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MATERNAL HEIGHT AND THE SEX RATIO

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Abstract. This paper tests the generalized Trivers Willard hypothesis in the spirit of KANAZAWA (2005), which predicts that parents with heritable traits that increase the relative reproductive success of males compared to females will have relatively more male than female offspring. We test whether taller mothers are more likely to have a male first-born using data on 400,302 mothers in a sample of Demographic Health and Surveys (DHS) from 46 developing countries. Despite using a plethora of statistical models that take into account the multi-level structure of the data, we find no strong evidence in favor of the hypothesis between and within communities, as well as on a country-by-country basis. Conversely, ANDREWS (1989)'s inverse power calculations suggest that the absence of a statistically small effect cannot be rejected.

Keywords: evolutionary psychology, sex ratio, generalized Trivers Willard hypothesis (gTWH), height, Demographic and Health Surveys

INTRODUCTION

This paper tests the generalized Trivers Willard hypothesis (henceforth gTWH) as proposed by KANAZAWA (2005), which predicts that parents with any heritable traits that increase the relative reproductive success of males compared to females will have a lower-than-expected offspring sex ratio. The gTWH is not to be conflated with the Trivers Willard Hypothesis (TWH). While the TWH considers phenotypic links, the gTWH is explicitly based on genetic correlation. Therefore, our study is a test of both the gTWH and TWH with respect to maternal height. In his test of the hypothesis, KANAZAWA (2005) finds that taller and bigger parents have relatively more boys than girls in Britain's National Child Development Survey and the British Cohort Survey. However the statistical (GELMAN 2007; GELMAN and WEAKLIEM 2007; DENNY 2008; HELLE 2008) and theoretical validity (RICKARD 2008) of this approach has been questioned. Nevertheless these findings are striking, as one expects that, in a modern society, body mass and height would lose some of their importance in terms of reproductive success.

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Overall evidence for the TWH is somewhat mixed and has been confined to a limited number of human populations (LAZARUS 2002). Most closely related to our paper, studies by DENNY (2008) on British data, as well as POLLET and NETTLE (2010) on British and Guatemalan anthropometric data of parents have rejected the gTWH. Others, such as HELLE (2008) have provided mixed evidence in favor in that there is no relationship between maternal height and the sex of her offspring, while there is an effect of maternal weight depending on maternal age at first birth and birth order. Similarly, MATHEWS et al. (2008) find that foetal sex is related to maternal diet at conception in a British dataset. In Ethiopia, a country that regularly suffers from food shortages, GIBSON and MACE (2003) detect a marked relationship between body mass index, as well as measures of fat and muscle mass and the sex of the most recent birth, while evidence by CAGNACCI et al. (2004) points to a role of pre-pregnancy weight as a driver of sex ratios. However, detecting a link between maternal nutrition and weight is problematic due to issues of endogeneity stemming from omitted variables, measurement error and reverse causality. Indeed, using the large Ethiopian Demographic and Health Survey from 2000, STEIN et al. (2004) fail to replicate results by GIBSON and MACE (2003), arguing that reverse causality or anticipation effects may obscure the estimation. Similarly, TAMIMI et al. (2003) forcefully argue that since girls weigh circa 100 g less than boys at birth, mothers with female embryos require 10% less energy intake. Therefore, the sex of the foetus influences maternal weight and nutrition rather than vice versa.

Concerning evidence from other species, CAMERON (2004) and SHELDON and WEST (2004) find little evidence of a link between maternal weight and sex ratios in horses and ungulates, respectively. However, these studies confirm important differences between pre- and post-conception conditions of mothers. We should note that many studies on animals such as the one by SHELDON and WEST (2004) pertain to the TWH and not to the gTWH. The reason is that they consider phenotypic rather than genetic correlations. ROSENFELD and ROBERTS (2004) provide evidence from laboratory mice indicating that maternal diet rather than maternal body condition influences sex ratios. More specifically, diets high in fat and low in carbohydrates increase the likelihood of male births.

An array of studies find that socio-deconomic factors and parental status correlate with sex ratios. VOLAND (1984) indicates that sex ratios co-move with land rights, inheritance practices and the marriage market. Work by ZALDIVAR et al. (1991) and CHACON and JAFFE (1996) find that parental status is either not or only weakly correlated with the likelihood of male births. Interpreting such findings is no easy task since socio-economic or parental status are correlated with a plethora of unobservables that might also be linked to, for instance, sex ratios, body stature and the timing of births.

Other theories such as GRANT (1994) and JAMES (1996) generate similar predictions and have linked testosterone levels to human sex ratios. For instance, Grant's Maternal Dominance hypothesis suggests that relatively more dominant mothers have a higher likelihood of male offspring. And dominance is associated with higher testosterone levels of women that are, for instance reflected in professional success (GRANT and YANG 2003). POLLET et al. (2009) provide indirect support for the dominance hypothesis by documenting that lower ranking wives in polygynous households in Rwanda have a higher likelihood of conceiving girls. Of course the decision to take on additional wives is endogenous to the birth record of existing wives and a wide array of observable and unobservable socio-economic variables may affect sex ratios (e.g. JACOBY 1995). If taller women are also more dominant, then our results have implications for Grant's Maternal Dominance hypothesis. Since we cannot observe dominance in our data this might be one of the unobservable channels through which height affects sex ratios. Conversely, LEIMAR (1996) outlines a general theory that is in contrast with TWH. Assume that, in a given species, high quality mothers provide better care and singlehandedly transmit quality to their offspring. Then high quality mothers should prefer daughters over boys, because grandchildren of daughters are more valuable than those of sons.

