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## Abstract

Health and income are strongly correlated both within and across countries, yet the extent to which improvements in income have a causal effect on health status remains controversial. We revisit this question with an investigation of short-term fluctuations in aggregate income and infant mortality using an unusually large dataset of 59 developing countries, covering over 1.7 million births. Our focus is on the effect of departures of income from trend on infant mortality, rather than on the relationship between long-term changes in income and infant mortality. Given that we use unit data, rather than country averages, we can control for the changing composition of women giving birth, and assess how aggregate income shocks interact with a variety of characteristics of mothers and children, such as mother's education and the gender of the child. We show that there is a large, negative association between per capita GDP and infant mortality—on average, a one percent decrease in per capita GDP is associated with an increase in mortality of between 0.24 and 0.40 infants per 1,000 children born. Female infant mortality is more sensitive than male infant mortality to economic fluctuations, especially during negative shocks to GDP, suggesting that policies that protect the health status of female infants may be especially important during economic downturns in much of the developing world.

## 1. Introduction

Health and income are strongly correlated across countries and, within countries, across individuals. In the United States, the life expectancy of people in the lowest ventile of the income distribution was about 25 percent lower than that of people in the highest ventile in 1980 (Rogot et al. 1992, cited in Cutler, Deaton, and Lleras-Muney 2006, pp. 111-112). In developing countries, dozens of studies have found that people with higher incomes have better health status and lower mortality (see Gwatkin et al. 2007 for a review). The seminal work by Preston (1975, 1980) shows that, as countries become richer, life expectancy rises, although many other factors are also important in explaining mortality declines.

Despite the association between income and health status, the extent to which improvements in income have a causal effect on health status remains controversial. In an early, influential article using cross-country data, Pritchett and Summers (1996) argued that "wealthier is healthier" but their identification and conclusions have been challenged by, among others, Jamison et al. (2004) and Deaton (2006). Part of the concern is the existence of feedbacks from health to income—for example, both Gallup and Sachs (2001) and the World Health Organization (2001) argue that improvements in health status would increase rates of economic growth. Countries with higher income levels also tend to have higher education levels, better-functioning health systems, and better institutions—all of which are likely to improve health outcomes independently of income (Cutler, Deaton and Lleras-Muney 2006).

In this paper, we revisit the discussion of the relationship between health and income with an investigation of the impact of short term fluctuations in per capita GDP on infant mortality for a large set of developing countries between 1975 and 2004. Infant mortality is pervasive in the developing world: In poor countries, approximately 30 percent of all deaths occur to children under the age of five, compared to less than 1 percent in rich countries (see Cutler, Deaton, and Lleras-Muney 2006, pp. 106-07). Infant mortality is also much less likely than adult mortality to be affected by reverse causality from health to income.

Our focus in this paper is on departures of income from trend, and the effect that these have on infant mortality, rather than on the relationship between long-term changes in income and infant mortality. This is an important distinction—even if long term improvements in infant mortality are primarily caused by improvements in medical technology, rather than by economic growth, short term shocks to GDP could have important consequences for child health. However, the effect of income shocks on infant mortality is hard to sign ex ante. In developing countries, negative shocks will reduce household consumption of nutritious foods, lower expenditures on other inputs into child health, and may involve serious disruptions of public health services; all of these would tend to increase infant mortality. On the other hand, aggregate shocks depress wages and may imply a lower opportunity cost of women's time. Many inputs into the production of child health are intensive in parental (especially maternal) time, including taking children for preventive health visits, breastfeeding, cooking healthy meals, or collecting clean water; because systemic (as opposed to idiosyncratic) shocks reduce the cost of engaging in these activities, they may improve child outcomes. The effect of negative income shocks on child health and mortality is therefore ambiguous in theory.

Since this paper focuses on the effect of GDP shocks on infant mortality, it is closely related to a literature on the health consequences of "booms" and "busts" in aggregate income. Dehejia and Lleras-Muney (2004) conclude that infant mortality is generally pro-cyclical in the United States. A variety of transmission mechanisms have been proposed to explain why economic recessions lead to improved child health, including reductions in air pollution (Chay and Greenstone 2003), reductions in health-damaging behaviors such as smoking and drinking, and increases in the probability that mothers engage in time-intensive activities such as exercise and prenatal care (Ruhm 2000; Ruhm and Black 2002).

In developing countries, our focus in this paper, the evidence on the relationship between economic downturns and infant mortality is more mixed—see the review by Ferreira and Schady (2009). Sharp economic downturns were associated with increases in infant mortality in Mexico (Cutler et al. 2002), Peru (Paxson and Schady 2005) and India (Bhalotra 2008). On the other hand, Miller and Urdinola (2007) find that arguably exogenous declines in the price of coffee, which resulted in aggregate income decreases in coffee-growing areas in Colombia, were associated with lower infant mortality, echoing the results from the US.

Our paper extends the existing literature in a number of important ways. The sample of countries, 59, is much larger than that from the country-specific studies. This allows us to estimate the effect of aggregate income shocks on health in a variety of settings across recent decades. In addition, and unlike the cross-country studies discussed above, we use individual level data on infant mortality rather than working with 5-year country averages. These data allow us to control for the changing composition of women giving birth, and to assess how aggregate income shocks interact with a variety of characteristics of mothers and children, such as mother's education and the gender of the child.

