Births between July 1999 and December 2000 also belong to the non-crisis period, but they are not

than that of the non-crisis periods in both urban and rural areas. Since we expect to encounter selection problems in reported birthweight as discussed earlier, the estimated effects of the crises on birthweight is subject to errors. For instance, if during a drought/smoke crisis a high proportion of mother shifted away from public healthcare to in-home care or offices of traditional midwife for delivery, it is more likely that babies were not weighed. Table 4 shows the proportions of delivery locations during pre-crisis, crisis, and post-crisis periods. The data suggest that during both crises, there is a jump in the proportion of deliveries that took place at home or at offices of traditional midwife from a decreasing trend. Babies born during the crises are, therefore, less likely to be weighed. If the unreported babies weighed less than the reported mean birthweight (e.g. mothers who switched to deliver at home because of high costs of public healthcare were the ones endowed with poor health), then the mean reported birthweight during the drought/smoke crisis understimates the adverse crisis impact. On the other hand, if unreported babies weighed more than the reported mean birthweight, (e.g. mothers who were healthy decided to save some money by giving birth at home), then the mean reported birthweight overestimates the adverse crisis impact. The direction of the bias is certainly an empirical one.

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included because those children were not yet exposed to one year mortality by the interview date.

Testing the Effect of Economic Crises on Birthweight and Infant Mortality

This paper tests the hypothesis of whether crises in Indonesia have any effects on child mortality and birthweight. Children born during the crisis may be affected differently than those conceived during the crisis. Neonatal mortality is influenced by conditions during pregnancy or at birth, while post-neonatal mortality is influenced more by external factors during child rearing after birth[45]. Since conditions during pregnancy and after births are more likely to affect neonatal and post-neonatal mortality differently, the analysis is carried out separately for neonatal and post-neonatal mortality.

Birthweight and child mortality may be affected by various socioeconomic characteristics other than conditions affected by economic crises. In this paper, mother's education, urban/rural residence, geographic location (provinces and communities), and child's gender are used as control variables.

According to Schultz,[46] women's education has external benefits to society. Higher mother's education reduces child mortality, improves child nutrition and schooling, and decreases fertility and population growth. Mother's education benefits child health in several ways. First, education helps make learning of childcare process more efficient, especially when the process is complex (*technical efficiency*). Second,

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[45] A. Razzaque, N. Alam, L. Wai, and A. Foster. " Sustained Effects of the 1974-5 Famine on Infant and Child Mortality in a Rural Area of Bangladesh." Population Studies. March 1990

M. Rahman, B. Wojtyniak, M. M. Rahaman and K.M.S. Aziz, "Impact of environmental sanitation and crowding on infant mortality in rural Bangladesh." The Lancet (1985), pp. 28-32

[46] T.P. Schultz. "Investments in the schooling and health of women and men: Quantities and returns**."** Journal of Human Resources**.** Fall 1993; Vol. 28, Issue. 4; pg. 694, 41.

education helps mothers allocate household resources efficiently to improve child's health (*allocative efficiency*). For example, mother's education increases the willingness to seek medical care and improves nutrition and sanitation practices.[47] Third, mothers with higher education are more likely to earn more income, and therefore can use this additional income to consume more or better-quality child's health inputs (*income effects*). Fourth, higher education may help improve women's bargaining power in household resource allocation. According to Schultz (1993), women may channel more of their income to expenditures on children than their husbands do. Improving women's education, therefore, could result in higher women's bargaining power, which in turn yields an allocation of more of the household income to expenditures on children.

When studying the effects of mother's education on child mortality, one should consider a possibility that education and health services are substitutes. According to a review by Basu and Aaby (1998)[48] that refers to studies by Palloni (1985)[49] and Rosenzweig and Schultz (1982),[50] the dominant theoretical stance on the education-mortality association is that the influence of personal characteristics, such as maternal

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[47] Schultz, T. Paul. 1989. Benefits of Educating Women. Washington, D.C.: World Bank, Background Papers Series, Education and Employment Division, Population and Human Resources Department."The Benefits of Education for Women HRO Dissemination Notes." Human Resources Development and Operations Policy. Number 2, March 8, 1993.

Mellington and Cameron (1999) find that mother's primary and secondary schooling significantly decrease the probability of child death in both rural and urban areas in Indonesia. "Female Education and Child Mortality in Indonesia." Melbourne- Department of Economics in its series papers number 693. 1999

[48] Basu, A.M. and Peter Aaby. *The Methods and Uses of Antropological Demography*. 1998. Clarendon Press. Oxford.

[49] Palloni, A. Health Conditions in latin America and policies for mortality change', in J. Vallin and A. Lopez (eds.), *Health Policy, Social Policy and Mortality Prospects*. Liege: Ordina. 1985

education, on child welfare attenuates when good health services are widely available. However, many studies give contrary examples in which access to services appears to make little difference to education differentials (for more detailed discussions, see Cleland and Van Ginneken, 1989[51]). For example, Bicego and Boerma (1991) find a stronger effect of maternal education in urban areas, where health services are assumed to be widely available, than in rural areas.

Table 7 shows the distribution of the mean education in years and percentages of different education levels of mothers over time. Formal education in Indonesia had been successfully improved. From our sample, we observe a significant increase in the mean education of mothers from 5.0 to 8.2 years during the eleven-year period. For example, the proportion of mothers with no formal education decreased from 19.2 percent to 4.6 percent (76 percent decrease). The proportion of mothers with 1-5 years of schooling decreased from 33.9 percent to 14.8 percent (56 percent decrease), while the proportion of mothers with 9-11 years of schooling increased from 7.9 percent to 20.4 percent (158 percent increase).

Table 8 shows that higher mother's education is associated with lower mortality. The effect seems to be stronger for post-neonates than neonates. If mother's education contributes to child mortality as previously argued, we must control for mother's education when studying the effects of crises. Since the children born in the crises

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[50]Rosenzweig, M and T.P. Shultz. "Child mortality and fetility in Colombia: individual and community effects." *Health Policy and Education*, 2: 305. 1982.

[51] Cleland, J. and J. Van Ginneken. 1989. "Maternal education and child survival in developing countries: the search for pathways of influences." Social Science and Medicine. 27: 1357-60.

periods are from mothers of later cohorts for whom education is higher, not controlling for mother's education, will underestimate the impact of the crises.

The probability of neonatal and post-neonatal mortality is estimated using linear probability and logit regressions. Dependent variables are whether a child born alive had neonatal mortality, whether the child had postneonatal mortality given he/she survived neonatal mortality[52] and whether a child born alive had infant mortality (either neonatal mortality or post-neonatal mortality). Province or communities dummies are used to control for time-invariant community-specific unobserved factors that may influence child mortality (such as local desease patterns and public health infrastructure[53]). Since the focus of this paper is to study the effects of the financial and the drought/smoke crises on child birhtweight and mortality, controls thought to be correlated with the crises such as household income and time-varying public health provision are excluded.

In addition to socioeconomic controls, a time trend is included in our analysis to control for mortality trends that can be observed in Figure 5. According to the data, both neonatal and post-neonatal mortality increased during the crises period, but the level of these mortality rates are still not as high as that of the earlier control period (1988-1992). Since neonatal and post-neonatal mortality had generally declined during the pre-crisis periods from a relatively higher rate in 1988, omitting the time trend would result in an underestimated effect of the crises.

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[52] Therefore, those born alive that experienced neonatal mortality are dropped from our sample when studying post-neonatal mortality.

[53] Pertersen, W. and R. Petersen with the collaboration of an International Panel of Demographers. 1986. Dictionary of Demography: Terms, Concepts, and Institution. Greenwood Press. New York. Westport, Connecticut. London.

As discussed earlier, the sample used includes children who were at least one year old at the interview date. Both results from the linear probability and the logistic regression are reported. Since urban and rural areas may be affected by different crises in different ways, regressions are performed separately for urban and rural areas.

Due to the nature of the dependent variables, the linear probability model violates one of the Gauss-Markov assumptions. When the dependent variable is a binary variable, its variance, conditional on the explanatory variables, depends on the explanatory variables (unless the probability of success does not depend on any of the explanatory variables):

$$\mathrm{Var}\,(y|x) = p(x)\,[1-p(x)]$$

Where, p(x) is the probability of success, which depends on x.

As a result, there must be heteroskedasticy in a linear probability model. Although heteroskedasticity does not cause any bias in the OLS estimators of the coefficients, homoskedasticity is crucial for justifying the usual t and F statistics.[54] Regressions robust to heteroskedasticity are, therefore, included to correct the conditional variances.

In addition, since our period of analysis includes children born in 1988-2000, children in our sample can be from the same mother. The outcomes of children within a group of mother are likely to be correlated. In this paper, observations are clustered at the mother level to allow for this type of heteroskedasticity problem in the variances of the estimated coefficients.

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[54] For further discussion, refer to Jeffrey M. Wooldrige. "Introductory Econometrics: A Modern Approach" Copyright 2000. by Sourth-Western College Publishing. USA.

The effects of economic crises on birthweight can be tested nonparametrically by comparing cumulative distribution of birthweight of children born during non-crisis period and that of children born during the crisis period. Special attention, however, should be paid to proportion of children who have "low-birth-weight," which is birthweight that is less than 2.5 kilograms. Note that when testing the differences of the distributions, data are restricted to exclude outliers. The tests include only birthweights between 1.5[55] and 3.0 kilograms. Computing for the difference in the values of the cumulative distribution at each weight point of interest was performed. The comparison follows directly the formulation offered by Davidson and Duclos (2000).[56] We will be able to conclude that the crisis has statistically significant adverse effects on birthweight when the cumulative distribution value of birthweight of those born during the crisis period at all birthweights of interests (that are considered low birthweight) is unambiguously higher than that of those born during non-crisis period. In other words, we are testing whether the cumulative distribution of low birthweights of those born during the crisis period first stochastically dominates that of the cumulative distribution of low birthweights of those born during the non-crisis period.

When the test of first order-stochastic dominance fails, we test for relative riskiness or dispersion of birthweight (second-order stochastic dominance). This is to test whether those born during the crisis period as a group exhibit relatively higher risk of low birthweight than those born during non crisis period. In other words, we test whether the

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[55] There are few birthweights that are less than 1.5 kilograms. These observations are considered outliers.

[56] Russel Davidson and Jean-Yves Duclos. "Statistical Inference for Stochastic Dominance and for the Measurement of Poverty and Inequality." Econometrica v86 n6. 2000.

probability of those born during the crisis period that have birthweight at or below a specific weight is significantly higher than that of those born during the non-crisis period. If this is true for all low birthwieght values tested, we can conclude that the distribution of birthweight of those born during the crisis period second-order dominates that of those born during the crisis, and that those born during the crisis exhibit relatively higher risk of having low birthweight. Computing for the difference in the values of the cumulative distribution at each weight point of interest was performed using a software for Distributative Analysis/Analyse Distributative (DAD).[57] The techniques used in this software follow directly the formulation offered by Davidson and Duclos (2000).[58]

In addition to nonparametric estimation of the effects of the crises on birthweight, we include a simple parametric estimation using Ordinary Least Square (OLS) regressions, using birthweight as a continuous variable. Explanatory variables are the same set used in mortality regressions. Separate regressions are carried out for rural and urban samples.

Results

Figure 6a-6c present the comparisons of the cumulative distributions of birthweights in the crisis and the non-crisis periods. The results confirm that the reduction in birthweight was only among urban children born in the economic crisis. Table 9 shows the results of testing the difference of cumulative distributions using both

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[57] Copyright by Jean-Yves Duclos, Abdelkrim Araar, and Carl Fortin.

first and second order stochastic dominance tests. The points of testing are between 1.5 and 3.0 Kilograms. The results from these tests show that, within this range, none of the comparisons show any significant first and second-order stochastic dominance. However, it is hard to reject that the economic crisis had no impact on urban children as one can observe that there is some evidence suggesting a second order stochastic dominance in the range of 2.1-3.0 Kilograms among urban sample. In rural areas, the results are opposite to what we expect, but none of the comparisons shows first nor second-order dominance.

Table 10 presents the results from the multivariate OLS birthweight regressions. The dependent variable in these regressions is birthweight in kilograms. The estimations are carried out for urban and rural samples separately. Four specifications are presented for each sample. The first specification is our base regression. In this specification, explanatory variables include only the crisis dummies and the time trend. Mother's education dummies are added in the second specification. The last two specifications use province and community dummies to controls for location-specific unobserved factors respectively. Regardless of which geographic location is used, the regressions essentially yield fixed-effect estimators.

The results from these regressions suggest that all of the crises had no statistically significant adverse effects on birthweights. We find no effects even after controlling for other factors that may affect birthweight. We observe a downward trend in birthweight in urban areas and an upward trend in rural areas, but the time trend variable does not appear to have statistically significant effects on birthweight. Male children had

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[58] Davidson, R. and Jean-Yves Duclos. "Statistical Inference for Stochastic Dominance for the

significantly higher birthweight than female children in both urban and rural areas. In any case, one should keep in mind that these regression results are based on birthweights that were reported. The observations are subjects to potential selection problems as discussed earlier.

The results of the multivariate analysis of infant mortality are shown in Tables 11 to 14. Tables 11 and 12 present results from the LPM and the Logistic regressions using the urban sample. Tables 13 and 14 present results from the LPM and the Logit regressions using the rural sample.[59] In each set of regressions, four specifications are presented. The first specification is our base regression. In this specification, explanatory variables include only the crisis dummies and the time trend. Mother's education dummies are added in the second specification. The last two specifications use province and community dummies to controls for location-specific unobserved factors respectively. Regardless of which geographic location is used, the regressions essentially yield fixed-effect estimators. In the Logit regressions, the last columns in each group of regressions estimates conditional Logit model with fixed province/community effects. To be able to identify the conditional effects, observations used in these regressions are only from those provinces/communities that experienced at least one incidence of mortality. The provinces/communities without mortality incidence are therefore

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Measurement of Poverty and Inequality." Econometrica, V86. n6. 2000.

