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AGGREGATE INCOME SHOCKS AND INFANT MORTALITY IN THE DEVELOPING WORLD

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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 investigate whether short-term fluctuations in aggregate income affect infant mortality using an unusually large data set of 1.7 million births in 59 developing countries. We show a large, negative association between per capita GDP and infant mortality. Female infant mortality is more sensitive than male infant mortality to negative economic shocks, suggesting that policies that protect the health status of female infants may be especially important during economic downturns.

I. 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% lower than that of people in the highest ventile in 1980 (Rogot et al., 1992). 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, Sandbu, and Wang (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 independent of income (Cutler, Deaton, & 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% of all deaths occur to children under the age of five, compared to less than 1% in rich countries (see Cutler et al., 2006). 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 caused primarily by improvements in medical technology rather than directly 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 reduce household consumption of nutritious foods and lower expenditures on other inputs into child health, and they may seriously disrupt public health services; all of these would tend to increase infant mortality. Aggregate shocks, however, depress wages and 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, and collecting clean water. Because systemic 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 concerns 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 procyclical 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 & 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 & 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 & Schady, 2009). Sharp economic downturns have been associated with increases in infant mortality in Mexico (Cutler et al., 2002), Peru (Paxson & Schady, 2005), and India (Bhalotra, 2010). Miller and Urdinola (2010), however, find that arguably exogenous declines in the price of coffee, which resulted in declines in aggregate income in coffee-growing

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areas in Colombia, were associated with lower infant mortality, echoing the results from the United States.

Our paper extends the existing literature in a number of important ways. The sample, 59 countries, 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 already discussed, we use individual-level data on infant mortality rather than working with five-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 shocks to per capita GDP and infant mortality: on average, a 1% 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 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.

Section II, 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 III, we discuss the basic estimation approach and present results. Section IV concludes.

II. 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 U.S. 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 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 data we use contain 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 is 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 1 if a child died at a reported age of 12 months or younger.[1] We discard information for children born within twelve 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 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 eleven years prior to the date of the survey. Thus, our birth data cover the period 1975 to 2003.[2]

Second, any given survey is representative of women ages 15 to 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 the country of study, no data would be available on births to women aged 40 or older in 1990. Children born to older women may respond to economic fluctuations differently from those born to younger women. To avoid the problem of varying age composition of birth mothers, we discard from the sample births to women age 40 or older. Our analysis therefore provides meaningful estimates of the relationship between income shocks and mortality of infants born to women aged 15 to 39. We note, however, that only 1.2% of births in our sample of DHS countries occur to women age 40 or older in any year of analysis. This retrospective construction of births and infant deaths to women aged 15 to 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]

[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 five-year recall is used instead of the ten-year recall. When the recall period is fifteen years rather than ten years, our estimates of the impact of GDP shocks on infant mortality fall by about one-third, but are still precisely estimated. These results are available from the authors on request.

[3] In the working paper version of this paper (Baird, Friedman, & Schady, 2007), 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) database (World Bank, 2007). However, the estimate of infant mortality we calculate is more useful to estimate the relationship between shocks to 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 five-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.