The main finding of this paper is that there is a robust relationship between per capita GDP and infant mortality: on average, a one percent decrease in per capita GDP results in an increase in infant mortality of between 0.24 and 0.40 per 1,000 children born. Changes in infant mortality during economic downturns cannot be explained by the changing composition of women giving birth. The paper also shows that there is important heterogeneity underlying these aggregate results. The mortality of girls is significantly more sensitive to aggregate economic shocks than that of boys. This difference is particularly apparent during economic contractions, especially when these are large. This heterogeneity has important implications for the design of policies to protect children during economic downturns.

The next section, Section 2, describes the data for our sample of countries and provides details on the construction of the variables used in the analysis, in particular our measure of infant mortality. In Section 3, we discuss the basic estimation approach and present results. Section 4 concludes.

## 2. Data and construction of variables

The data on per capita GDP used for this paper are taken from World Bank (2007). The values correspond to real per capita GDP in 2000 US dollars, adjusted for differences across countries in purchasing power parity (PPP). The data on births and deaths are based on 123 Demographic and Health Surveys (DHS) covering 59 countries. The surveys include countries in Africa (33 countries, 68 surveys), Latin America (12 countries, 31 surveys), and Asia (14 countries, 27 surveys). The earliest surveys in our sample were carried out in 1986, the latest ones in 2004. Taken together, the surveys we use collect information on approximately 760,000 women and 1.7 million births. However, the sample sizes vary considerably—for example, the 1999 India DHS covers approximately 90,000 births, while the sample size for the 1987 DHS for Trinidad and Tobago is just over 3,800 births. The list of specific surveys and their sample sizes are given in Table 1.

The DHS ask women a set of questions about the date of birth, current vital statistics, and date of death (if deceased) of all children ever born. We use the responses to these questions to construct retrospective birth and death histories, closely following Paxson and Schady (2005). Our measure of infant mortality is an indicator that takes on the value of one if a child died at a reported age of 12 months or younger.[1] We discard information for children born within 12 months of the survey when calculating mortality rates to avoid complications with censored data.

Although the DHS are a rich source of data, they also have some limitations for our analysis. We briefly discuss two of these limitations, both of which are related to the use of *retrospective* information in the DHS to construct birth and death histories. First, recall bias may be a concern if women are less likely to accurately remember more distant births and deaths. To minimize recall errors, we do not use information on births that occurred more than 11 years prior to the date of the survey. Thus, our birth data cover the period 1975-2003.[2]

Second, any given survey is representative of women ages 15-49 at the time of the survey, but is not representative of all births and child deaths in earlier years. To see this, note that a woman aged 49 in a survey carried out in 2000 would have been 39 in 1990. If no surveys were carried out between 1990 and 2000 in this country, no data would be available on births to women aged 40 or older in 1990.

[1] We use this measure of infant mortality, rather than the standard definition of mortality for children *younger than* 12 months, because of age heaping in reports of mortality.
[2] The results reported in this paper are very similar when a 5-year recall is used instead of the 10-year recall. When the recall period is 15 years rather than 10 years, our estimates of the impact of GDP shocks on infant mortality fall by about one-third, but the level of statistical significance is not affected; these results available from the authors upon request.

Children born to older women may respond to economic fluctuations differently than those born to younger women. To avoid this problem, we discard from the sample births to women age 40 or older. Our analysis therefore provides meaningful estimates of the relationship between income fluctuations and infant mortality for women aged 15 to 39; we note, however, that only 1.2 percent of births in our sample of DHS countries occur to women age 40 or older. This retrospective construction of births and infant deaths to women aged 15-39 results in series of varying lengths and with varying start periods depending on the number and dates of DHS surveys in each country.[3]

The DHS collect a great deal of *current* information on mothers (for example, their education levels, whether they are employed) and children (for example, the gender and birth order and, in the most recent surveys, height and weight). Some DHS also ask respondents about their use of health services—for example prenatal check-ups and the place of delivery. However, these data are not collected in a comparable fashion in every survey, and typically are available only for the last birth. The degree to which we can analyze possible transmission mechanisms from income to infant mortality with our data is therefore limited.

## 3. Econometric specification and results

### I. *Basic results*

To estimate the effect of per capita GDP on infant mortality in our data we pool all surveys and run regressions of the following form:

[3] In the Working Paper version of this paper (Baird, Friedman, and Schady 2008) we show that our estimates of aggregate infant mortality are internally consistent and are highly correlated with other sources of data that have been used to assess the relationship between per capita GDP and infant mortality, in particular data from the World Development Indicators (WDI) data base (World Bank 2007). However, the estimates of infant mortality we calculate are more useful to estimate the relationship between fluctuations in per capita GDP and infant mortality for a variety of reasons. First, we have constructed annual series of infant mortality to look at higher-frequency changes than what can be observed in the 5-year averages in the WDI series. Data like those in WDI will have smoothed some of the year-on-year variation in infant mortality in the DHS. A share of the variation that is smoothed is likely to be measurement error, but the remainder likely reflects genuine annual fluctuations in infant mortality. Second, the data in WDI would not allow us to adjust for the changing composition of women giving birth during economic expansions or contractions, or to estimate the heterogeneity of responses to economic fluctuations by characteristics of the mother and child.