[59] None of those born during financial crisis 2 period (122 obs) experienced neonatal mortality. Since we are interested in the estimated coefficients of the crisis dummies (fiancial and drought/smoke crises), the "pregnancy in fiancial crisis2" dummy variable was dropped in the Logit regressions instead of dropping these observations. The dummy variable was also dropped in the corresponding LPM regressions.

excluded from the sample.[60] To be consistent with the Logit regressions in terms of observations used, the last two columns of the LPM regressions limit the samples to be the same as those in the Logit regressions. Also, when community dummies are included, crisis dummies pick up effects that vary within communities. Then, if there is no variation in the communities, observations belonged to these communities do not help identifying the crisis effects.

The estimated coefficients from the linear probability models are the estimated partial effects. For Logit regressions, the odds ratios[61] are reported. As discussed earlier, explanatory variables of interest in the neonatal mortality regressions are whether the child was conceived during the crises, while explanatory variables of interest for post-neonatal and infant mortality regressions are whether the child was born during the crises. The LPM and the Logit regressions give similar results.

In urban areas, the estimated coefficients of the time trend are statistically significant in neonatal, post-neonatal, and infant mortality regressions regardless of whether or not the regressions control for other fact.ors The estimates suggest that both neonatal and post-neonatal mortality rates in urban areas decreased by approximately 0.2 percentage points per years (by 2 deaths per 1,000 live births per year). Infant mortality declined by approximately 4 deaths per 1,000 live births per year. The estimated effects of the time trend are significantly larger in the community fixed-effects regressions.[62]

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[60] The data show that none of the infants born in urban areas of Lumpung and Bali provinces experienced post-neonatal mortality. Similarly, none of the infants born in rural Yogyakarta experienced post-neonatal mortality.

[61] The odd ratios, e^{β}, show impacts in terms of Prob (mortality | x) / Prob (survival | x).

[62] Recall that observations used in the LPM community fixed-effect regressions and the conditional Logit regressions includes only communities with at least one incidence of mortality.

In urban areas, children born during the peak of the financial crisis (crisis1) exhibited higher odds of neonatal, post-neonatal, and infant mortality. The effects of the crisis were more precisely estimated in the Logit regressions. Controlling for mother's education and province of residence does not change our point estimates by much. when controlling for the community fixed effects, these point estimates, however, decrease in the Logit regression and increased in the LPM regressions .

Results from the linear probability regressions show that children conceived during the peak of the financial crisis (crisis1) had approximately 1.7 percentage points higher probability of neonatal mortality than those conceived during the non-crisis periods. Similarly, those born during crisis1 exhibited approximately 1.6-1.7 percentage points higher probability of post-neonatal mortality than those born during the non-crisis periods. However, the impacts of the crisis are not statistically significant when community fixed-effects are controlled for.

Urban children conceived during the drought/smoke haze crises exhibited higher neonatal mortality than those conceived during the non-crisis periods. The adverse effects are estimated to be about 2.1 percentage points in the LPM regressions. The adverse effects were statistically significant at 10 percent level even after controlling for the community fixed effects using Logit regression. The 1997 drought/smoke crisis did not have any statistically significant effects on post-neonatal mortality in urban areas.

The results from the infant mortality regressions suggest that only economic crisis had statistically significant impact on infant mortality in urban areas. Similar to results from the neonatal mortality regressions, the effects of the crisis were not statistically significant in the fixed-effect estimation for infant mortality.

In urban areas, male children had higher neonatal, post-neonatal, and infant mortality rates. However, the estimates are not statistically significant. In addition, we found that mother's education did not play a significant role in reducing neonatal mortality.[63] On the other hand, the effects were felt in post-neonatal mortality. Although in some specifications the points estimates are not statistically precise, we do observe that children born to mothers with at least 9 years of experienced lower probability of post-neonatal mortality by approximately 2.8-3.5 percentage points. When controlling for the community fixed effects, the magnitude of the effects of twelve or more years of education more than doubled the effects in province fixed-effect regressions.[64]

In rural areas, both neonatal and post-neonatal mortality rates decreased over time. The estimated coefficients of the time trend, however, suggest that rural areas experienced a decline in these mortality rates at a slightly slower rate than the urban areas. The community fixed-effect estimations suggest that the reduction in mortality risks was at a faster rate in communities that experienced at least one mortality incidence, controlling for community fixed effects.

In rural areas, the financial crisis 1 period had adverse effects on neonatal and post-neonatal mortality. However, the estimated effects are statistically significant in only neonatal mortality regressions. In these regressions, the estimated effects are statistically significant at 10 percent level even after controlling for community fixed-effects. Rural infant mortality was affected by financial crisis 2 instead of financial crisis

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[63] Except at 12 or more years of education in the Logit regression.

[64] Recall that observations used in the LPM community fixed-effect regressions and the conditional Logit regressions include only communities with at least one incidence of mortality.

1. The increase in infant mortality rate is estimated to be approximately 3.2 percentage points. In those communities where we applied community fixed-effect estimation, the effects are higher (4.9 percentage points) and statistically significant at 5 % level.

In rural areas, the drought/smoke haze crisis exhibited statistically significant adverse effects on post-neonatal mortality, but not on neonatal mortality. The effects of the crisis on post-neonatal mortality are stronger than the effects of the financial crisis. The LPM estimations show that infants born during the drought/smoke haze crisis had approximately 2.7-3.1 percentage points higher probability of post-neonatal mortality than those born during the non-crisis periods. The estimated effects are higher (6.3 percentage points in the LPM) and still statistically significant when controlling for community fixed-effects in both the LPM and the Logit regressions.

Mother's education played a greater role in reducing neonatal mortality in rural areas than in urban areas. Our estimates from the LPM regressions suggest that mother's primary education is associated with approximately 1.6-1.8 percentage points lower neonatal mortality rate. The effects of education was much stronger (4 percentage points) using community fixed-effect estimation for those communities that experienced neonatal mortality. If a higher level of mother's education increases effectiveness of prenatal care, these results imply that increasing mother's education in rural areas will help reduce neonatal mortality. An explanation of our finding that higher mother's education did not reduce neonatal mortality in urban areas is that urban mothers' ability and effectiveness in providing prenatal care could be substituted by better services and less expensive healthcare that were more readily available in urban areas.

Similar to urban areas, education of rural mothers reduced post-neonatal mortality. Moreover, our estimates suggest that in rural areas mother's education started to have statistically significant effects on post-neonatal mortality at secondary education level, much earlier than in urban areas.

In rural areas, we found that male infants had a higher chance of both neonatal and post-neonatal mortality than female infants. Recall that in urban areas, a child's gender had no statistically significant effects on neonatal mortality. If male infants are biologically more prone to neonatal mortality than female infants, the result from our regressions suggest that better healthcare and more exposure to prenatal services (such as those available in urban areas) can help overcome higher risks of neonatal mortality among male infants.

How did mothers with different education levels cope with the crises ?

An important question we might ask when studying the effects of the crises on infant mortality is how households with different socio-economic backgrounds coped with adverse short-term shocks. For instance, we ask whether poor households responded differently to the financial crisis than rich households did. Nevertheless, since the financial and the drought/smoke haze crises had direct impacts on household's income, distinguishing households by income level may result in biased estimates of the effects of the crises as the crisis itself determines which income group a household belonged to. To avoid this selection problem, we want to use determinant of household' income that are not affected by the crises (in the short-run). In this paper, we experiment with different mother's education levels as such determinants.[65]

Tables 15-17 show LPM mortality regression results of children born to mothers with 0-8 years and 9+ years of education. Table 15 shows the results from infant mortality regressions of infants born to mothers with different levels of education. We observe from the point estimates that in urban areas, infant mortality trends are different between the two groups, even after controllong for community fixed effects. Those born to mothers with lower education levels experienced a decline in infant mortality at a rate that is twice as much as the rate for those born to mothers with higher levels (0.6 vs 0.3 percentage points per year). However, a test of the differential trend effects between these two groups in urban areas indicates that the differential effects are not statistically significant at the 10 percent level.[66] In contrast, the difference in the time trend effects between low and high education groups is small in rural areas. The estimated effect of the time trends for both groups is approximately 0.4 percentage points per year.

In urban areas, the point estimates of the financial crisis effects on infant mortality suggest that the adverse effects are stronger for infants born to mothers with lower education. However, a test of the diffential effects cannot reject that the financial crisis effects on the two groups are statistically similar.[67] In rural areas, the effects of the financial crisis on infant mortality were slightly higher for the lower education group. A

[65] Alternatively, one can ask directly how mothers of different education levels cope with the crises.

[66] The t-statistic for the differential effects is 1.50 (p = 0.13).

[67] The t-statisitics for the differential effects are 1.43 (p = 0.15) for financial crisis 1 and 0.03 (p =0.98) for financial crisis 2.

test of the differential effects indicates no statistically difference between the effects on the two groups.[68]

In rural areas, the drought/smoke crisis had statistically significant adverse effects on infant mortality only for infants born to low education group. The magnitude of the drought/smoke effects are also much larger for those belonged to mothers with lower education than for those belonging to mothers with higher education group (6.0 versus 1.3 percentage points). The differential effects are larger when controlling for community fixed effects. However, the effects of the droughts/smoke crisis are not statistically different between these two education groups.[69]

Table 16 shows results from neonatal mortality regressions of infants born to mothers with low and high education. We found that in urban areas the mortality trends of the low and the high education groups are similar. Mothers with higher education levels experience a decline in neonatal mortality of their infants at a rate of approximately 0.28 percentage points per year, slightly higher than that of mothers with lower than six years of education. We observe a larger differential between the two groups after controlling for province and community fixed effects. However, the differential effects of the time trends are still statistically insignificant.[70] In rural areas, neonatal mortality had not been statistically reduced over time regardless of the education of the mother. When community fixed effects are controlled for, the differential effects of the time trend between the two groups is larger, but the point estimates are still statistically

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[68] The t-statisitics for the differential effects are 0.36 (p =0.72) for financial crisis 1 and 0.08 (p =0.94) for financial crisis 2.

[69] The t-statisitic for the differential effects is 1.50 (p = 0.13).

[70] The t-statistic for the differential effects is 0.61 (p =0.54).

insignificant. A test of the differential effects cannot reject that the neonatal mortality trends among the two education groups are similar.[71] In addition, we observe higher and more significant time-trend effects in urban areas than in rural areas for both education groups.

In urban areas, financial crisis 1 had statistically significant adverse effects on only infants with high education group. Our point estimates suggest that infants born to mothers with lower education were more adversely affected by the drought/smoke crisis in their risks of neonatal mortality than infants born to mothers with low education. However, the differential effects are not statistically significant.[72] In rural areas, infants born to mothers with higher education during the second phrase of the financial crisis and the drought/smoke crisis, however, had a lower neonatal mortality rate than those born during the non-crisis periods. A plausible explanation is that even though the neonatal mortality rate for this groups of infants increased during these crisis periods, the increase in mortality risk is not so high to surpass the rates of the pre-crisis period. This explanation is support by the fact that the negative time trends were not statistically significant for this group of infants. When the community fixed effects are controlled for (for those with 9+ years of education), the crisis dummies still have negative signs while the estimated effect of the time trend is negative and significant. This result is puzzling. However, our test of the differential effects of the crises suggest that there is no statistically significant differential effects between the two education groups.[73]

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[71] The t-statistic for the differential effects is 0.42 ($p = 0.67$).

[72] The t-statistics for the differential effects are 0.62 ($p = 0.54$).

[73] The t-statistic for the differential effects is 0.04 ($p = 0.97$), 0.83 ($p = 0.41$), and 1.00($p = 0.32$) for financial crisis 1, financial crisis 2, and drought/smoke crisis.

The neonatal mortality regressions give another interesting result when we look at infants of different genders. In urban areas, we observe small and statistically insignificant gender discrimination in the mortality rate regardless of levels of mother's education. In rural areas, among those born to mother with lower education, male infants had 1.4 percentage points higher neonatal mortality risks than female infants. The the differential effect is higher after controlling for community fixed effect. In contrast, male infants had less neonatal mortality risk that female infants among those born to mothers with higher education. Our test indicates that the gender differential in neonatal mortality risk among infants born to mothers with low education is significantly different form that among infants born to mothers with high education.[74]

Table 17 shows results from post-neonatal mortality regressions of infants born to mothers with low and high education. We found that in urban areas, the mortality trends are significantly different among the low and the high education groups at the 10% level.[75] In urban areas, infants born to mothers with less than nine years of education exhibited a much larger declining trend than those born to mothers with higher education. The estimated effects of the time trend are estimated to be approximately 0.4-1.2 percentage points per year for lower education group. In contrast to the trend difference in urban areas, the difference between the effects of the time trends between these two education groups are less pronounced in rural areas. Rural infants born to mothers with lower education experienced a faster decline in post-neonatal mortality risk. The

[74] The t-statistic for the differential effects is 1.79 (p = 0.07).

[75] The t-statistic for the differential trend effects is 1.67 (p = 0.09).

estimated effects of the time trend are lower in rural areas than in urban areas (-0.24 versus -0.33 percentage points per year) for lower education group. On the contrary, the time trend effects are higher in rural areas than in urban areas for higher education group (-0.13 versus -0.04 percentage points per year).

In urban areas, the adverse effects of the financial and the drought/smoke crises on post-neonatal mortality seemed larger for those born to mothers with lower level of education than for those born to mother with high education. However, none of the estimated coefficients of the crisis dummies are statistically significant in either high or low education groups. The differential effects of the crises between the two samples, in turn, are not statistically significant.

In rural areas, unlike in urban areas, the adverse effects of the financial crisis appear to be similar for both education groups. As expected, the drought/smoke haze crisis adversely affected children born to both groups of mothers. However, the drought/smoke crisis had much worse effects on post-neonatal mortality for infants in low education groups than that of infants in high education group. Our point estimates suggest that infants belonging to mothers with less than 9 years of education tended to suffer from the drought/smoke haze crisis much more than those belonging to mothers with at least 9 years of education (4.3 versus 0.0 percentage points). A test of the differential effects indicates statistically significant differential effects between the two education groups.[76]

In urban areas, our estimates show a discrimination in post-neonatal mortality between male and female infants born to mothers with at least nine years of education.