TABLE 1.—DHS DATA SETS USED IN THE ANALYSIS

Country	Survey Years	Total Number of Mothers	Total Number of Births
Armenia	2000	2,446	4,234
Bangladesh	1994, 1997, 1999, 2004	26,313	51,071
Benin	1996, 2001	7,515	18,891
Bolivia	1989, 1994, 1998, 2004	24,574	54,474
Brazil	1986, 1992, 1996	11,672	23,590
Burkina Faso	1993, 1999, 2003	16,362	39,410
Burundi	1987	2,416	6,464
Central African Republic	1995	3,373	7,962
Cameroon	1991, 1998, 2004	11,444	27,350
Chad	1997	4,655	11,829
Colombia	1986, 1990, 1995, 2000	17,149	31,010
Comoros	1996	1,405	3,838
Cote d'Ivoire	1994, 1999	6,660	15,993
Dominican Republic	1986, 1991, 1996, 1999, 2002	23,486	48,458
Ecuador	1987	2,536	5,817
Egypt	1993, 1996, 2000, 2003	33,988	73,378
Ethiopia	2000	8,436	20,484
Gabon	2001	3,371	7,084
Ghana	1988, 1994, 1999, 2003	11,841	25,675
Guatemala	1987, 1995, 1999	13,496	33,832
Guinea	1999	4,549	11,224
Haiti	1995, 2000	7,764	18,283
India	1993, 1999	103,669	208,690
Indonesia	1987, 1991, 1994, 1997, 2003	81,673	153,661
Kazakhstan	1995, 1999	3,971	6,624
Kenya	1989, 1993, 1998, 2003	18,457	44,289
Kyrgyz Republic	1997	2,131	4,100
Liberia	1986	3,419	8,669
Madagascar	1992, 1997	7,592	19,195
Malawi	1992, 2000	11,368	27,292
Mali	1987, 1996, 2001	17,915	47,710
Mexico	1987	4,528	10,177
Morocco	1987, 1992, 2004	14,775	33,052
Mozambique	1997	5,535	12,468
Namibia	1992, 2000	6,674	13,550
Nepal	1996, 2001	12,058	27,569
Nicaragua	1998, 2001	14,098	29,598
Niger	1992, 1998	9,468	26,714
Nigeria	1990, 1999, 2003	14,333	36,543
Pakistan	1991	4,874	13,255
Paraguay	1990	3,208	7,752
Peru	1986, 1992, 1996, 2000	40,330	84,225
Philippines	1993, 1998, 2003	20,621	46,551
Rwanda	1992, 2000	9,317	23,607
Senegal	1986, 1993, 1997	11,881	30,636
South Africa	1998	6,017	9,970
Sri Lanka	1987	4,121	8,250
Sudan	1990	4,242	11,314
Tanzania	1992, 1996, 1999	12,826	29,743
Thailand	1987	4,294	7,516
Togo	1988, 1998	7,611	18,582
Trinidad and Tobago	1987	1,786	3,588
Tunisia	1988	3,224	8,318
Turkey	1993, 1998	7,897	15,306
Uganda	1989, 1995, 2001	11,883	30,062
Uzbekistan	1996	2,315	4,744
Vietnam	1997, 2002	7,643	13,012
Zambia	1992, 1997, 2002	13,776	32,044
Zimbabwe	1989, 1994, 1999	9,346	19,913
Total		764,327	1,668,640

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 most recent birth. The degree to which we can analyze possible transmission mechanisms from income to infant mortality with our data is therefore limited.

III. Econometric Specification and Results

A. 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:

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

where D_{imct} is an indicator variable that takes the value of 1 if child i born to mother m in country c in year t died in the first year of life, and 0 otherwise; α_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 ε_{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 a country. In this specification, β is the impact of GDP on infant mortality, after removing country-specific trends and intercepts from the data.[4]

In principle, two mechanisms 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 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 for our understanding and for the design of corrective policy.

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

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

where X_{imc} 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 equation (2). In practice, we control for third-order polynomials in mother's years of education, maternal age at the time of birth, and birth order and binary indicators for place of residence (urban or rural) at the time of the survey, the gender of the child, and whether 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 1% decrease in per capita GDP is associated with a 0.24 to 0.40 increase in infant mortality per 1,000 children born.[6] On average, the country-specific year-on-year decrease in infant mortality in our data is 2.5 per 1,000 live births. A 1% shortfall in per capita GDP from expected trends therefore results in an increase in infant mortality of between 10% and 15% of the average annual mortality decline in our data. Note also that those regressions that 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 effect of economic shocks on 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

[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, & Schady, 2007, for a presentation of these estimates).

[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).