(1) $$D_{imct} = \alpha_c + \beta \log GDP_{ct} + f_c(t) + \varepsilon_{imct}.$$

where $D_{imct}$ is an indicator variable that takes on the value of one if child $i$ born to mother $m$ in country $c$ in year $t$ died in the first year of life, zero otherwise; $\alpha_c$ is a set of country fixed effects; $\log GDP_{ct}$ is the natural logarithm of per capita GDP; $f_c(t)$ is a flexible, country-specific formulation of time (in practice, we present results that include linear, quadratic, and cubic terms); and $\varepsilon_{imct}$ is the error term. Standard errors are clustered at the country level in order to correct for autocorrelation of arbitrary form in shocks to infant mortality across years within country. In this specification, $\beta$ is the impact of GDP on infant mortality, after removing country-specific trends from the data.[4]

In principle, there are two mechanisms that could account for a negative association between infant mortality and aggregate economic circumstances. First, it is possible that a child born to a woman of given characteristics is more likely to die if economic circumstances are unfavorable. Second, it is possible that the composition of women giving birth changes with economic circumstances. Clearly, these two causes for a possible countercyclical relationship between GDP and infant mortality—changes in mortality risk for a child born to a given woman or changes in the pool of women giving birth—have very different implications.

An obvious way to adjust for compositional changes is to include the characteristics of women, children, and births in equation (1), which gives us the following:

(2) $$D^{adj}_{imct} = \alpha_c + \beta \log GDP_{ct} + f_c(t) + \delta X_{im} + \varepsilon_{imct}$$

where $X_{im}$ is a vector of characteristics of child $i$ born to mother $m$. Recall that child births and deaths are calculated on the basis of retrospective questions asked of mothers at the time of the survey, which limits the variables that can be included in (2). In practice, we control for a cubic term for mother's years of education, maternal age at the time of birth, and birth order; and binary indicators for place of residence

[4] We obtain very similar results from a two-step process in which we first collapse the data to the level of the country-year cell, and then account for secular trends in various ways—including regressions in first differences, with a formal Error Correction Model (ECM), and smoothing the data with standard time-series filters such as the Hodrick-Prescott and Baxter-King filters (see Baird, Friedman, and Schady 2008 for a discussion of these estimates).

(urban or rural) at the time of the survey, the gender of the child, and whether or not the child was a multiple birth. All of these variables are highly correlated with the probability of child survival.[5] This approach implicitly assumes that place of residence at the time of the survey is correlated with place of residence at the time of child birth, and that schooling has been completed by age 15; these should be reasonable approximations for most of the countries and years in our sample. In addition, as an alternative means of controlling for compositional effects, we include a set of mother fixed effects, as well as birth-specific characteristics (child gender and an indicator for multiple births). This approach has the advantage that it controls for all time-invariant mother characteristics, not just education and place of residence, but limits the sample to women who have had at least two live births.

Our main set of results is presented in Table 2. The first row in Table 2, which reports the results from estimates of Equation (1), implies that a one percent decrease in per capita GDP is associated with a 0.24 to 0.40 increase in infant mortality per thousand children born. On average, the country-specific year-on-year decrease in infant mortality in our data is 2.5 per 1000 live births. A one percent decline in per capita GDP from expected trends therefore results in an increase in infant mortality of between 10 and 15 percent of the average annual mortality decline in our data. Note also that those regressions which more flexibly account for underlying secular trends result in larger (in absolute value) estimates of the association between per capita GDP and infant mortality. Previous studies have generally adjusted only for linear trends (as in Jamison et al. 2004) and hence may underestimate the contemporaneous relationship between detrended GDP and infant mortality.

Results from regressions that include the vector of covariates $X_{im}$ (Equation (2)) are presented in the second row of Table 2. These results show that including these covariates has a negligible effect on estimates of the association between log per capita GDP and infant mortality. The third row of the table reports the results from regressions that do not include mother characteristics or fixed effects for the sample of women that have had at least two live births. These results are presented to put the fixed effects

[5] There is an extensive literature on this topic. See, for example, the review papers by Behrman and Deolalikar (1988); Strauss and Thomas (1998); and Schultz (2002).

estimates in context; they show that the association between GDP and mortality in this smaller sample is very similar to that observed in the full sample of live births. The fourth row of Table 2, finally, reports the results that include the mother fixed effects and birth-specific characteristics. These coefficients are very similar to those without fixed effects. In sum, Table 2 makes clear that the changing composition of women cannot account for the bulk of the association between infant mortality and log per capita GDP that we observe in our data. Rather, when there are negative economic shocks, there is an increase in mortality risk for infants born to a given mother.