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[76] The t-statistic for the differential effects is 1.84 (p = 0.07).

The discriminating effect is large when controlling for community fixed effects (4.7 percentage points). A test of the differences in gender discrimination suggest no differential effects between the two education groups.[77] In rural areas, gender discrimination in post-neonatal mortality risks is statistically significant for only the low education group. Male infants born to mothers with lower education experienced about 1.4 percentage points higher post-neonatal mortality rate than female infants. The difference between the two education groups are, however, not statistically significant.[78]

In sum, our findings indicate that infants born to mothers with different levels of education exhibited no significantly different trends in mortality risks over time in both urban and rural areas. We also find that even though the economic crises had adverse effects on infant mortality, infants born to mothers with different education levels did not experience statistically different adverse financial crisis effects in either urban or rural areas. There were statistically significant differencial effects of the drought/smoke crisis on post-neonatal mortality between the low and the high education groups in rural areas. Those born to mothers with lower education were more adversely affected by the drought/smoke crisis than those born to mothers with higher education.

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[77] The t-statistic for the differential effects is 0.93 (p = 0.35).

[78] The t-statistic for the differential effects is 0.71 (p = 0.48).

Hazard Models

The findings obtained in the previous section are based on sample observations that include only mortality of children who were exposed to at least one year of life. Many children born in 1999 and all children born in 2000 are excluded by this criterion.

Since we are interested in estimating the probability of dying before one month and within 1-11 months, we can use duration models to directly extract the hazard rates (the probability of dying in the next period given survival up to the current period) without having to exclude those who were not yet one year old at the time of the interview. Using the duration model is also a better way to capture the precise time of death assuming reported dates are accurate. By using duration models, each month's survival status information of each child is used to estimate the mortality hazard. The hazard rate of each month can then be obtained.

Figure 7a-7c show nonparametric estimates of discrete monthly child mortality hazard rates from Nelson-Aalen cumulative hazard function. In urban areas, those born during the financial crisis exhibited higher hazard rates that those born during the non-crisis periods. Similar to results from the previous regressions, in rural areas, the drought crisis had more overall effects on child mortality than the financial crisis. The results from testing the equality of the survivor functions of children born in these periods, however, show that none of crises had statistically significant adverse effects on infant mortality. These insignificant crisis effects are inconsistent with our findings from the OLS and logit regressions. Nevertheless, one should keep in mind that the hazard rates used in this section to calculate the crisis effects are estimated without any controls.

Parametric Estimation of Child Mortality Using Hazard Models

Let T ≥ 0 denotes the length of time the child lived in months. T has some distribution over the population. Consider time invariant covariates.

Let F(t; x) = conditional cdf of T where x = covariates:

F(t; x) = P(T ≤ t; x), t ≥ 0

The survivor function is defined as

$$S(t;x) = 1 - F(t;x) = P(T > t;x)$$

Then, the probability of leaving the initial state in the time interval (t, t+h) is

$P(t \le T < t+h \mid T \ge t;x)$ for h > 0

Define the hazard function as

$$\lambda(t;x) = P(t \le T < t+h \mid T \ge t;x) = P(t \le T < t+h;x) / P(T \ge t;x) = \frac{F(t+h,x) - F(t;x)}{1 - F(t;x)}$$

If the cdf is differentiable, then the hazard function is

$$\lambda(t;x) = \frac{F(t+h,x) - F(t;x)}{1 - F(t;x)} = \frac{f(t;x)}{1 - F(t;x)} = \frac{f(t;x)}{S(t;x)}$$

Then all probabilities can be computed using this hazard function. For example, from time a to time b, a < b is

$$P(a \le T < b \mid T \ge a;x) = 1 - \exp\left[-\int_a^b \lambda(s;x)ds\right]$$

Since we did not observe the end of the survival period of every child in the sample, the survival data obtained from the survey are considered "flow" data, which are subject to time censoring. In this case, the data are right-censored at the interview date. For example if the child born on January 1, 2000 was still alive on the interview date of June 30, 2000. We only know that the child's survival time was at least six months. We never observed real survival time of this child. The model can be adjusted to include this time censoring by defining censored flow data in the following way.

Observed duration t*

Define t_i* as the length of time in the initial state that has a continuous conditional density

$$f(t_i \mid x_i;\theta), t \geq 0$$

where θ is the vector of unknown parameters.

The observed length of time, t, in the initial state is

$$t_i = \min(t^*{}_i, c_i)$$

where c_i is censoring time for individual i. In this case, the censoring time is age of the child the interview date.

Table 18 and Table 19 show results of parametric estimations of hazard rates of urban and rural children respectively. Hazard regressions use Weibull distribution to allow for both positive and negative monthly hazard rates. The standard error and the z statistics of the hazard ratios are robust to heteroskedasticity of the variance-covariance matrix at the mother level. In addition, we also present results of these hazard regressions assuming heterogeneity in our observation. In our regressions, the observations are assumed to have inverse-Gaussian heterogeneity (frailty).

The estimated results are similar to those obtained from previous LPM and logit regressions. Those born during the financial and the drought/smoke haze crises appear to have been negatively affected by the crises. The estimated effect of the financial crisis is statistically significant at 10 percent level while the estimated effect of the drought crisis is significant at 5 percent level. Those with mother of at least 12 years of education have lowers estimated hazard rates than those with no education. The estimated coefficient of the male dummy suggests boys had a higher mortality risk than girls.

In rural areas, those born during the drought/smoke haze crisis appear to have been adversely affected. Similar to results found in the LPM and the Logit regressions, the financial crisis did not have statistically significant effects on rural children. Education was also more important in rural areas than in urban areas. In rural areas, having some education, regardless of the education level, helped reduced the risk of mortality. Similar to the results from urban areas, male children exhibited higher risk of mortality.

The results from the hazard regression allowing for heterogeneity adjustment confirms that the observations used are heterogeneous.[79] However, the degree of the statistically significance of the estimated coefficients are similar to those in the regressions that assume homogeneity of the observations.

Figure 8a-8c show estimated child mortality hazard rates using parametric hazard models. Results from hazard models confirm that the financial crisis and the drought

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[79] We reject that the variance is zero with p-value = 0.000 for both urban and rural observations.

crisis had adverse effects on neonatal and post-neonatal mortality. When comparing rural and urban samples of the same period, rural children exhibited higher probabilities of both neonatal mortality and post-neonatal mortality than urban children. The financial crisis increased the odds of both neonatal and post-neonatal mortality for urban children more than for rural children. As expected, in rural areas, the crisis-noncrisis differential of the mortality risk at specific age (months) is relatively smaller for financial crisis than that of the drought crisis.

Discussions

Several child's health outcomes are examined in this paper to assess whether the Indonesian financial and drought/smoke crises negatively impacted young children in Indonesia. Birthweights, neonatal, post-neonatal, and infant mortality are examined in this paper. The results from both the birthweight cumulative distribution comparison and the OLS estimation similarly suggest that none of the crises had negative impacts on birthweight in both urban and rural areas. Nevertheless, we realize that this evidence is drawn from reported birthweights only. An investigation of the probability of reporting birthweight in various delivery locations in Indonesia suggest that the birthweight distribution drawn from reported birthweights may be biased due to unreported birthweights of infants that were born at home or in offices of traditional midwives. Further, biased results could come from selection problems when some mothers switched from hospital to home delivery in the time of crisis. Unfortunately, we do not have any plausible instrument to correct for this selection problem in reported birthweight. It is difficult to find a factor that affects delivery location (and whether birthweight was reported), but does not affect birthweight itself.

Unlike birthweight, our findings indicate that the financial and the drought/smoke crises had significant adverse effects on infant mortality. The overall effects were different in urban and rural areas. Although the fianancial crisis had adverse effects on neonatal mortality in both urban and rural areas, the effecs on post-neonatal mortality were felt by only urban infants. Contrary to the effects of the financial crisis, the effects of the drought/smoke crisis on post-neonatal mortality were felt by only rural infants.[80]

Results from multivariate regressions, nonparametric, and parametric hazard estimations show similar effects on infant mortality. In urban areas, infant mortality were affected by only the fiancial crisis. In rural areas, infant mortality was affected by both the financial and the drought/smoke crises, but the drought/smoke crisis appeared to have worse effects than the financial crisis.

These differentials results are consistent to our expectations. Since a large increase in the food prices and an adverse income shock during the financial crisis resulted in a sharp reduction of real incomes for most of the Indonesian population, who are net food purchasers, we expect to find some adverse effects on child health outcomes among both urban and rural population. Since the smoke haze that resulted from erupted fires during late 1997 and early 1998 affected only parts of urban areas in our sample, we do not expect to find strong effects of the smoke haze on overall urban population. Notice that when community fixed effects are controlled for in our analysis, the effects of the

[80] We observed that neonatal mortality in urban areas ws affected by the drought/smoke crisis, but the level of statistical significance is at 10%. After controlling for community fixed effects, the crisis appears to have no effects on neonatal mortality.

drought/smoke crisis on neonatal mortality in urban areas lost statistical significance. However, instead of combining the drought and the smoke/haze crisis, one could distinguish these two crises by mapping exact locations and levels of the fire smoke haze to each community in our sample in different periods. This task is plausible, but it would involve elaborative work in data collection.

Since the focus of this paper is to examine whether economic crises have any effects on child mortality, controls correlated with crises such as household income and public health provisions are excluded. Including these correlated variables in the regression, however, may answer different questions.

In addition, regressions in this paper do not control for fertility decisions. Since economic crises may affect fertility decisions, including fertility decisions in the regressions might lead to endogeneity problems. Demographic theory and empirical evidence from different countries suggest that mortality change should call forth some fertility response (Preston, 1975). Major examples of studies that suggest mortality-fertility link are given by Schultz (1969) and Federicksen (1966). According to Schultz, parents try to compensate for the average incidence of death by seeking the number of births that will give them the desired number of surviving children. Federicksen (1966) infers from the regional birth rates and the rates of population growth in Ceylon, Mauritius, and British Guiana that the improvement in health conditions that were responsible for the death rates led to a subsequent reduction in birth rates. On the other hand, some studies such as those by Adelman (1963) and Coale and Hoover (1958) suggest no significant link between mortality and fertility. For instance, Coale and Hoover (1958) in their classic text, *Population Growth and Economic Development in*

Low Income Countries, found the absence of major fertility decline in several developing countries that had experienced a prolonged mortality decline. The direction of the change in fertility decision during the Indonesian crises is, of course, an empirical one that needs to be further studied.

Since economic crises may affect people with different socioeconomic backgrounds differently, the study of the effects of economic crises can be extended to focus on these differential effects. Our preliminary findings suggest that even though some of the crises had adverse effects on infant mortality, infants born to mothers with different education levels did not experienced different adverse crisis effects.

As for infant mortality, we found that in urban areas, the financial crisis seems to have worse effects on infant mortality for those born to mothers with lower education. In rural areas, the effects of the financial crisis on infant mortality were slightly higher for the lower education group. However, when community fixed effects are accounted for, the financial crisis effects is higher (and statistically significant) for the higher education group. In rural areas, even though the drought/smoke crisis had adverse effects on infant mortality for both education groups, the magnitude of the effects are larger for those belonged to mothers with lower education. We found the differential effects between the two groups to be larger when controlling for community fixed effects.

Even though the true underlying causes of the differentials effects of the crises are not completely explored in this paper, we view our results from the estimations of the differential effects between different mother's education levels as some evidence that suggests differential effects of the crises on children with different socio-economic backgrounds. Since male and female children exhibited different mortality rates, studying the differential effects of the crises between male and female children may be of interest. This could be carried out by adding interaction terms of the crisis periods and a sex dummy or by estimating regressions separately for each group of children.

Conclusion

This paper examines the impacts of the recent Asian financial crisis and the 1997/98 drought and smoke haze crises on infant mortality and birthweight in Indonesia. The paper uses data from three waves of the Indonesian Family Life Survey: IFLS1 (1993), IFLS2 (1997), and IFLS3 (2000), utilizing rich data on socio-economic backgrounds as well as detailed information on children's birthdates, birthweights, mortality status at the time of interview, and ages at death if they died.

The methodology used in this paper is to compare health conditions of newborns of different birth cohorts. Specifically, this paper examines whether those conceived/born during the crisis periods exhibited higher risk of neonatal mortality and post-neonatal mortality and whether their birthweights were lower than birthweights of those born during the non-crisis periods.

The estimations of both neonatal and post-neonatal mortality risks are carried out using multivariate regressions with socio-economic control variables such as mother's education, place of residence (province/community), and gender of the child. In addition, mortality risks are estimated using hazard models to capture the mortality risks at different age (in months). The paper uses both nonparametric and parametric hazard models to estimates the hazard rates. The effects of the crises on birthweights are analyzed using multivariate regressions and comparisons of birthweight cumulative distributions. In both mortality and birthweight analyses, urban and rural samples are analyzed separately since the economic crises could have affected mortality and birthweight differently in rural and urban areas.

Estimated results on mortality outcomes show that the financial crisis had adverse impacts on neonatal mortality in both urban and rual areas. Urban infant conceived during the peak of the financial crisis exhibited approximately 1.7 percentage points higher neonatal mortality risk (17 per thousand live births more) than those conceived during non-crisis periods. The increase in neonatal mortality risk was approximately 2.2 percent for rural infants. The adverse effects of the financial crisis on post-neonatal mortality risks were larger and more statistically significant for urban infants than for rural infants. Overall, the financial crisis increased infant mortality risks by about 3.2 percent in both urban and rural areas.

The drought/smoke crisis adversely affected post-neonatal mortality risks in rural areas. The increase in the post-neonatal mortality risk is about 3.1 percent. When community fixed effects are controlled for, the drought/smoke crisis appears to have had much larger effects. Overall, the drought/smoke crisis had no significant advese effects on infant mortality in urban areas, while the effects in rural areas were large. Our estimates show that rural infants born during the drought/smoke crisis experienced approximately 4.4 percent increase in their infant mortality risks (44 per 1,000 live births). The magnitude of the effects almost doubled after controlling for community fixed effects.