[6] The results in table 2, as well as in the other tables in the paper, differ slightly from those in the working paper version of our paper (Baird et al., 2007) for two reasons. First, the results in the working paper version clustered standard errors at the country-year level; in this version, we use the more conservative approach of clustering standard errors at the country level. Second, the results in the working paper version were based on weighted regressions that used the within-country weights provided in the DHS documentation. In this version, we report the results from unweighted regressions. We choose to show the unweighted regressions since the DHS weights are constructed to draw inferences that are representative of a country's population at the time of survey, but our retrospective birth histories extend to eleven years before year of survey, and it is not clear how appropriate the weights provided are to the earlier period. More important, we are not conducting a country-by-country analysis, but rather pooling data from all available surveys. The weights provided by the DHS make no adjustment for the fact that the underlying populations across countries are very different. For both of these reasons, regressions without weights are more transparent and intuitive. We note, however, that none of the main messages in the paper are affected by these changes.

TABLE 2.—INCOME SHOCKS AND INFANT MORTALITY

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

Number of observed births equals 1,634,360 in first two panels and 1,356,738 in bottom two panels. Mother and birth characteristics are 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 the 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. *$p < .10$, **$p < .05$, ***$p < .01$.

sample of women who have had at least two live births. These results are presented to place the fixed-effects 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 reports the results that include the mother fixed effects and birth-specific characteristics. These coefficients are very similar to those without fixed effects. Table 2 makes clear that the changing composition of women cannot account for the bulk of the association between infant mortality and aggregate income that we observe in our data. Instead, when there are negative economic shocks, there is an increase in mortality risk for an infant born to a given mother.

B. *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.[7] 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 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 fifteenth 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.[8] These results underscore that

[7] 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 relevant for the measure 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 (the average birth point in our data), 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, and lead GDP reflects conditions after the infant is six months of age.

[8] We also experimented with breakdowns of the in utero period. For example, in a study of the effect of the Chernobyl nuclear disaster, Almond, Edlund, and Palme (2009) 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 postconception 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 to 25 weeks after conception significant once we controlled for conditions in the last three months of pregnancy and after birth.

economic conditions around birth appear to matter most for infant survival.

The importance of economic conditions around birth for infant survival also yields clues about the likely transmission mechanisms from aggregate economic shocks to infant mortality. Low birth weight is considered an important risk factor in predicting neonatal and infant death (see, for example, the review by Lawn, Cousens, & Zupan, 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.[9] On the other hand, to have a skilled birth attendant during birth, or access to health care for children who face health shocks shortly after, may help explain our findings. Approximately 36% of neonatal deaths worldwide are a result of severe infections during birth, and another 23% are a result of asphyxia (Lawn et al., 2005). Poor economic conditions around birth could result in either a deterioration in public health services or a decrease in households' ability to pay for or otherwise access health services related to delivery, as well as pre- and post-natal care (as suggested, for example, by Paxson & Schady, 2005, in their analysis of infant mortality in Peru), both of which could lead to increased mortality in the first year of life.[10]

C. Heterogeneity

Up to this point, we have implicitly assumed that aggregate income shocks affect all mothers and children equally. However, this need not be so for a host of reasons. 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 hence 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 the child, the education and age of the mother, place of residence (urban or rural), birth parity, and the overall income of the country of residence.

TABLE 4.—INFANT MORTALITY BY MOTHER, CHILD, AND COUNTRY CHARACTERISTICS

Characteristic	Estimated Infant Mortality Rate		
Child gender		Boys	Girls
		88.4	78.5
		(833,545)	(800,814)
Mother's education		Less Than Primary	Primary or Greater
		99.5	51.3
		(1,093,757)	(540,603)
Mother's location		Urban	Rural
		61.4	94.9
		(555,742)	(1,078,618)
Mother's age	15–19	20–34	35–39
	105.4	77.0	90.1
	(296,461)	(1,151,038)	(144,052)
Birth order	First	Second to Fourth	Fifth or More
	79.1	75.6	101.5
	381,176	(804,593)	(448,591)
Country income		Low Income	Middle Income
		94.5	67.7
		(964,446)	(669,914)

Numbers in the sample are in parentheses.