### *II. Timing of shocks to GDP*

The discussion so far has focused on the contemporaneous relationship between GDP and infant death, without giving explicit attention to the timing of shocks. As a first step to clarifying this issue, we include terms in lagged and lead per capita GDP in our basic regression.[6] The top panel in Table 3 shows that the coefficients on both of these terms are small, and are not significant at conventional levels. Only the coefficient on current GDP in the top panel of Table 3 is significant. This suggests that it is not the economic conditions early in the pregnancy that are most important in determining infant mortality—these conditions are loaded on lagged GDP for most children, and the coefficient on lagged GDP is insignificant. Similarly, it does not appear to be that conditions in the later part of a child's first year in life substantially affect the probability of survival—these conditions are loaded on to lead GDP for most children in our sample, and the coefficient on lead GDP is also insignificant. Rather, it appears that

[6] To see how this speaks to the issue of the effects of shocks to GDP at different times in an infant's life, it is useful to work out what the coefficients on lagged, current, and lead GDP imply for children born at different times in the year. For a child born early in the year (say, in January), the coefficient on lagged GDP mainly reflects conditions before conception and in utero, the coefficient on current GDP reflects conditions in the first year of life, and the coefficient on lead GDP reflects conditions in the second year—beyond the period that is relevant for the calculation of infant mortality. By contrast, for a child born late in the year (say, in December), the coefficient on lagged GDP reflects conditions before conception, the coefficient on current GDP reflects conditions in utero, while the coefficient on lead GDP reflects the conditions after birth. Finally, for a child born at the midpoint of the year, on June 30, lagged GDP reflects conditions before conception and during the first three months in utero, current GDP reflects conditions in the last six months in utero and the first six months after birth, while lead GDP reflects conditions after the infant is six months of age.

economic conditions in those months shortly before and shortly after birth have the biggest effect on the probability that a child survives.

We make a further attempt to clarify issues about the window of vulnerability that infants face with regard to GDP shocks. Mothers report the year and month of birth of each child, and we assign the 15th day of the relevant month as the birth date for each child. Using these data, we then construct birth-month specific exposure windows for economic conditions in utero, in the first month of life, and in the next eleven months. The results from these regressions are presented in the lower panel of Table 3. The coefficients on economic conditions in utero and after the first month of life are both small and insignificant. By contrast, the coefficient on per capita GDP in the first month is large, significant, and very close in magnitude to that reported in Table 2.[7] These results underscore that economic conditions around birth appear to matter most for infant survival.

The importance of economic conditions "around" birth for infant survival also yields some clues about the likely transmission mechanisms from trend deviations in per capita GDP to infant mortality. Low birthweight is considered an important risk factor in predicting neonatal and infant death (see for example the review by Lawn et al. 2005). However, the fact that the coefficient on economic conditions for much of the in utero period is not significant in Table 3 suggests that this is unlikely to be the main reason for elevated infant mortality during economic downturns.[8] On the other hand, skilled attendance during birth, or health care for children who face infections shortly thereafter, may help explain our findings. Approximately 36 percent of neonatal deaths worldwide are a result of severe infections during birth, and another 23 percent are a result of asphyxia (Lawn et al. 2005). Poor economic conditions around birth could result either in a deterioration in public health services or in a decrease in households'

[7] We also experimented with breakdowns of the in utero period. For example, in a study of the effect of the Chernobyl nuclear disaster, Almond et al. (2007) show that radiation exposure was particularly damaging during the period between 8 and 25 weeks after conception. The emphasis of our paper is on economic conditions, rather than radiation exposure, but it is conceivable that the period of 8 to 25 weeks post-conception is one in which health insults more generally are particularly damaging. However, in none of the specifications we ran was the coefficient on economic conditions in the period corresponding to 8-25 weeks after conception significant once we controlled for conditions in the last three months of pregnancy and after birth.

ability to pay for, or otherwise access, skilled birth assistance (as suggested for example by Paxson and Schady 2005 in their analysis of infant mortality in Peru), both of which could lead to increased mortality in the first year of life. It is also possible that maternal mortality around the time of birth plays a role: Children whose mothers die are themselves much more likely to die (for example, Anderson et al. 2007), and death in childbirth may increase during poor economic times.[9]

## *III. Heterogeneity*

Up to this point, we have implicitly assumed that aggregate income shocks affect all mothers and children equally. Yet, there are a host of reasons why this need not be so. For example, more educated and wealthier mothers may be better able to smooth consumption of critical inputs into child health; there may also be within-household discrimination so that boys are better protected from negative health shocks than girls; families in richer countries may have greater access to credit markets, and may be better able to smooth consumption of essential items. We now turn to the question of heterogeneity of impacts, focusing on differences by the gender of a child, the education and age of the mother, place of residence (urban or rural), birth parity, and the overall income of country of residence.

To motivate our results, we first present the mean infant mortality rates in Table 4 for each mother, child, or country characteristic we use in our analysis. The first row of the table shows that girls are almost ten percentage points less likely to die in the first year of life than boys—a well-known finding in the demographic literature.[10] The other coefficients show that children born in rural areas are more likely to die than those born in urban areas; that the mortality of children born to mothers with less than primary schooling is more than twice as high as that of children born to mothers with completed primary

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[8] Selection may be important if poor economic conditions in utero lead to a higher rate of spontaneous abortions. The sample of children born alive may then have higher health endowments, introducing a downward bias to the association we estimate between economic conditions in utero and infant mortality.

[9] We thank an anonymous referee for this suggestion. Our birth data is reported retrospectively by mothers alive at time of survey, and so we do not observe maternal mortality in our data. As a result, if maternal mortality is countercyclical our estimates could be biased downwards.