Our findings on differential effects on infants born to mothers with different levels of education indicates that infants born to mothers with different levels of education exhibited no significantly different trends in mortality risks over time in both urban and rural areas. We also find that even though the economic crises had adverse effects on infant mortality, infants born to mothers with different education levels did not experience statistically different adverse financial crisis effects in either urban or rural areas. There were statistically significant differencial effects of the drought/smoke crisis on post-neonatal mortality between the low and the high education groups in rural areas. Those born to mothers with lower education were more adversely affected by the drought/smoke crisis than those born to mothers with higher education.

Results from the cumulative distribution comparisons of birthweights suggest that the financial crisis also had adverse impacts on birthweight in urban areas. However, under multivariate analyses, the adverse effect seems to disappear. None of the crises affected birthweights in rural areas. The lack of an evidence on the adverse effects maybe due to a selection problem in reported birthweights. The data show that from 1988 to 2000, 56 percent of women gave birth at home or at a family member's house. For those who were born at home, only 52 percent have reported birthweights, whereas 99 percent of those born in hospitals have reported birthweights.

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Figure 1: Timing of IFLS and the Rp/USD Exchange Rate

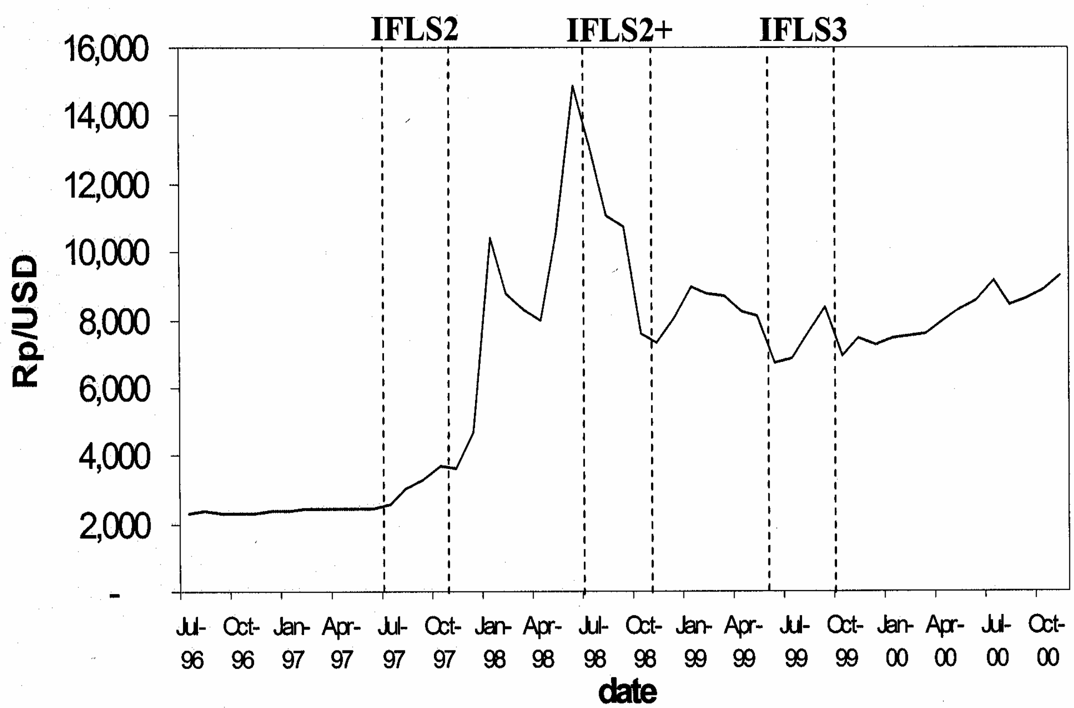


Figure 2: Food Price Index (January 1997=100)

Source: Kaiser K., T.A.Choesni, P. Gerter, L. Levine, and J. Molyneaux

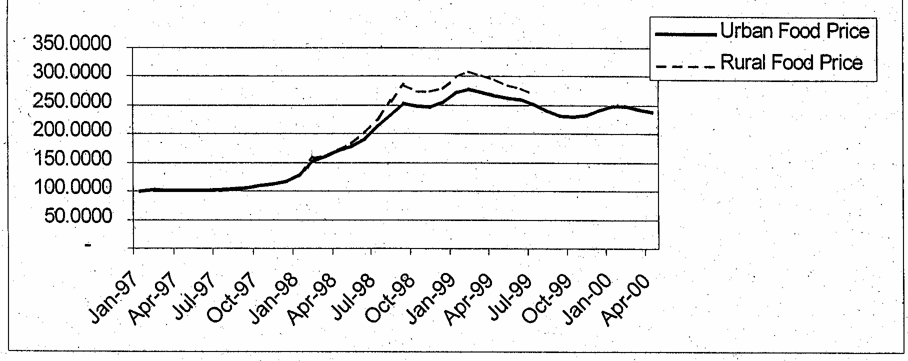
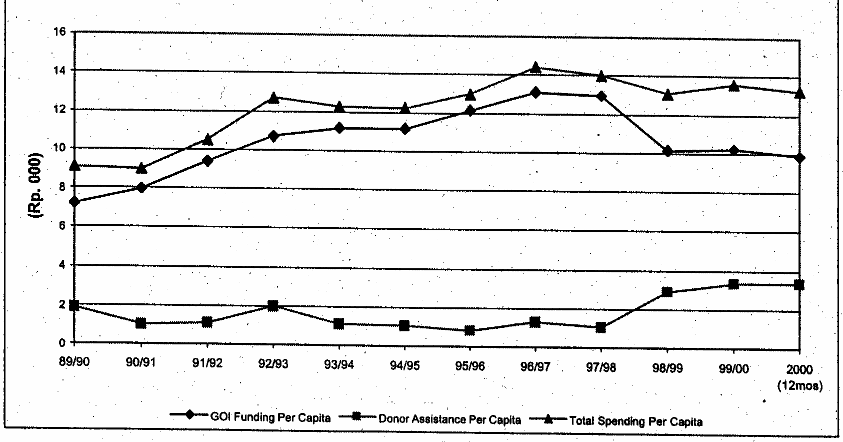


Figure 3

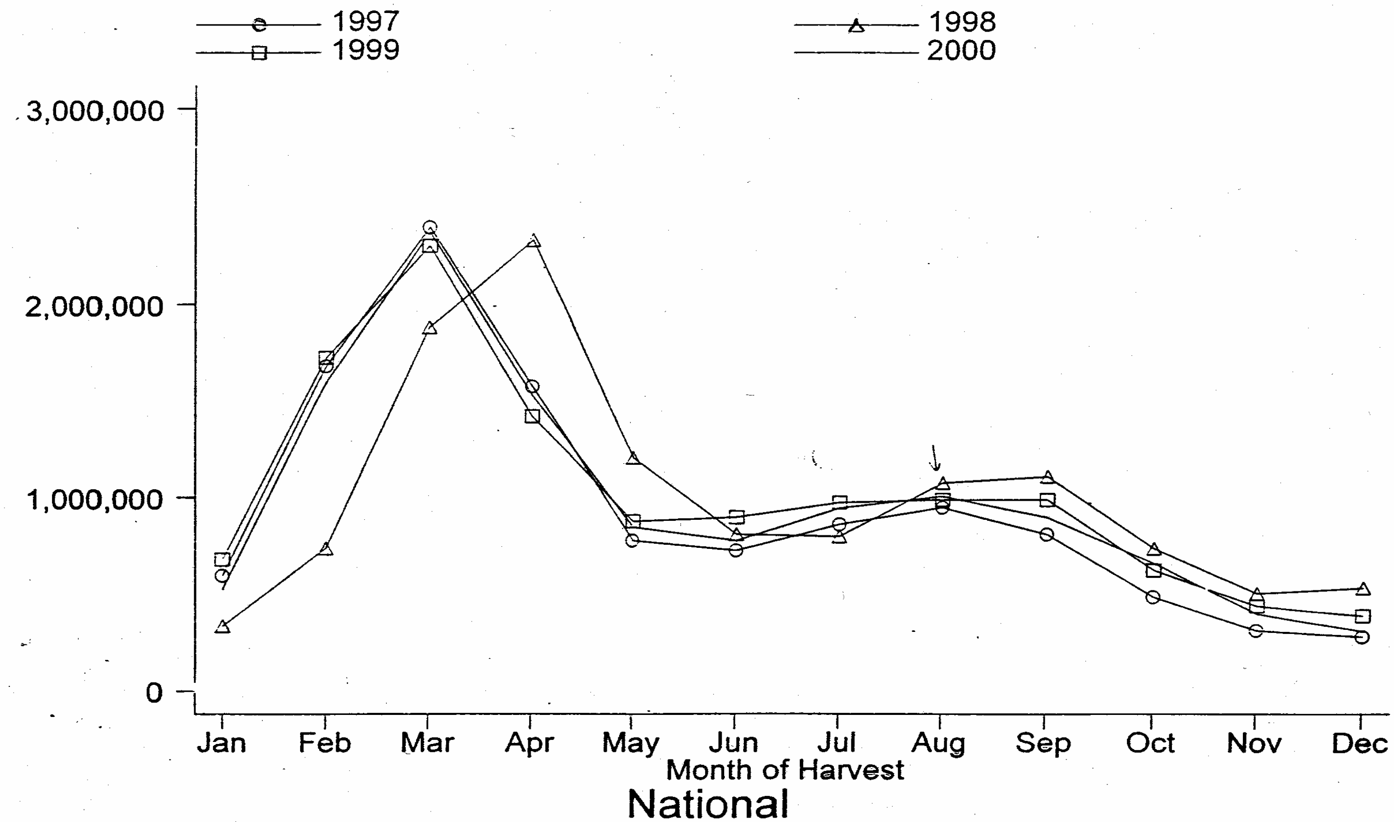
Government Health Expenditures, Overall and by Major Component, Per Capita
Constant Prices, 1993=100



Source: Figure 1. Lieberman, S., M. Juwono, and P. Marzoeki. "Government health expenditures in Indonesia through December 2000: An Update." World Bank East Asia and the Pacific Region Watching Brief.

Figure 4

Rice Area Harvested (Hectares), 1997-2000, by Month



Source: Courtesy of Jack Molyneaux

Table 1: Neonatal and Post-neonatal Mortality Rates Comparisons (per 1000 live births)
(Data from 1997 Demographic Household Survey of Indonesia and Indonesian Family Life Surveys)

1a. DHS 1997 Publication

	1992-1997		1987-1992		1982-1987	
Years preceding	0-4		5-9		10-14	
	Neonatal	Post-neonatal	Neonatal	Post-neonatal	Neonatal	Post-neonatal
DHS1997	21.8	23.9	28.2	30.3	28.4	37.0

1b. DHS Data (Replicated Numbers)

	1992-1997		1987-1992		1982-1987	
Years preceding 1997	0-4		5-9		10-14	
	Neonatal	Post-neonatal	Neonatal	Post-neonatal	Neonatal	Post-neonatal
Full sample (unweighted)	23.9	23.9	28.9	32.2	29	36.8
Full sample (weighted)	22.4	24.1	28.1	28.2	28.3	35.4
IFLS provinces (unweighted)	23.5	23.2	28.7	30.6	29.3	34.3
IFLS provinces (weighted)	21.9	23.6	27.5	27.1	28.0	34.7

1c. IFLS Data

	1992-1997		1987-1992		1982-1987		1995-2000	
Years preceding 1997	0-4		5-9		10-14		***	
	Neonatal	Post-neonatal	Neonatal	Post-neonatal	Neonatal	Post-neonatal	Neonatal	Post-neonatal
IFLS1	---	---	25.3	28.2	29.7	36.4	---	---
IFLS1 and IFLS2	15.5	17.9	28.0	31.1	30.6	38.5	---	---
IFLS2	16.1	20.1	---	---	---	---	---	---
IFLS2 and IFLS3	---	---	---	---	---	---	16.3	20.1
IFLS3	---	---	---	---	---	---	20.0	23.9

Figure 5: Mortality Rates (per 1,000 live births)

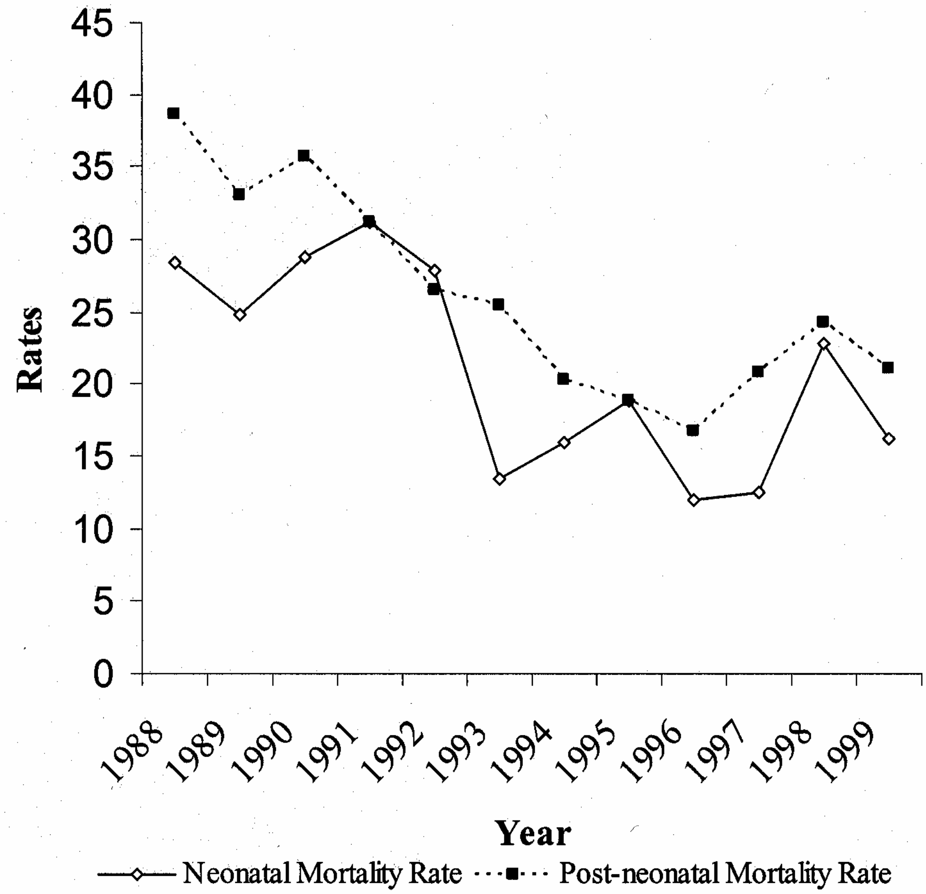


Table 2: Proportion of births that have reported birthweight by birth location (%)