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 eleven percent less likely to die in the first year of life than boys, a well-known finding in the demographic literature.[11] 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 almost twice as high as that of children born to mothers with completed primary schooling or more; that children born to young mothers (ages 15–19) and older mothers (ages 35–39) are more likely to die than those born to prime-age mothers (ages 20–34); that high-parity births (fifth 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.[12]

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 reports the coefficients on the main effect for log per capita

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

[10] It is also possible that maternal mortality can play a role in the countercyclical relation between infant survival and GDP. Children whose mothers die in birth are themselves much more likely to die (Anderson et al., 2007), and maternal death during childbirth may increase during poor economic times. We thank an anonymous referee for this suggestion. Our birth data are 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 would be biased downward.

[11] 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.

[12] The GDP data we use in this paper are measured in constant 2000 U.S. 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, a date before the start 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 that could arise had we adopted a later date by which to categorize countries.

TABLE 5.—HETEROGENEITY IN INFANT MORTALITY RATE AND GDP 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 years)	−44.20	22.52
	[10.15]***	[15.86]
Older mother (>34 years)	−36.05	−19.37
	[8.24]***	[18.51]
First births	−49.24	42.68
	[10.81]***	[16.26]**
High birth order (above fourth)	−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.

GDP and on the interaction between log per capita GDP and the given characteristic from these regressions. 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 in rural areas is significantly more sensitive to changes in economic conditions than that of children born to urban mothers. In part, this is the result of the higher mortality rates among infants born to rural women, although this does not fully explain the differences in the magnitudes we estimate.[13] A similar pattern can be seen in a comparison between mothers with less than primary education and those with primary education or greater—the increase in infant mortality during economic downturns is larger for less educated women, but from a higher base—although this difference is not significant at standard levels. Also, the point estimate on the interaction term for middle-income countries suggests larger absolute increases in infant mortality during economic downturns in low-income countries, although this difference is also 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 sensitive to changes in economic circumstances than that of boys: a 1% change in per capita GDP changes the mortality of boys by approximately 0.27 per 1,000 children born and that of girls by 0.53 per 1,000—a remarkable difference by any standard.[14]

D. *Magnitude and Sign of Shocks to Per Capita GDP*

In addition to heterogeneity by household characteristics, it is possible that the impacts on infant mortality may vary by the magnitude of the GDP shock and, perhaps, the direction (either positive or negative). To investigate this, we estimate a series of gender-specific, continuous spline regressions.[15] The results from these estimations are reported in table 6, separately for boys (upper panel) and girls (lower panel). The top row for each panel presents the results from a spline regression with a knot at 0, which allows 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. Negative shocks, however, have much larger effects on the mortality of girls than boys. For example, for negative shocks of -1.5σ or larger, a 1% decrease in log per capita GDP results in an increase in mortality of 1.05 per 1,000 girls born (with a standard error of 0.30), but an increase in mortality of only 0.53 per 1,000 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 experienced a GDP contraction of 5.9%. (Note that there are 122 such country-year events in our data.) The predicted increase in girl infant mortality during these negative shocks to aggregate income is 7.4 deaths per 1,000, approximately three times the magnitude of the average country-specific annual reduction in mortality.

[13] A 1 log-unit decrease in per capita GDP would increase the infant mortality rate of children born to rural women from 95 to 143 and that of children born to urban women from 61 to 82. The proportional, not just the absolute, change among rural mothers is thus larger.