[10] For example, the World Health Organization (2006) estimates that the male-to-female ratio in neonatal mortality and in early neonatal mortality in developing countries is 1.3.

schooling or more; that children born to young mothers (age 15-19) and older mothers (age 35-39) are more likely to die than those born to "prime-age" mothers (age 20-34); that high-parity births (5$^{th}$ birth or higher) are also more likely to die than lower parity births; and that children born in lower income developing countries are more likely to die than children born in middle income countries.[11]

We next analyze heterogeneity in the relationship between detrended per capita GDP and infant mortality along these observable dimensions of mothers and children. Our approach is straightforward. In each case, we generate an indicator for the characteristic in question—for example, an indicator for the birth of a girl—and then interact this indicator with the measure of log per capita GDP. Table 5 then reports the coefficients on the main effect for log per capita GDP and on the interaction between log per capita GDP and the given characteristic. We focus on the specification that includes country-specific cubic time trends, as these account for underlying time trends most flexibly.

Table 5 shows that the mortality of infants born to mothers with low education levels is significantly more sensitive to changes in economic conditions than that of children born to mothers with higher education levels. In part, this is the result of the higher mortality rates among women with lower education levels, although this does not fully explain the differences in the magnitudes we estimate.[12] A similar pattern can be seen in a comparison between rural and urban areas—the increase in mortality during economic downturns is larger in rural areas, but from a higher base. Also, the point estimate on the interaction term for middle income countries suggests larger increases in infant mortality during economic downturns in low income countries, although this difference is not significant at standard levels. The most striking result in the table relates to differences by gender. Although the average mortality among boys is higher than among girls, Table 5 shows that the mortality of girls is much more

[11] The GDP data we use in this paper are measured in constant 2000 US dollars. The World Bank (2001) classifies countries as "low income" if per capita GDP in constant 2000 dollars is below $755. To classify countries as "low" or "middle" income, we apply the World Bank threshold to the 1980 per capita GDP data. Using 1980, which corresponds to a year before the beginning of our infant mortality series for the bulk of the countries we analyze, limits the potential for possible simultaneity biases induced by feedback from health to income.

[12] A one log-unit decrease in per capita GDP would increase the infant mortality rate of children born to low education women from 104 to 160.7, and that of children born to high education women from 51.7 to 69.5. The proportional, not just the absolute, change among low-education mothers is thus larger.

sensitive to changes in economic circumstances than that of boys: A one percent change in per capita GDP changes the mortality of boys by approximately 0.27 per thousand children born, and that of girls by 0.53 per thousand—a remarkable difference by any standard.[13]

## *IV. Magnitude and sign of shocks to per capita GDP*

In addition to heterogeneity by household characteristics, there may also be heterogeneity in impacts by the sign (positive or negative) and magnitude of the GDP shock. To investigate this, we estimate a series of gender-specific, continuous spline regressions. [14] The results from these estimations are reported in Table 6, separately for boys (upper panel) and girls (lower panel). The first (top) row for each panel presents the results from a spline regression with a knot at zero, which allows for different slopes for positive and negative changes in GDP. We then turn to spline regressions with two knots. In the second row these knots are fixed at -1σ and 1σ (where σ stands for standard deviations of GDP trend deviations in our sample), in the third row they are fixed at -1.5σ and 1.5σ, and in the fourth (bottom) row the knots are fixed at -2σ and 2σ.

Table 6 shows that positive shocks to per capita GDP affect girls and boys in a similar fashion. On the other hand, negative shocks have much larger effects on the mortality of girls than boys. For

[13] We also considered gender differences in the impact of shocks to GDP on infant mortality separately by region. These results suggest that the mortality of girls is more sensitive to GDP shocks than that of boys in every region in the developing world for which we have data. Thus, in Sub-Saharan Africa, a one percent decrease in GDP increases the mortality of boys by 0.33 per thousand, and that of girls by 0.62 per thousand; comparable figures for Latin America and the Caribbean are 0.29 per thousand (boys) and 0.46 per thousand (girls); for Southeast Asia, these are 0.15 per thousand (boys) and 0.24 per thousand (girls); for South Asia, 0.72 per thousand (boys) and 1.43 per thousand (girls); and for the Middle East and North Africa region, a one percent decline in GDP results in a *decrease* of mortality of 0.18 per thousand boys born, and an increase of mortality of 0.78 per thousand girls born. All of the coefficients on GDP in the regressions for girls are significant at the 10 percent level or higher; in the regressions for boys, only the coefficients in the regressions for Southeast Asia and Latin America and the Caribbean are significant. We conclude that the gender differences in the effect of GDP shocks on infant mortality we observe in our sample of developing countries as a whole is not driven by a single region, including regions where a preference for boys has been well documented (for example, South Asia).

[14] For this purpose, we regress the indicator for infant death on country fixed effects and a country-specific cubic polynomial in time and predict the residual from this regression. We also regress log per capita GDP on country fixed effects and a country-specific cubic polynomial in time, and predict the residuals from this regression. Finally, we use locally weighted least squares to depict the relation between the residual from the infant death regression and the residual from the GDP regression. For presentational purposes, the figure is trimmed at the 1st and 99th percentiles of GDP deviation from cubic trend.

example, for negative shocks of -1.5σ or larger, a one percent decrease in log per capita GDP results in an increase in mortality of 1.05 per thousand girls born (with a standard error of 0.30), but an increase in mortality of only 0.53 per thousand boys born (with a standard error of 0.23). On average, countries with a negative shock to per capita GDP of 1.5σ or larger had a contraction of 5.9 percent. (There are 122 such country-year events in our data.) The average increase in girl infant mortality during these negative shocks to aggregate income is 7.4 deaths per 1000, approximately three times the magnitude of the average country-specific annual reduction in mortality. These simple back of the envelope calculations suggest that the magnitude of the effects of large negative income shocks on infant mortality, in particular of girls, is large by any standard.