Birth Location	Year Born 1988-1992	1993-1997	1998-2000	Average
Own house/family members house	38.4	55.7	64.6	51.6
Office/house of traditional midwife	41.7	45.8	54.6	46.2
Clinic of Physician/ Clinic or office of midwife	97.8	97.3	99.4	98.1
Community health center/village delivery post	98.1	98.6	97.0	98.0
Public/private/delivery hospital	99.7	99.1	99.5	99.4
Others	47.8	71.4	87.5	55.1

Source: IFLS1, IFLS2, and IFLS3

Table 3: Birth Location (%)

Birth Location	Year Born 1988-1992	1993-1997	1998-2000	Average
Own house/family members house	63.2	55.3	49.3	56.2
Office/house of traditional midwife	1.3	2.0	1.0	1.5
Clinic of Physician/ Clinic or office of midwife	17.1	23.0	28.6	22.6
Community health center/village delivery post	3.9	4.0	4.4	4.1
Public/private/delivery hospital	11.9	15.3	16.5	14.5
Others	2.5	0.4	0.4	1.1
Number of obs. with reported birth location	2682	3535	2275	8492

Table 4. Birth Location Comparison (%)

Birth Location	Year Born: Noncrisis 1995	Noncrisis 1996	Noncrisis 1997	Financial Crisis	Drought/smoke Crisis	Non-crisis 1999	Non-crisis 2000
Own house/family members house	55.4	52.5	48.4	50.6	50.4	48.8	47.7
Office/house of traditional midwife	1.8	2.2	1.2	1.0	2.2	1.6	0.6
Clinic of Physician/ Clinic or office of midwife	22.8	24.0	27.6	27.5	27.1	28.4	30.2
Community health center/village delivery post	3.6	4.0	4.0	3.9	3.2	5.1	4.5
Public/private/delivery hospital	16.1	17.2	17.2	16.8	16.9	15.6	16.6
Others	0.3	0.2	1.6	0.2	0.2	0.6	0.4
Number of Obs.	781	834	250	1033	498	514	728

Source: IFLS1, IFLS2, and IFLS3

Table 5: Mortality Rate (per 1000 live births)

	Neonatal mortality	Post-neonatal mortality	Infant mortality	Obs
Born during financial crisis	**23.1**	**23.1**	**46.2**	**1040**
rural	25.4	29.0	54.4	552
urban	20.5	16.4	36.9	488
Born during 97 drought period (rural)	**29.0**	**47.1**	**76.1**	**276**
Born during non-crisis period				
1988-1992	**28.2**	**33.0**	**61.2**	**4117**
rural	30.6	38.7	69.3	2324
urban	25.1	25.7	50.8	1793
1993-1997, 1999	**13.3**	**19.2**	**32.5**	**3542**
rural	15.7	25.8	41.5	1976
urban	10.2	10.9	21.1	1566

financial crisis = January 1998 to June 1999

drought97 = May 1997 to December 1997

non-crisis = January 1988 to April 1997 and July 1999 to December 2000

Obs include children that had a chance to live >=356 days by the interview date.

Table 6: Birthweight Statistics

	Mean Birth Weight (Kg.)	% of Low Birthweight(<2.5kg)	Total Obs
Born during financial crisis	**3.13**	**8.7**	**841**
rural	3.17	6.9	390
urban	3.11	10.2	451
Born during 97 drought period	**3.16**	**6.7**	**386**
rural	3.15	8.1	186
urban	3.17	5.5	200
Born during non-crisis period			
1988-1992	**3.15**	**9.4**	**1568**
rural	3.11	11.1	624
urban	3.17	8.3	944
1993-1997, 1999	**3.18**	**8.1**	**2635**
rural	3.19	8.4	1245
urban	3.17	7.7	1390

financial crisis = January 1998 to June 1999

drought97 = May 1997 to December 1997

non-crisis = January 1988 to April 1997 and July 1999 to December 2000

Table 7: Mother's Education

Year	Mean (years)	No edu (%)	1-5 years (%)	6-8 years (%)	9-11 years (%)	12+ years (%)	Obs
1988	5.0	19.2	33.9	26.2	7.9	12.9	776
1989	5.5	15.2	32.8	26.9	9.6	15.6	845
1990	5.6	15.9	31.4	26.5	10.7	15.5	867
1991	5.8	14.2	27.2	30.1	12.1	16.5	832
1992	5.9	14.5	26.6	29.6	12.3	17.1	791
1993	6.1	11.7	27.8	30.0	13.4	17.2	746
1994	6.6	11.0	23.4	30.6	14.5	20.5	689
1995	7.2	9.5	28.0	32.1	14.0	26.5	794
1996	7.3	9.1	19.0	30.5	15.3	26.1	836
1997	7.6	8.1	15.9	30.2	17.7	28.1	718
1998	7.7	5.2	18.1	31.6	17.5	27.7	653
1999	8.2	4.9	14.8	28.2	20.0	32.1	614

Obs include mothers of children that had a chance to live >=356 days by the interview date.

Table 8: Child Mortality Rates (per 1,000 live births) by Mother's Education

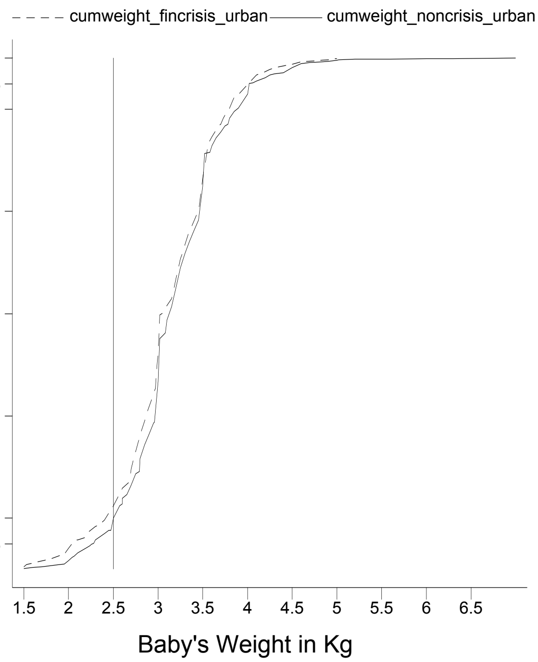
	Neonatal Mortality	Post-neonatal Mortality	Infant Mortality	Obs
No edu	34.1	40.6	74.7	1084
1-5 years	19.6	48.2	67.8	2241
6-8 years	23.5	20.9	44.4	2685
9-11 years	22.7	17.8	40.5	1236
12+ years	12.5	6.3	18.8	1915

Obs include children that had a chance to live >=356 days by the interview date.

sample: 1988-1999

Figure 6a-6c: Cummulative Distribution of Birthweight (Kg.)

Urban



Rural

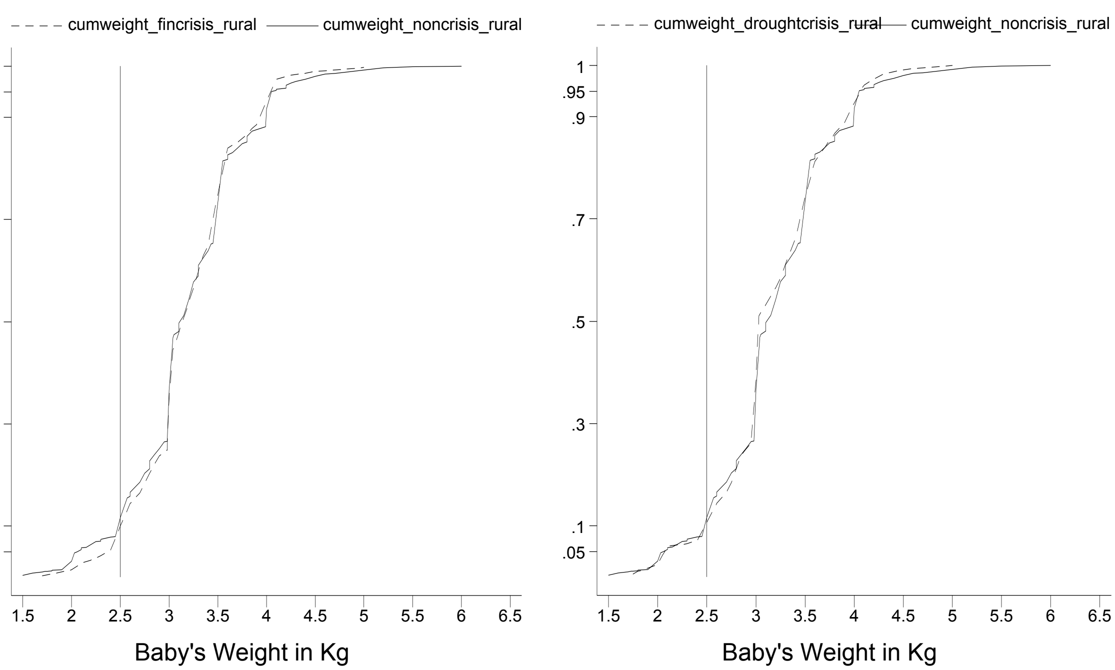


Table 9: Baby's Birthweight First and Second Order Stochastic Dominance

	First crossing point and difference between curves (first and second order stochastic dominance)											
	Urban				Rural							
	Financial Crisis - Noncrisis				Financial Crisis - Noncrisis				Drought Crisis - Noncrisis			
	s = 1	(sd)	s = 2	(sd)	s = 1	(sd)	s = 2	(sd)	s = 1	(sd)	s = 2	(sd)
First crossing point	4.499	(0.106)	-	-	3.310	(0.143)	-	-	2.100	0.683	-	-
Points of testing (Kg.)												
1.5	4.01	(5.08)	0.00	(0.00)	-6.13	(1.69)	0.00	(0.00)	-6.13	(1.69)	0.00	(0.00)
1.6	9.92	(6.36)	0.56	(0.53)	-8.48	(1.99)	-0.61	(0.17)	-8.48	(1.99)	-0.61	(0.17)
1.7	8.40	(6.40)	1.55	(1.12)	-7.78	(3.39)	-1.46	(0.35)	-10.37	(2.19)	-1.46	(0.35)
1.8	17.23	(8.65)	2.39	(1.75)	-11.08	(3.67)	-2.32	(0.61)	-2.80	(8.08)	-2.31	(0.62)
1.9	17.18	(8.94)	4.11	(2.45)	-12.02	(3.72)	-3.43	(0.94)	-3.74	(8.10)	-2.59	(1.25)
2.0	29.09	(11.50)	5.92	(3.23)	-24.99	(8.57)	-4.63	(1.29)	-13.10	(15.75)	-2.96	(2.02)
2.1	22.79	(12.42)	8.63	(4.11)	-29.94	(10.00)	-7.25	(1.85)	1.35	(19.56)	-3.87	(3.13)
2.2	29.24	(13.95)	10.95	(5.12)	-33.00	(11.12)	-10.27	(2.65)	-9.49	(19.68)	-3.76	(4.72)
2.3	25.87	(15.23)	13.68	(6.24)	-32.30	(12.06)	-13.62	(3.58)	-11.13	(20.41)	-4.75	(6.51)
2.4	24.45	(15.98)	16.26	(7.49)	-19.59	(13.54)	-16.85	(4.60)	-8.52	(21.06)	-5.86	(8.39)
2.5	22.52	(18.55)	18.67	(8.83)	-17.74	(19.45)	-18.83	(5.72)	-19.17	(27.31)	-6.74	(10.34)
2.6	20.50	(19.38)	20.91	(10.22)	-23.87	(20.01)	-20.62	(7.03)	-24.80	(28.17)	-8.67	(12.36)
2.7	39.61	(21.31)	22.97	(11.72)	-21.06	(21.54)	-23.00	(8.58)	-20.46	(30.32)	-11.15	(14.61)
2.8	41.09	(23.05)	27.29	(13.25)	-14.95	(23.41)	-25.51	(10.26)	-6.64	(33.08)	-13.33	(16.98)
2.9	65.61	(24.17)	31.26	(14.87)	-18.26	(24.06)	-27.03	(12.06)	-8.94	(34.12)	-13.74	(19.45)
3.0	48.33	(25.32)	37.77	(16.52)	-23.77	(27.16)	-28.94	(13.97)	35.56	(38.20)	-14.50	(22.10)
# observations	Financial Crisis:		450		Financial Crisis:		386		Financial Crisis:		184	
	Noncrisis:		2634		Noncrisis:		2122		Noncrisis:		2122	

Source: IFLS I, IFLS2, and IFLS3.

Standard errors reflect clustering at the mother level.

S=1 is the difference in cummulative distribution for the first order stochastice dominance (grams).

S=2 is the difference in the cummulative distribution for the second order stochastice dominance (grams).

Observations include birthweight betweeen 1.5 kilograms and 5.0 kilograms.

Financial Crisis = January 1998 - June 1999

Drougt Crisis = May 1997 - December 1997

Noncrisis = January 1988 - April 1997 and July 1999 - December 2000

Dash (-) indicates that the curves do not cross.

Formulation for the standard deviation is from Russel Davidson and Jean-Yves Duclos (2000), Statistical Inference for Stochastic Dominance and for the Measurement of Poverty and Inequality," Economometrica v86 n6. Computing for the table above was performed using "DAD: A software for Distributative Analysis/Analyse Distributative." copyright by Jean-Yves Duclos, Abdelkrim Araar, and Carl Fortin.