[14] 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 1% decrease in GDP increases the mortality of boys by 0.33 per 1,000, and that of girls by 0.62 per 1,000. Comparable figures for Latin America and the Caribbean are 0.29 per 1,000 (boys) and 0.46 per 1,000 (girls). For Southeast Asia, these are 0.15 per 1,000 (boys) and 0.24 per 1,000 (girls); for South Asia, 0.72 per 1,000 (boys) and 1.43 per 1,000 (girls); and for the Middle East and North Africa region, a 1% decline in GDP results in a decrease of mortality of 0.18 per 1,000 boys born and an increase of mortality of 0.78 per 1,000 girls born. All of the coefficients on GDP in the regressions for girls are significant at the 10% 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 are not driven by a single region, including regions where a preference for boys has been well documented (for example, South Asia).

[15] 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. We then estimate with a spline regression the relationship between these two residuals.

Table 6.—Heterogeneity in Infant Mortality Rate and GDP Relation by Size of GDP Deviation from Cubic Trend, for Male and Female Children

Dependent variable	Magnitude and Direction of GDP Deviation				
Infant Mortality Rate for boys (N = 833,545)		<= 0		> 0	
		−20.75		−31.00	
		[10.10]**		[16.46]*	
	<= −1 s.d.		> −1 s.d. and <= 1 s.d.		> 1 s.d.
	−38.22		−14.73		−40.69
	[19.34]*		[21.44]		[25.10]
	<= −1.5 s.d.		> −1.5 s.d. and <= 1.5 s.d.		> 1.5 s.d.
	−52.81		−13.16		−67.25
	[22.82]**		[15.48]		[19.67]***
	<= −2 s.d.		> −2 s.d. and <= 2 s.d.		> 2 s.d.
	−59.09		−18.07		−70.82
	[26.47]**		[12.64]		[19.65]***
Infant Mortality Rate for girls (N = 800,814)		<= 0		> 0	
		−55.43		−43.71	
		[13.90]***		[14.05]***	
	<= −1 s.d.		> −1 s.d. and <= 1 s.d.		> 1 s.d.
	−75.81		−39.24		−47.27
	[24.41]***		[22.00]*		[21.84]**
	<= −1.5 s.d.		> −1.5 s.d. and <= 1.5 s.d.		> 1.5 s.d.
	−104.71		−36.60		−58.66
	[30.03]***		[15.97]**		[19.93]***
	<= −2 s.d.		> −2 s.d. and <= 2 s.d.		> 2 s.d.
	−148.47		−36.52		−69.12
	[51.38]***		[14.04]**		[19.72]***

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.

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. In other words, these findings 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.[16]

[16] In the working paper version of our paper (Baird, Friedman, & Schady, 2007), 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 also suggests that our estimates of the effect of aggregate income shocks on Infant Mortality Rate are not driven by the omitted variables that have received the most attention in the empirical literature. 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.

IV. 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 & Gopinath, 2007). A recent review stresses the welfare costs of volatility for 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. The difference with the findings for the United States, where infant mortality appears to be procyclical (Chay & Greenstone, 2003; Dehejia & Lleras-Muney, 2004) is striking but perhaps not unexpected. Economic shocks in the developing world are often much deeper than those experienced in developed countries, and it is the largest (negative) shocks that have the most serious consequences for infant mortality. Moreover, developing countries are, by definition, poorer than developed countries, and we show that within developing countries, the biggest effects of shocks on infant mortality occur in the poorest countries.

Even within developing countries, there is variation in the effects of aggregate economic shocks on infant mortality. Miller and Urdinola (2010) argue that decreases in the price of coffee in Colombia, and the attendant reduction in income in coffee-growing areas, led to reductions in infant mortality. In their review, Ferreira and Schady (2009) contrast the cases of crises in Indonesia in the late 1990s and Peru in the late 1980s. They argue that the much larger increase in infant mortality in Peru than in Indonesia 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 suggest 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 (2010) 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 suggest that it is macroeconomic conditions around birth, rather than in the early in utero period or the later half of a child's first year of life, that matter most for a child's survival in her first year. However, with annual economic 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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