In sum, Table 6 is consistent either with girls being more fragile in their first year of life than boys, which seems unlikely, or with families protecting boys more than girls during economic downturns. Put differently, these results suggest that household behavioral responses to negative shocks play an important role in determining infant survival. Finally, Table 6 underscores that our results are unlikely to be driven by omitted variables, as any potential omitted variables would have to interact with both the gender of the child and the direction of the income shock. It is hard to imagine what such an omitted variable would be.[15,16]

## 4. Conclusion

Macroeconomic volatility is a fact for most developing countries. In recent decades, the standard deviation of income over time has been approximately twice as large in developing as developed countries (Aguiar and Gopinath 2007). A recent review stresses the welfare costs of volatility for

[15] In the Working Paper version of our paper (Baird, Friedman and Schady 2008) we also show that our results are insensitive to the inclusion of controls for a number of possible omitted variables such as rainfall; conflict, including civil war; and inflation and other measures of the quality of governance. This further suggests that our results are not driven by the omitted variables that have received the most attention in the literature.

[16] The observed asymmetry by gender also makes it very unlikely that our results are driven by recall bias in the DHS. For example, in principle one might be concerned that mothers use a salient event like an economic crisis to date an infant death, which could induce a spurious association between negative income shocks and mortality. It seems unlikely, however, that this sort of recall bias would be present with female deaths, but not male deaths.

developing countries in terms of their inability to smooth consumption (Loayza et al. 2007). In this paper we document another way in which aggregate economic fluctuations can have dramatic welfare consequences. In developing countries, infants, in particular girls, are more likely to die when there is a negative economic shock.

Aggregate macroeconomic shocks involve both income and substitution effects for individual households. Given positive income gradients in child health, the income effect would generally result in an increase in mortality. But there is also a substitution effect, as economic shocks decrease the opportunity cost of time and may free up mothers for time-intensive tasks that have positive effects on child health—for example, collection of clean water, preparation of food, or regular visits to health centers. The effect of aggregate economic contractions on child health is therefore hard to sign ex ante. The literature on the United States suggests that child health generally improves, and infant mortality declines during economic contractions.

Our results suggest that economic shocks in the developing world generally lead to more infant deaths, especially of girls, and especially when these shocks are severe. Of course, there is variation across countries—for example, Ferreira and Schady (2009) argue that the much larger increase in infant mortality in Peru during the crisis of the late 1980s than was the case in Indonesia during the crisis of the late 1990s may have been a result, in part, of the protection of health expenditures in Indonesia (but not in Peru). We cannot systematically explore these differences with the data at hand. Nevertheless, our results make clear that the findings from a handful of country-specific studies, including Cutler et al. (2002) on Mexico, Paxson and Schady (2005) on Peru, and Bhalotra (2008) on India, hold for a much larger sample of developing countries and time periods. We also show that the effects of crises on infant mortality appear to be much more severe for girls than boys.

We conclude by discussing two areas where our data impose limitations on the possible analysis we can conduct. The first of these is the timing of the GDP shocks. Our results suggests that it is macroeconomic conditions around birth, rather than in the early in utero period or in the later half of a child's first year of life, which matter most for a child's survival in her first year. However, with annual data like those we use it is not possible to tease out the relative importance of conditions in narrow windows of exposure. Our results on the timing of the shock to aggregate income should therefore be viewed as suggestive rather than definitive. Second, because we construct birth and death histories retrospectively, we do not have data on the utilization of health services before, during and after birth for the majority of births (and deaths) we observe. Further, the DHS data we use do not include information on other potential inputs into child health, such as the consumption of nutritious foods. We are therefore unable to explore in a comprehensive manner the transmission mechanisms from income shocks to infant mortality. Nevertheless, our results clearly indicate that short-term fluctuations in aggregate income can have important consequences for the likelihood that a child survives her first year of life. Policies that reduce the volatility of per capita GDP in developing countries, or that protect health status during sudden economic downturns, may have significant benefits for child survival, especially that of girls.