Table 10: Birthweight Regressions

Birthweight (Kg.)	Urban				Rural			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
OLS	weight	weight	weight	weight	weight	weight	weight	weight
t	-0.002	-0.001	-0.001	-0.001	0.005	0.005	0.006	0.005
	(0.44)	(0.33)	(0.25)	(0.20)	(1.11)	(1.14)	(1.29)	(1.06)
preg crisis1	-0.043	-0.043	-0.049	-0.025	0.000	-0.003	-0.002	0.009
	(0.97)	(0.98)	(1.11)	(0.52)	(0.01)	(0.06)	(0.04)	(0.17)
preg crisis2	0.032	0.031	0.028	0.046	-0.005	-0.012	-0.015	-0.013
	(0.85)	(0.80)	(0.74)	(1.05)	(0.13)	(0.30)	(0.37)	(0.27)
preg_cs97	-0.064	-0.062	-0.061	-0.073	-0.011	-0.016	-0.023	0.011
	(1.46)	(1.41)	(1.40)	(1.46)	(0.27)	(0.38)	(0.57)	(0.22)
male	0.073	0.074	0.069	0.070	0.067	0.066	0.064	0.055
	(3.53)***	(3.57)***	(3.37)***	(3.04)***	(2.76)***	(2.76)***	(2.67)***	(2.10)**
Mother's edu								
1-5 yrs		-0.094	-0.082	-0.088		0.103	0.064	0.017
		(1.26)	(1.14)	(1.12)		(1.64)	(1.00)	(0.24)
6-8 yrs		-0.129	-0.107	-0.117		0.049	0.019	-0.037
		(1.74)*	(1.49)	(1.53)		(0.84)	(0.33)	(0.55)
9-11 yrs		-0.106	-0.085	-0.128		0.118	0.088	0.022
		(1.41)	(1.19)	(1.65)		(1.89)*	(1.39)	(0.30)
12+ yrs		-0.103	-0.076	-0.102		0.056	0.012	-0.051
		(1.42)	(1.10)	(1.34)		(0.93)	(0.20)	(0.71)
Constant	6.333	5.712	5.213	4.947	-6.818	-7.159	-7.936	-7.729
	(0.86)	(0.76)	(0.70)	(0.59)	(0.76)	(0.80)	(0.90)	(0.75)
Province dummies	No	No	Yes	No	No	No	Yes	No
Community dummies	No	No	No	Yes	No	No	No	Yes
Observations	3293	3293	3293	3293	2701	2701	2701	2701
R-squared	0.010	0.010	0.030	0.200	0.000	0.010	0.030	0.220
p-value(crises)	0.196	0.222	0.207	0.192	0.994	0.981	0.945	0.973
p-value(mother's edu)		0.510	0.629	0.517		0.192	0.256	0.450

Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all livebirths born 1998-2000.
preg_cs1 = Pregnant in January 1998 - September 1998
preg_cs2 = Pregnant in October 1998 - June 1999
preg_cs97 = Pregnant in May 1997 - December 1997
"No-education" and "North Sumatra" are omitted categories.

Table 11: Mortality Linear Probability Regressions (Urban)

LPM	(1) Neonatal	(2) Neonatal	(3) Neonatal	(4) Neonatal	(1) Post-Neonatal	(2) Post-Neonatal	(3) Post-Neonatal	(4) Post-Neonatal	(1) Infant	(2) Infant	(3) Infant	(4) Infant
t (x100)	-0.261	-0.235	-0.228	-0.756	-0.252	-0.174	-0.174	-0.751	-0.482	-0.379	-0.360	-0.855
	(3.63)***	(3.25)***	(3.20)***	(2.56)**	(2.98)***	(2.15)**	(2.01)**	(2.06)**	(4.40)***	(3.55)***	(3.37)***	(3.08)***
preg crisis1 (x100)	1.769	1.757	1.640	5.011								
	(1.91)*	(1.90)*	(1.76)*	(1.50)								
preg crisis2 (x100)	0.828	0.802	0.913	3.488								
	(0.85)	(0.82)	(0.94)	(0.63)								
preg_cs97 (x100)	2.119	2.135	2.050	7.071								
	(1.87)*	(1.88)*	(1.86)*	(1.28)								
born crisis1 (x100)					1.781	1.759	1.757	3.385	3.502	3.492	3.240	5.356
					(1.61)	(1.60)	(1.47)	(0.68)	(2.33)**	(2.33)**	(2.18)**	(1.34)
born crisis2 (x100)					0.817	0.815	0.575	0.231	2.366	2.350	1.976	3.419
					(1.00)	(1.00)	(0.64)	(0.06)	(1.93)*	(1.92)*	(1.59)	(1.04)
born 97 crisis (x100)					-0.241	-0.269	-0.249	-0.962	-0.430	-0.468	-0.526	-0.351
					(0.39)	(0.43)	(0.37)	(0.26)	(0.40)	(0.43)	(0.47)	(0.10)
male (x100)	0.435	0.434	0.463	1.119	0.512	0.512	0.563	1.285	0.940	0.939	0.986	1.462
	(1.05)	(1.05)	(1.12)	(0.69)	(1.15)	(1.16)	(1.18)	(0.68)	(1.56)	(1.57)	(1.63)	(0.99)
Mother's edu (x100)												
1-5 yrs		-0.946	-0.498	-2.709		-0.283	0.150	3.328		-1.173	-0.186	0.177
		(0.68)	(0.39)	(0.75)		(0.15)	(0.08)	(0.80)		(0.54)	(0.09)	(0.05)
6-8 yrs		-1.666	-1.139	-3.830		-2.461	-2.118	-3.036		-4.022	-2.919	-4.552
		(1.26)	(0.94)	(1.10)		(1.48)	(1.25)	(0.83)		(1.99)**	(1.60)	(1.50)
9-11 yrs		-0.999	-0.622	-1.862		-2.824	-2.426	-3.889		-3.715	-2.691	-3.690
		(0.72)	(0.49)	(0.48)		(1.67)*	(1.42)	(0.95)		(1.80)*	(1.45)	(1.08)
12+ yrs		-1.936	-1.605	-6.404		-3.537	-3.167	-8.194		-5.340	-4.330	-8.314
		(1.48)	(1.37)	(1.75)*		(2.16)**	(1.91)*	(2.11)**		(2.71)***	(2.46)**	(2.59)***
Constant	5.224	4.717	4.557	15.163	5.043	3.499	3.512	15.062	9.633	7.621	7.213	17.154
	(3.63)***	(3.27)***	(3.21)***	(2.58)**	(2.99)***	(2.18)**	(2.03)**	(2.07)**	(4.41)***	(3.58)***	(3.38)***	(3.10)***
Province dummies	No	No	Yes	No	No	No	Yes	No	No	No	Yes	No
Community dummies	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes
Observations	4032	4032	4032	997	3961	3961	3731	890	4032	4032	4032	1586
R-squared	0.000	0.010	0.010	0.070	0.000	0.010	0.020	0.070	0.010	0.010	0.020	0.080
p-value(crises)	0.091	0.093	0.104	0.310	0.253	0.249	0.382	0.895	0.031	0.031	0.062	0.468
p-value(mother's edu)		0.253	0.262	0.247		0.000	0.001	0.003		0.000	0.000	0.001

Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97

Table 12: Mortality Logistic Regressions (Urban)

Logit	(1) Neonatal	(2) Neonatal	(3) Neonatal	(4) Neonatal	(1) Post-Neonatal	(2) Post-Neonatal	(3) Post-Neonatal	(4) Post-Neonatal	(1) Infant	(2) Infant	(3) Infant	(4) Infant
t	0.851	0.863	0.867	0.875	0.866	0.903	0.907	0.894	0.864	0.889	0.893	0.889
	(3.81)***	(3,38)***	(3.31)***	(2.44)**	(2.99)***	(2.06)**	(1.93)*	(2.08)**	(4.41)***	(3.47)***	(3.29)***	(3.06)***
preg crisis1	3.262	3.271	3.006	2.542								
	(2.09)**	(2.09)**	(1.91)*	(1.36)								
preg crisis2	1.586	1.578	1.648	1.994								
	(0.43)	(0.43)	(0.47)	(0.60)								
preg_cs97	3.761	3.809	3.649	3.142								
	(2.40)**	(2.41)**	(2.37)**	(1.65)*								
born crisis1					2.901	2.931	2.706	1.891	3.018	3.048	2.823	2.119
					(1.86)*	(1.86)*	(1.66)*	(0.85)	(2.73)***	(2.73)***	(2.52)**	(1.49)
born crisis2					1.596	1.654	1.396	1.030	2.211	2.246	1.971	1.659
					(0.67)	(0.72)	(0.46)	(0.04)	(1.78)*	(1.81)*	(1.47)	(0.97)
born 97 crisis					0.540	0.538	0.547	0.740	0.559	0.557	0.531	0.802
					(0.59)	(0.59)	(0.57)	(0.28)	(0.57)	(0.57)	(0.61)	(0.29)
male	1.289	1.283	1.312	1.198	1.333	1.321	1.337	1.130	1.317	1.309	1.336	1.182
	(1.04)	(1.02)	(1.11)	(0.68)	(1.12)	(1.08)	(1.10)	(0.47)	(1.53)	(1.49)	(1.57)	(0.89)
Mother's edu												
1-5 yrs		0.708	0.898	0.652		0.928	1.153	1.435		0.831	1.050	1.039
		(0.74)	(0.25)	(0.80)		(0.16)	(0.33)	(0.77)		(0.56)	(0.15)	(0.11)
6-8 yrs		0.484	0.635	0.520		0.413	0.532	0.672		0.437	0.572	0.599
		(1.54)	(0.99)	(1.20)		(1.92)*	(1.41)	(0.81)		(2.46)**	(1.72)*	(1.38)
9-11 yrs		0.709	0.848	0.731		0.322	0.428	0.582		0.482	0.613	0.680
		(0.70)	(0.36)	(0.55)		(2.06)**	(1.60)	(0.96)		(2.01)**	(1.41)	(0.96)
12+ yrs		0.384	0.447	0.324		0.135	0.181	0.252		0.238	0.298	0.308
		(1.97)**	(1.73)*	(1.84)*		(3.38)***	(2.90)***	(2.01)**		(3.90)***	(3.39)***	(2.60)***
Province dummies	No	No	Yes	No	No	No	Yes	No	No	No	Yes	No
Conditional Logit	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes
Observations	4032	4032	4032	997	3961	3961	3731	890	4032	4032	4032	1586
Log-Likelihood	-350.07	-346.91	-332.57	-172.35	-352.15	-336.47	-324.11	-167.50	-603.55	-588.41	-574.21	-326.27
Pseudo R-Squared	0.020	0.029	0.069	0.039	0.022	0.065	0.088	0.067	0.023	0.048	0.071	0.048
p-value(crises)	0.054	0.054	0.070	0.333	0.227	0.224	0.310	0.824	0.028	0.027	0.053	0.415
p-value(mother's edu)		0.206	0.281	0.311		0.000	0.001	0.021		0.000	0.000	0.011

Partial odd ratios are reported.
Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97
"No-edu" and "North Sumatra" are omitted categories.

Table 13: Mortality Linear Probability Regressions (Rural)

	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
LPM	Neonatal	Neonatal	Neonatal	Neonatal	Post-Neonatal	Post-Neonatal	Post-Neonatal	Post-Neonatal	Infant	Infant	Infant	Infant
t (x100)	-0.172	-0.174	-0.176	-0.425	-0.265	-0.166	-0.176	-0.329	-0.468	-0.371	-0.376	-0.581
	(2.33)**	(2.37)**	(2.40)**	(2.54)**	(2.78)***	(1.72)*	(1.79)*	(1.92)*	(4.01)***	(3.15)***	(3.22)***	(3.57)***
preg crisis1 (x100)	2.187	2.183	2.184	5.448								
	(1.76)*	(1.75)*	(1.76)*	(1.83)*								
preg_cs97 (x100)	0.226	0.232	0.317	-1.210								
	(0.25)	(0.25)	(0.35)	(0.68)								
born crisis1 (x100)					1.676	1.555	1.848	2.836	2.059	1.941	2.224	1.974
					(1.34)	(1.24)	(1.43)	(1.29)	(1.39)	(1.31)	(1.51)	(0.97)
born crisis2 (x100)					0.906	0.857	0.788	0.942	3.352	3.307	3.223	4.953
					(0.82)	(0.78)	(0.70)	(0.41)	(2.06)**	(2.03)**	(2.00)**	(1.96)**
born 97 crisis (x100)					2.747	2.848	3.109	6.345	4.206	4.277	4.434	7.707
					(1.99)**	(2.07)**	(2.22)**	(2.39)**	(2.39)**	(2.43)**	(2.53)**	(2.92)***
male (x100)	1.038	1.035	1.006	2.055	1.419	1.470	1.531	2.218	2.389	2.431	2.407	2.930
	(2.38)**	(2.36)**	(2.30)**	(2.20)**	(2.64)***	(2.73)***	(2.77)***	(2.33)**	(3.54)***	(3.60)***	(3.57)***	(3.20)***
Mother's edu (x100)												
1-5 yrs		-1.607	-1.838	-4.375		1.057	1.332	2.357		-0.486	-0.424	-0.572
		(1.73)*	(1.91)*	(2.29)**		(1.11)	(1.36)	(1.54)		(0.37)	(0.32)	(0.35)
6-8 yrs		-0.380	-0.477	-1.129		-1.737	-1.352	-2.215		-2.048	-1.754	-2.208
		(0.39)	(0.49)	(0.54)		(2.06)**	(1.53)	(1.51)		(1.64)	(1.38)	(1.30)
9-11 yrs		-0.680	-0.804	-0.372		-1.630	-1.111	-1.256		-2.233	-1.822	-0.852
		(0.62)	(0.72)	(0.15)		(1.58)	(1.01)	(0.59)		(1.48)	(1.19)	(0.38)
12+ yrs		-1.585	-1.726	-2.305		-3.242	-2.719	-4.660		-4.709	-4.207	-4.434
		(1.53)	(1.58)	(0.97)		(3.89)***	(3.03)***	(2.19)**		(3.60)***	(3.08)***	(2.08)**
Constant	3.448	3.485	3.529	8.537	5.314	3.333	3.517	6.598	9.364	7.440	7.546	11.640
	(2.35)**	(2.38)**	(2.42)**	(2.56)**	(2.79)***	(1.74)*	(1.80)*	(1.94)*	(4.03)***	(3.17)***	(3.24)***	(3.60)***
Province dummies	No	No	Yes	No	No	No	Yes	No	No	No	Yes	No
Community dummies	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes
Observations	5100	5100	5100	2241	4978	4978	4844	2738	5100	5100	5100	3619
R-squared	0.000	0.000	0.010	0.070	0.000	0.010	0.020	0.070	0.010	0.010	0.020	0.070
p-value(crises)	0.213	0.215	0.212	0.128	0.152	0.146	0.096	0.085	0.025	0.026	0.019	0.009
p-value(mother's edu)		0.139	0.102	0.022		0.000	0.000	0.001		0.000	0.003	0.167

Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97
"No-edu" and "North Sumatra" are omitted categories.

Table 14: Mortality Logistic Regressions (Rural)

Logit	(1) Neonatal	(2) Neonatal	(3) Neonatal	(4) Neonatal	(1) Post-Neonatal	(2) Post-Neonatal	(3) Post-Neonatal	(4) Post-Neonatal	(1) Infant	(2) Infant	(3) Infant	(4) Infant
t	0.926	0.926	0.925	0.912	0.918	0.945	0.942	0.937	0.911	0.927	0.926	0.913
	(2.37)**	(2.37)**	(2.40)**	(2.46)**	(2.77)***	(1.80)*	(1.89)*	(1.92)*	(3.99)***	(3.18)***	(3.26)***	(3.43)***
preg crisis1	2.457	2.460	2.462	2.654								
	(2.18)**	(2.18)**	(2.19)**	(2.13)**								
preg_cs97	1.082	1.078	1.125	0.580								
	(0.14)	(0.14)	(0.21)	(0.71)								
born crisis1					1.767	1.732	1.872	1.613	1.548	1.522	1.608	1.274
					(1.43)	(1.37)	(1.55)	(1.04)	(1.34)	(1.28)	(1.45)	(0.63)
born crisis2					1.349	1.351	1.287	1.176	1.996	1.997	1.968	2.132
					(0.67)	(0.67)	(0.56)	(0.33)	(2.21)**	(2.20)**	(2.15)**	(2.21)**
born 97 crisis					2.269	2.372	2.567	2.886	2.218	2.261	2.353	2.924
					(2.44)**	(2.55)**	(2.78)***	(2.89)***	(2.87)***	(2.91)***	(3.04)***	(3.69)***
male	1.580	1.577	1.565	1.534	1.563	1.593	1.605	1.511	1.581	1.597	1.597	1.540
	(2.41)**	(2.38)**	(2.34)**	(2.09)**	(2.61)***	(2.71)***	(2.73)***	(2.41)**	(3.54)***	(3.60)***	(3.58)***	(3.25)***
Mother's edu												
1-5 yrs		0.510	0.459	0.404		1.275	1.393	1.498		0.934	0.945	0.953
		(1.95)*	(2.18)**	(2.93)***		(1.11)	(1.45)	(1.66)*		(0.35)	(0.28)	(0.25)
6-8 yrs		0.893	0.853	0.872		0.572	0.647	0.625		0.705	0.748	0.750
		(0.36)	(0.50)	(0.48)		(2.19)**	(1.65)*	(1.61)		(1.72)*	(1.39)	(1.39)
9-11 yrs		0.796	0.749	1.013		0.595	0.705	0.803		0.674	0.728	0.930
		(0.58)	(0.71)	(0.03)		(1.49)	(0.98)	(0.57)		(1.43)	(1.14)	(0.26)
12+ yrs		0.494	0.457	0.660		0.204	0.258	0.348		0.321	0.354	0.449
		(1.48)	(1.54)	(0.80)		(3.26)***	(2.71)***	(1.99)**		(3.31)***	(2.92)***	(2.15)**
Province dummies	No	No	Yes	No	No	No	Yes	No	No	No	Yes	No
Conditional Logit	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes
Observations	5100	5100	5100	2241	4978	4978	4844	2738	5100	5100	5100	3619
Log-Likelihood	-569.16	-564.43	-558.80	-333.74	-725.42	-708.04	-689.04	-453.80	-1097.35	-1086.87	-1068.60	-742.54
Pseudo R-Squared	0.012	0.020	0.030	0.039	0.012	0.036	0.056	0.039	0.014	0.024	0.040	0.025
p-value(crises)	0.088	0.088	0.088	0.057	0.085	0.072	0.036	0.034	0.015	0.014	0.011	0.002
p-value(mother's edu)		0.171	0.118	0.022		0.000	0.000	0.001		0.007	0.028	0.166

Partial odd ratios are reported.
Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97
"No-edu" and "North Sumatra" are omitted categories.

Table 15: Infant Mortality : Low Education and High Education

Urban	Mother's education = 0 - 8 years			Mother's education >= 9 years		
LPM	(1)	(2)	(3)	(1)	(2)	(3)
t (x100)	-0.594	-0.537	-1.172	-0.271	-0.268	-0.654
	(3.34)***	(3.04)***	(3.06)***	(2.22)**	(2.20)**	(1.57)
born crisis1 (x100)	6.624	6.177	13.530	1.172	1.457	-1.167
	(2.18)**	(2.09)**	(1.88)*	(0.83)	(1.01)	(0.25)
born crisis2 (x100)	2.712	1.852	1.662	1.958	1.943	5.603
	(1.16)	(0.74)	(0.36)	(1.41)	(1.38)	(1.18)
born cs97 (x100)	0.729	0.826	6.008	-1.173	-1.175	-3.411
	(0.27)	(0.31)	(0.90)	(2.38)**	(2.26)**	(1.39)
male (x100)	0.001	0.164	-0.702	1.848	1.857	4.455
	0.00	(0.16)	(0.34)	(2.82)***	(2.86)***	(2.10)**
Constant	11.885	10.716	23.450	5.417	5.343	13.077
	(3.35)***	(3.04)***	(3.08)***	(2.23)**	(2.20)**	(1.58)
Province dummies	No	Yes	No	No	Yes	No
Community dummies	No	No	Yes	No	No	Yes
Observations	2006	2006	953	2026	2026	638
R-squared	0.010	0.030	0.160	0.010	0.020	0.170
p-value (crises)	0.189	0.172	0.699	0.225	0.177	0.193

Rural	Mother's education = 0 - 8 years			Mother's education >= 9 years		
LPM	(1)	(2)	(3)	(1)	(2)	(3)
t (x100)	-0.431	-0.432	-0.634	-0.386	-0.408	-0.741
	(3.12)***	(3.13)***	(3.63)***	(1.80)*	(2.02)**	(1.70)*
born crisis1 (x100)	2.538	2.714	1.400	0.734	1.721	5.154
	(1.31)	(1.40)	(0.59)	(0.37)	(0.87)	(1.29)
born crisis2 (x100)	3.229	3.086	4.400	3.575	3.337	6.696
	(1.55)	(1.49)	(1.48)	(1.41)	(1.32)	(1.12)
born cs97 (x100)	5.830	6.034	11.191	1.134	1.288	1.163
	(2.42)**	(2.51)**	(3.26)***	(0.55)	(0.63)	(0.29)
male (x100)	2.976	2.950	3.386	0.452	0.461	-0.130
	(3.70)***	(3.67)***	(3.32)***	(0.40)	(0.41)	(0.06)
Constant	8.622	8.631	12.701	7.731	8.169	14.826
	(3.13)***	(3.14)***	(3.64)***	(1.81)*	(2.03)**	(1.71)*
Province dummies	No	Yes	No	No	Yes	No
Community dummies	No	No	Yes	No	No	Yes
Observations	3982	3982	3014	1118	1118	613
R-squared	0.010	0.020	0.110	0.000	0.030	0.160
p-value (crises)	0.038	0.031	0.006	0.575	0.567	0.510

Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97
"North Sumatra" is omitted.

Table 16: Neonatal Mortality : Low Education and High Education

Urban	Mother's education = 0 - 8 years			Mother's education >= 9 years		
LPM	(1)	(2)	(3)	(1)	(2)	(3)
t (x100)	-0.239	-0.199	-0.685	-0.273	-0.285	-1.276
	(2.20)**	(1.93)*	(1.68)*	(2.82)***	(2.96)***	(3.08)***
preg crisis1 (x100)	1.966	1.684	0.826	1.773	1.833	8.920
	(1.10)	(0.91)	(0.18)	(1.67)*	(1.69)*	(1.80)*
preg crisis2 (x100)	-0.330	-0.537	-0.362	1.528	1.715	12.461
	(0.48)	(0.72)	(0.10)	(1.04)	(1.16)	(1.01)
preg_cs97 (x100)	3.180	3.103	10.963	1.460	1.597	4.012
	(1.43)	(1.45)	(1.14)	(1.23)	(1.36)	(0.52)
male (x100)	-0.034	0.070	0.927	0.910	0.930	2.936
	(0.05)	(0.11)	(0.43)	(1.68)*	(1.72)*	(1.08)
Constant	4.781	3.974	13.717	5.448	5.680	25.487
	(2.21)**	(1.93)*	(1.69)*	(2.82)***	(2.96)***	(3.08)***
Province dummies	No	Yes	No	No	Yes	No
Community dummies	No	No	Yes	No	No	Yes
Observations	2006	2006	565	2026	2026	434
R-squared	0.000	0.020	0.150	0.000	0.010	0.150
p-value (crises)	0.462	0.496	0.501	0.051	0.086	0.312

Rural	Mother's education = 0 - 8 years			Mother's education >= 9 years		
LPM	(1)	(2)	(3)	(1)	(2)	(3)
t (x100)	-0.001	-0.001	-0.003	-0.002	-0.002	-0.006
	(1.36)	(1.36)	(1.53)	(1.15)	(1.28)	(1.46)
preg crisis1 (x100)	0.021	0.021	0.049	0.019	0.020	0.060
	(1.26)	(1.23)	(1.29)	(1.05)	(1.11)	(0.92)
preg crisis2 (x100)	-0.016	-0.020	-0.037	-0.010	-0.010	-0.037
	(2.45)**	(2.80)***	(1.87)*	(1.11)	(0.99)	(0.96)
preg_cs97 (x100)	0.007	0.007	-0.014	-0.013	-0.008	0.000
	(0.50)	(0.54)	(0.64)	(1.60)	(1.05)	(0.01)
male (x100)	0.014	0.014	0.027	-0.004	-0.004	-0.020
	(2.80)***	(2.74)***	(2.60)***	(0.43)	(0.42)	(0.86)
Constant	2.592	2.595	6.146	3.865	4.198	12.289
	(1.36)	(1.37)	(1.54)	(1.16)	(1.29)	(1.46)
Province dummies	No	Yes	No	No	Yes	No
Community dummies	No	No	Yes	No	No	Yes
Observations	3982	3982	1850	1118	1118	396
R-squared	0.000	0.010	0.110	0.000	0.020	0.160
p-value (crises)	0.003	0.002	0.073	0.001	0.232	0.526

Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97
"North Sumatra" is omitted.

Table 17: Post-neonatal Mortality : Low Education and High Education

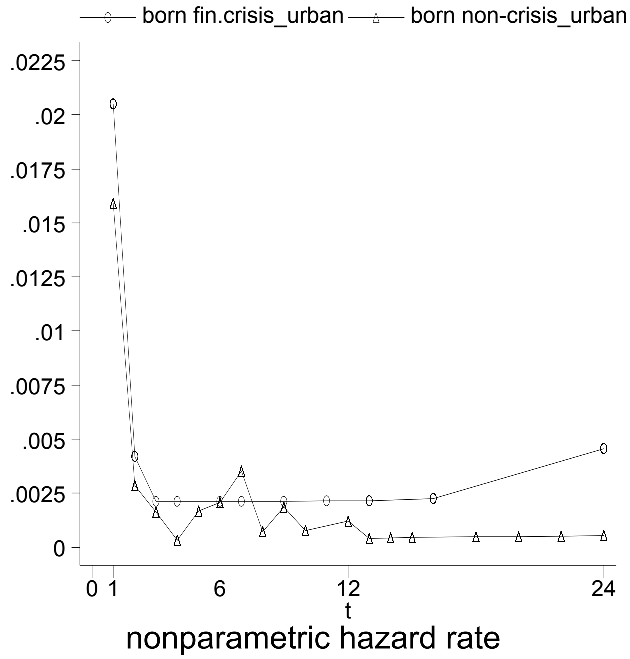
Urban	Mother's education = 0 - 8 years			Mother's education >= 9 years		
LPM	(1)	(2)	(3)	(1)	(2)	(3)
t (x100)	-0.350	-0.332	-1.156	-0.059	-0.048	0.066
	(2.33)**	(2.19)**	(2.24)**	(0.85)	(0.68)	(0.13)
born crisis1 (x100)	3.856	3.686	13.298	0.195	0.352	-4.331
	(1.59)	(1.51)	(1.46)	(0.23)	(0.41)	(1.06)
born crisis2 (x100)	0.650	-0.017	1.907	0.644	0.595	1.755
	(0.41)	(0.01)	(0.33)	(0.74)	(0.70)	(0.34)
born cs97 (x100)	0.157	0.437	3.888	-0.537	-0.500	-4.440
	(0.11)	(0.29)	(0.64)	(1.91)*	(1.72)*	(1.53)
male (x100)	0.031	0.077	-1.990	0.958	0.939	5.434
	(0.04)	(0.09)	(0.76)	(2.48)**	(2.47)**	(2.32)**
Constant	6.998	6.637	23.128	1.181	0.970	-1.298
	(2.34)**	(2.19)**	(2.25)**	(0.85)	(0.69)	(0.13)
Province dummies	No	Yes	No	No	Yes	No
Community dummies	No	No	Yes	No	No	Yes
Observations	1965	1965	585	1996	1996	308
R-squared	0.000	0.020	0.140	0.000	0.010	0.230
p-value (crises)	0.144	0.211	0.263	0.000	0.001	0.106

Rural	Mother's education = 0 - 8 years			Mother's education >= 9 years		
LPM	(1)	(2)	(3)	(1)	(2)	(3)
t (x100)	-0.244	-0.241	-0.435	-0.127	-0.133	-0.326
	(2.15)**	(2.12)**	(2.38)**	(0.78)	(0.86)	(0.64)
born crisis1 (x100)	1.633	1.778	2.097	1.517	2.069	5.752
	(1.02)	(1.11)	(0.85)	(0.80)	(1.09)	(1.24)
born crisis2 (x100)	0.669	0.514	0.463	1.175	0.898	2.473
	(0.47)	(0.37)	(0.18)	(0.68)	(0.51)	(0.42)
born cs97 (x100)	4.161	4.291	9.276	0.004	0.041	0.281
	(2.13)**	(2.21)**	(2.74)***	0.00	(0.03)	(0.07)
male (x100)	1.629	1.634	2.129	0.844	0.879	1.938
	(2.49)**	(2.49)**	(2.01)**	(1.06)	(1.12)	(0.79)
Constant	4.892	4.814	8.725	2.539	2.663	6.532
	(2.16)**	(2.13)**	(2.40)**	(0.78)	(0.86)	(0.64)
Province dummies	No	Yes	No	No	Yes	No
Community dummies	No	No	Yes	No	No	Yes
Observations	3882	3882	2320	1096	1096	421
R-squared	0.000	0.010	0.100	0.000	0.030	0.160
p-value (crises)	0.158	0.128	0.051	0.802	0.720	0.664

Absolute values of t-statistics robust to heteroskedasticity and clustering at mother level in parentheses.
* significant at 10%;** significant at 5%; *** significant at 1%
Observations include all individuals that were born 1988 - 2000 who were exposed to at least 365 days of life.
crisis1 = Jan 98 - Sept 98, crisis2 = Oct 98 - June 99, crisis97 = May 97 - Dec 97
"North Sumatra" is omitted.