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Table 1. List of DHS datasets used in the analysis, including information on country, year of survey, number of mothers, and number of births

| Country | Survey Years | Total Mothers | Total Births |
|---|---|---|---|
| Armenia | 2000 | 2446 | 4234 |
| Bangladesh | 1994, 1997, 1999, 2004 | 26313 | 51071 |
| Benin | 1996, 2001 | 7515 | 18891 |
| Bolivia | 1989, 1994, 1998,2004 | 24574 | 54474 |
| Brazil | 1986, 1992, 1996 | 11672 | 23590 |
| Burkina Faso | 1993, 1999, 2003 | 16362 | 39410 |
| Burundi | 1987 | 2416 | 6464 |
| Central African Republic | 1995 | 3373 | 7962 |
| Cameroon | 1991, 1998, 2004 | 11444 | 27350 |
| Chad | 1997 | 4655 | 11829 |
| Colombia | 1986, 1990, 1995, 2000 | 17149 | 31010 |
| Comoros | 1996 | 1405 | 3838 |
| Cote d'Ivoire | 1994, 1999 | 6660 | 15993 |
| Dominican Republic | 1986, 1991, 1996, 1999, 2002 | 23486 | 48458 |
| Ecuador | 1987 | 2536 | 5817 |
| Egypt | 1993, 1996, 2000, 2003 | 33988 | 73378 |
| Ethiopia | 2000 | 8436 | 20484 |
| Gabon | 2001 | 3371 | 7084 |
| Ghana | 1988, 1994, 1999, 2003 | 11841 | 25675 |
| Guatemala | 1987, 1995, 1999 | 13496 | 33832 |
| Guinea | 1999 | 4549 | 11224 |
| Haiti | 1995, 2000 | 7764 | 18283 |
| India | 1993, 1999 | 103669 | 208690 |
| Indonesia | 1987, 1991, 1994, 1997, 2003 | 81673 | 153661 |
| Kazakhstan | 1995, 1999 | 3971 | 6624 |
| Kenya | 1989, 1993, 1998, 2003 | 18457 | 44289 |
| Kyrgyz Republic | 1997 | 2131 | 4100 |
| Liberia | 1986 | 3419 | 8669 |
| Madagascar | 1992, 1997 | 7592 | 19195 |
| Malawi | 1992, 2000 | 11368 | 27292 |
| Mali | 1987, 1996, 2001 | 17915 | 47710 |
| Mexico | 1987 | 4528 | 10177 |
| Morocco | 1987, 1992, 2004 | 14775 | 33052 |
| Mozambique | 1997 | 5535 | 12468 |
| Namibia | 1992, 2000 | 6674 | 13550 |
| Nepal | 1996, 2001 | 12058 | 27569 |
| Nicaragua | 1998, 2001 | 14098 | 29598 |
| Niger | 1992, 1998 | 9468 | 26714 |
| Nigeria | 1990, 1999, 2003 | 14333 | 36543 |
| Pakistan | 1991 | 4874 | 13255 |
| Paraguay | 1990 | 3208 | 7752 |
| Peru | 1986, 1992, 1996, 2000 | 40330 | 84225 |
| Philippines | 1993, 1998, 2003 | 20621 | 46551 |
| Rwanda | 1992, 2000 | 9317 | 23607 |
| Senegal | 1986, 1993, 1997 | 11881 | 30636 |
| South Africa | 1998 | 6017 | 9970 |

Table 1 (continued)

| | | | |
|---|---|---|---|
| Sri Lanka | 1987 | 4121 | 8250 |
| Sudan | 1990 | 4242 | 11314 |
| Tanzania | 1992, 1996, 1999 | 12826 | 29743 |
| Thailand | 1987 | 4294 | 7516 |
| Togo | 1988, 1998 | 7611 | 18582 |
| Trinidad and Tobago | 1987 | 1786 | 3588 |
| Tunisia | 1988 | 3224 | 8318 |
| Turkey | 1993, 1998 | 7897 | 15306 |
| Uganda | 1989, 1995, 2001 | 11883 | 30062 |
| Uzbekistan | 1996 | 2315 | 4744 |
| Vietnam | 1997, 2002 | 7643 | 13012 |
| Zambia | 1992, 1997, 2002 | 13776 | 32044 |
| Zimbabwe | 1989, 1994, 1999 | 9346 | 19913 |
| Total: | | 764327 | 1668640 |

Table 2. Income shocks and infant mortality

| Dependent variable | Linear | Quadratic | Cubic |
|---|---|---|---|
| | | *Unadjusted* | |
| IMR | -23.96 | -32.88 | -39.81 |
| | [8.11]*** | [7.40]*** | [9.84]*** |
| | | *Controlling for mother and birth characteristics* | |
| IMR | -23.46 | -30.78 | -37.83 |
| | [7.73]*** | [6.99]*** | [9.82]*** |
| | | *Unadjusted, restricted to mothers with multiple births* | |
| IMR | -26.34 | -31.08 | -38.25 |
| | [9.08]*** | [7.59]*** | [11.60]*** |
| | | *Mothers' fixed effects* | |
| IMR | -29.46 | -32.33 | -36.22 |
| | [9.43]*** | [8.69]*** | [11.45]*** |

Number of observed births equals 1,634,360 in first two panels, 1,356,738 in bottom two panels. Mother and birth characteristics include indicators for rural location, gender of child, and multiple birth, and cubic terms for mothers' age, years of education, and infant birth order. Robust standard errors are clustered at the country level - there are 59 countries. GDP is measured in year 2000 international (PPP) dollars.
* p<.10, ** p<.05, *** p<.01

Table 3. GDP shocks and infant mortality, including possible lead and lag effects

| Independent variable | Linear | Quadratic | Cubic |
|---|---|---|---|
| | *Lagged, Current, and Lead GDP* | | |
| Lagged GDP | -1.08 | -6.66 | -5.45 |
| | [10.93] | [11.19] | [9.85] |
| GDP | -31.26 | -36.59 | -38.74 |
| | [11.59]*** | [10.43]*** | [11.20]*** |
| Lead GDP | 10.93 | 6.69 | 6.19 |
| | [7.26] | [8.59] | [8.48] |
| | *GDP Series Reweighted to Approximate Exposure over Course of In Utero Development and First Year* | | |
| In Utero | 7.5 | 2.37 | 3.93 |
| | [19.86] | [17.58] | [14.61] |
| First Month | -38.71 | -40.61 | -41.77 |
| | [20.59]* | [20.86]* | [18.97]** |
| Next 11 Months | 9.84 | 3.74 | 1.54 |
| | [10.23] | [11.01] | [12.73] |

Robust standard errors are clustered at the country level - there are 1,549,745 observations distributed across 840 country-year cells and 59 countries. GDP is measured in year 2000 international (PPP) dollars.
* p<.10, ** p<.05, *** p<.01

Table 4. Infant mortality by mother, child, and country characteristics

| Characteristic | Estimated IMR | | |
|---|---|---|---|
| Child gender | Boys | | Girls |
| | 88.4 | | 78.5 |
| N | 833,545 | | 800,814 |
| Mothers' education | Less than primary | | Primary or greater |
| | 99.5 | | 51.3 |
| N | 1,093,757 | | 540,603 |
| Mother location | Urban | | Rural |
| | 61.4 | | 94.9 |
| N | 555,742 | | 1,078,618 |
| Mothers' age | 15-19 | 20-34 | 35-39 |
| | 105.4 | 77.0 | 90.1 |
| N | 296,461 | 1,151,038 | 144,052 |
| Birth order | 1st | 2nd-4th | 5th+ |
| | 79.1 | 75.6 | 101.5 |
| N | 381,176 | 804,593 | 448,591 |
| Country income | Low-Income | | Middle-Income |
| | 94.5 | | 67.7 |
| N | 964,446 | | 669,914 |

Table 5. Heterogeneity in IMR and GDP relation by mother, birth, or country characteristic, cubic trend

| Characteristic | GDP | Interaction (GDP*characteristic) |
|---|---|---|
| Female infant | -27.22 | -25.52 |
| | [10.40]** | [10.20]** |
| Low mother's education | -31.31 | -12.32 |
| | [11.98]** | [14.00] |
| Rural location | -21.28 | -26.33 |
| | [11.81]* | [10.45]** |
| Young mother (<20) | -44.20 | 22.52 |
| | [10.15]*** | [15.86] |
| Older mother (>34) | -36.05 | -19.37 |
| | [8.24]*** | [18.51] |
| First births | -49.24 | 42.68 |
| | [10.81]*** | [16.26]** |
| High birth order (>4) | -30.42 | -29.78 |
| | [9.55]*** | [14.54]** |
| Middle income country | -46.14 | 14.14 |
| | [18.49]** | [23.32] |

Low mother's education is defined as less than primary attainment. Robust standard errors are clustered at the country level. GDP is measured in year 2000 international (PPP) dollars. In this currency measure, the World Bank threshold for middle-income country status is a per-capita GNI of $755.
* p<.10, ** p<.05, *** p<.01

Table 6. Heterogeneity in IMR and GDP relation by size of GDP deviation from cubic trend, separate for male and female children

| Dependent variable | Magnitude and direction of GDP deviation | | |
|---|---|---|---|
| IMR for Boys (N=833,545) | *<= 0* | | *> 0* |
| | -20.75 | | -31.00 |
| | [10.10]** | | [16.46]* |
| | *<= -1 sd* | *> -1 sd & <= 1 sd* | *> 1 sd* |
| | -38.22 | -14.73 | -40.69 |
| | [19.34]* | [21.44] | [25.10] |
| | *<= -1.5 sd* | *> -1.5 sd & <= 1.5 sd* | *> 1.5 sd* |
| | -52.81 | -13.16 | -67.25 |
| | [22.82]** | [15.48] | [19.67]*** |
| | *<= -2 sd* | *> -2 sd & <= 2 sd* | *> 2 sd* |
| | -59.09 | -18.07 | -70.82 |
| | [26.47]** | [12.64] | [19.65]*** |

| Dependent variable | Magnitude and direction of GDP deviation | | |
|---|---|---|---|
| IMR for Girls (N=800,814) | *<= 0* | | *> 0* |
| | -55.43 | | -43.71 |
| | [13.90]*** | | [14.05]*** |
| | *<= -1 sd* | *> -1 sd & <= 1 sd* | *> 1 sd* |
| | -75.81 | -39.24 | -47.27 |
| | [24.41]*** | [22.00]* | [21.84]** |
| | *<= -1.5 sd* | *> -1.5 sd & <= 1.5 sd* | *> 1.5 sd* |
| | -104.71 | -36.60 | -58.66 |
| | [30.03]*** | [15.97]** | [19.93]*** |
| | *<= -2 sd* | *> -2 sd & <= 2 sd* | *> 2 sd* |
| | -148.47 | -36.52 | -69.12 |
| | [51.38]*** | [14.04]** | [19.72]*** |

Note: Slope coefficients are estimated from a continuous spline specification. Robust standard errors are clustered at the country level. GDP is measured in year 2000 international (PPP) dollars.
* p<.10, ** p<.05, *** p<.01