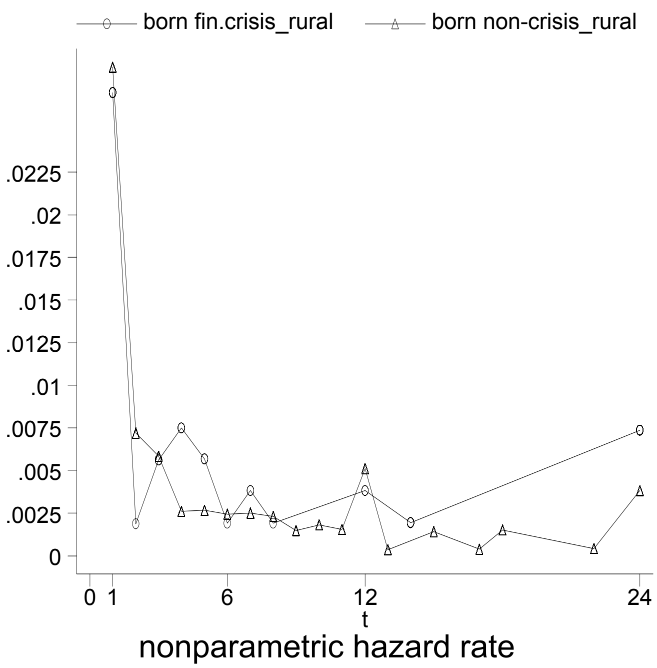
Figure 7a-7c: Non-parametric Hazard Rates (by Month)

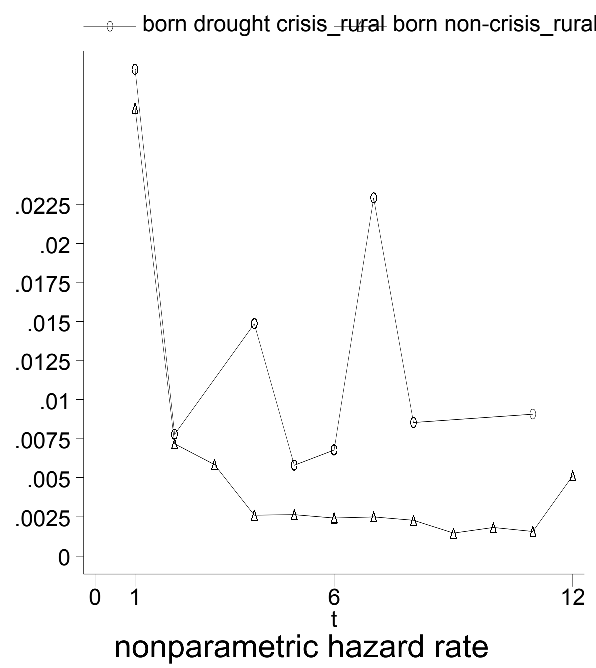
Urban



Log-rank test for equality of survivor functions: Pr>chi2 = 0.358

Rural





Log-rank test for equality of survivor functions: Pr>chi2 = 0.823

Log-rank test for equality of survivor functions: Pr>chi2 = 0.187

Table 18: Relative Risk Coefficients (Urban)

Weibull Regressions	Log Relative-hazard Form	Log Relative-hazard Form with Inverse Gaussian Heterogeneity
	Hazard Ratio (Robust Standard Error)	Hazard Ratio (Robust Standard Error)
t	0.970	0.945
	(0.029)	(0.040)
born financial crisis	1.721	2.232
	(0.503)*	(0.958)*
born drought crisis	2.746	4.735
	(1.000)***	(2.674)***
mother's edu		
1-5 years	1.057	0.993
	(0.352)	(0.526)
6-8 years	0.719	0.549
	(0.250)	(0.289)
9-11 years	0.836	0.718
	(0.300)	(0.395)
12+ years	0.409	0.265
	(0.145)**	(0.140)**
male	1.399	1.834
	(0.252)*	(0.484)**
/ ln p	-0.852	-0.400
	(0.038)***	(0.042)***
Log Likelihood	-843.03	-833.57
Observations	3956	3956
Times at risk	117276	117276

Standard errors are robust to heteroskedasticity clustering at the mother level.

Province dummies are included in the estimations, but not reported.

Table 19: Relative Risk Coefficients (Rural)

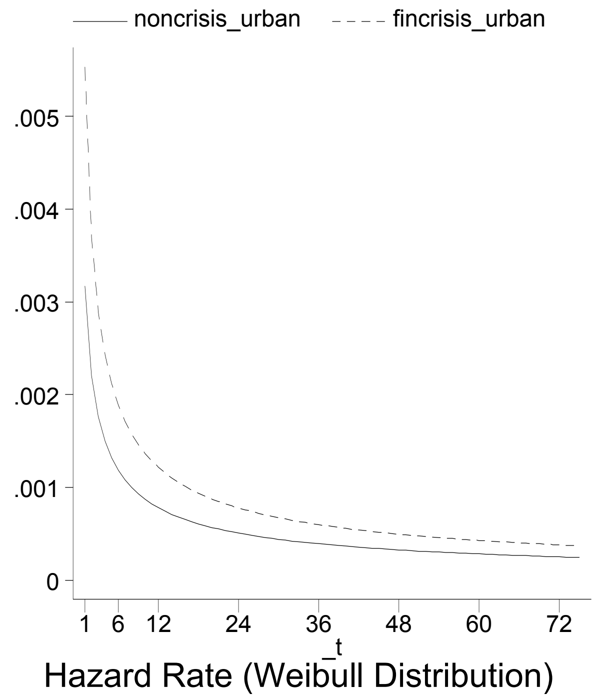
Weibull Regressions	Log Relative-hazard Form	Log Relative-hazard Form with Inverse Gaussian Heterogeneity
	Hazard Ratio (Robust Standard Error)	Hazard Ratio (Robust Standard Error)
t	1.017	1.013
	(0.020)	(0.029)
born financial crisis	1.105	1.159
	(0.225)	(0.363)
born drought crisis	1.621	2.144
	(0.411)*	(0.823)**
mother's edu		
1-5 years	0.744	0.596
	(0.123)*	(0.154)*
6-8 years	0.537	0.388
	(0.094)***	(0.104)***
9-11 years	0.539	0.395
	(0.130)**	(0.143)***
12+ years	0.275	0.152
	(0.082)***	(0.063)***
male	1.402	1.652
	(0.162)***	(0.287)***
/ ln p	-0.806	-0.355
	(0.027)***	(0.030)***
Log Likelihood	-1788.26	-1767.35
Observations	5033	5033
Times at risk	144635	144635

Standard errors are robust to heteroskedasticity clustering at the mother level.

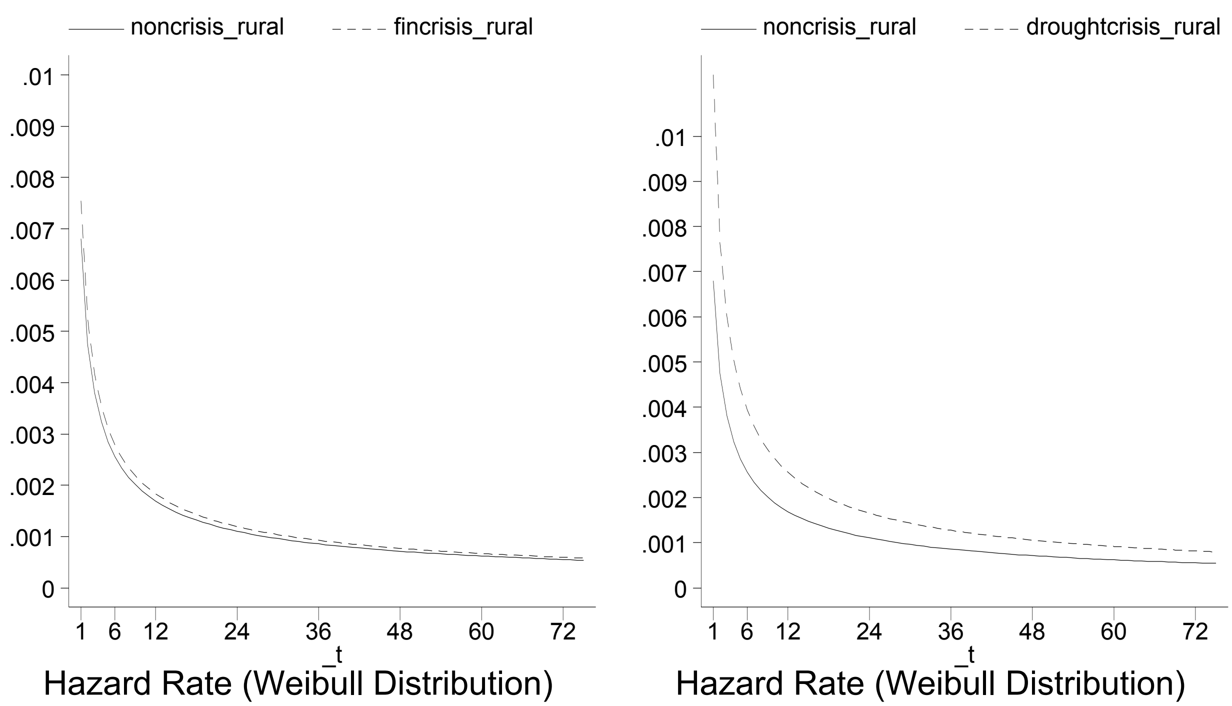
Province dummies are included in the estimations, but not reported.

Figure 8a-8c: Parametric Estimated Hazard Rate (Weibull Distribution)

Urban



Rural



Appendix I : Mortality Crosswalk from the Indonesian Family Life Surveys (IFLS1-IFLS3)

ch00=1	panel respondent with no preprinted child roster
ch00=2	panel respondent with preprinted child roster
ch00=3	new respondent

	Number of Obs		
FILS2 and 3 (1995-2000)	Neonatal	Post-neonatal	Total live births
All obs	59	69	3093

	Number of Obs		
FILS3 (1995-2000)	Neonatal	Post-neonatal	Total live births
All obs	41	49	2034
ch00==2 or 3	29	36	1443
ch00==1	12	13	591

	Rates (per 1000 live births)	
FILS3 (1995-2000)	Neonatal	Post-neonatal
All obs	20.2	24.1
ch00==2 or 3	20.1	24.9
ch00==1	20.3	22.0

Number of Obs by Year

	IFLS3 (1995)			IFLS2 (1995)		
	Neonatal	Post-neonatal	Total live births	Neonatal	Post-neonatal	Total live births
All obs	3	5	214	12	10	582
ch00==2 or 3	3	4	200	8	5	296
ch00==1	0	1	14	4	5	286
	IFLS3 (1996)			IFLS2 (1996)		
	Neonatal	Post-neonatal	Total live births	Neonatal	Post-neonatal	Total live births
All obs	4	4	271	6	10	472
ch00==2 or 3	3	1	239	0	7	253
ch00==1	1	3	32	0	3	219
	IFLS3 (1997)					
	Neonatal	Post-neonatal	Total live births	Neonatal	Post-neonatal	Total live births
All obs	9	11	297	0	0	5
ch00==2 or 3	6	9	245	0	0	5
ch00==1	3	2	52	0	0	0
	IFLS3 (1998)			IFLS2 (1998)		
	Neonatal	Post-neonatal	Total live births	Neonatal	Post-neonatal	Total live births
All obs	15	16	656	NA	NA	NA
ch00==2 or 3	10	14	390	NA	NA	NA
ch00==1	5	2	266	NA	NA	NA
	IFLS3 (1999)			IFLS2 (1999)		
	Neonatal	Post-neonatal	Total live births	Neonatal	Post-neonatal	Total live births
All obs	10	13	596	NA	NA	NA
ch00==2 or 3	7	8	369	NA	NA	NA
ch00==1	3	5	227	NA	NA	NA

NOTE:

Too few obs to calculate and compare mortality rates for each year

No obs in born in 2000 that had lived 365 days by the interview date of IFLS3

Similar reason for IFLS2 data with those born in 1997

(5 obs of 1997-borns were interviewed in 1998. These have age>=365 day by the interview)