We are not the first to provide cross country evidence on the TWH. For example, WILLIAMS and GLOSTER (1992) report a positive Pearson correlation between food availability and human sex ratios at birth at the country level. Using multivariate models and country level data, DAMA (2011) showed that there is a positive correlation between sex ratios at birth and life expectancies. This result is robust to controlling for confounding factors such as fertility rates, measures of wealth and latitude.

Our work complements POLLET and NETTLE (2010), who point out that further evidence is needed using data from different human populations and ecologies. Here we provide the first test of the hypothesis for a large number of countries, and more particularly for a wide range of developing countries. We test whether taller mothers have a higher likelihood of having a male first-born using data on 400,302 mothers drawn from Demographic and Health Surveys (DHS) from 46 developing countries. Our paper uses survey data that is comparable across countries and provides pooled, country-by-country, fixed effect, as well as multi-level random intercept and random slopes estimates.

METHODOLOGY

Data

We use cross-sectional data from the Demographic and Health Surveys (DHS). Pooling DHS datasets is a common approach in a vast number of anthropometric studies. For instance, OEZALTIN et al. (2010) study the associations of maternal height and child mortality, stunting and underweight across 54 countries. The DHS are nationally representative surveys in developing countries financed primarily by, and carried out at the request of, USAID. Data such as maternal birth history are collected by interviews using standardized questionnaires. The DHS data allow one to test the gTWH using highly comparable, nationally representative data. To in-

Table 1: Sample of mothers from Demographic and Health Surveys (DHS) across 46 countries

Country	DHS-Code	N	Perc. of Sample
Armenia	AM4	4,143	1.03
Azerbaijan	AZ5	5,083	1.27
Bangladesh	BD5	9,738	2.43
Burkina Faso	BF4	9,348	2.34
Benin	BJ5	13,093	3.27
Bolivia	BO5	11,457	2.86
DRC	CD5	3,439	0.86
Congo	CG5	5,034	1.26
Cameroon	CM4	3,729	0.93
Columbia	CO4	24,973	6.24
Egypt	EG5	14,691	3.67
Ethiopia	ET4	4,432	1.11
Gabon	GA3	2,795	0.7
Ghana	GH5	3,249	0.81
Guinea	GN4	3,129	0.78
Honduras	HN5	13,521	3.38
Haiti	HT5	3,195	0.8
India	IA5	81,154	20.27
Jordan	JO5	4,759	1.19
Kenya	KE5	6,020	1.5
Cambodia	KH5	5,383	1.34
Kazakhstan	KK3	1,626	0.41
Lebanon	LB5	5,625	1.41
Lesotho	LS4	2,356	0.59
Morocco	MA4	8,597	2.15
Moldova	MB4	4,828	1.21
Madagascar	MD5	6,311	1.58
Mali	ML5	11,412	2.85
Malawi	MW4	8,885	2.22
Mozambique	MZ4	9,250	2.31
Nicaragua	NC4	8,980	2.24
Nigeria	NG5	23,211	5.8
Niger	NI5	3,555	0.89
Namibia	NM5	6,487	1.62
Nepal	NP5	7,753	1.94
Peru	PE4	18,217	4.55
Rwanda	RW4	3,510	0.88
Sierra Leone	SL5	2,873	0.72
Senegal	SN4	2,994	0.75
Swaziland	SZ5	3,390	0.85
Chad	TD4	3,582	0.89
Turkey	TR4	3,289	0.82
Tanzania	TZ4	7,518	1.88
Uganda	UG5	2,170	0.54
Zambia	ZM5	5,343	1.33
Zimbabwe	ZW5	6,174	1.54
Total		400,302	
Missing height		156,509	

crease comparability, we restrict ourselves to countries that have had a standard DHS during the last two rounds (at least DHS-5 or DHS-4) and feature maternal height. The DHS focuses on child and maternal health, providing a complete birth history of all women that are of reproductive age (15–49 years), along with anthropometric information obtained by trained interviewers. Unfortunately, paternal height and weight are only being collected systematically in the currently ongoing DHS surveys. As such, this information is not available to us in the DHS rounds that are currently available.

The DHS uses a probabilistic sampling scheme and is stratified by urban, rural and, in some countries, geographic and administrative regions. The primary sampling units come from a sampling framework based on enumeration areas in the latest population census. Within an enumeration area, all households are listed and a fixed number of them is selected at random for the survey. Our results are based on unweighted data, as the official DHS statistics guide recommends that the "use of sample weights is inappropriate for estimating relationships, such as regression and correlation coefficients." Our inference is based upon clustered standard errors since observations in a given cluster are not independent due to the sampling design. We also provide a wide array of fixed effect and multi-level models. While one may well find reasons against the official DHS stance on the use of sampling weights, our results are qualitatively insensitive to whether we weight or not.

The final data set pools surveys from 46 developing countries, which are listed in *Table 1*. Our regression sample consists of 400,302 mothers for which we have information on height and a complete birth history. In principle, the height of all interviewed women between 15–49 years is measured by trained surveyors using UNICEF boards. Note that there are 156,509 mothers whose height is missing. Women report their number of surviving and dead children at the time of the survey. We then use the birth order to determine the sex of the first-born child. While our results are qualitatively similar when we drop mothers with multiple births at first birth, we count "all male" multiple births as a male first-born. The mean height of mothers is 1.56 m (0.07 SD). Mean age is 32.6 years (8.6 SD), ranging from 13 to 49 years. Taking the logarithm of maternal height or dropping outliers leaves our results qualitatively unaffected. The dependent variable in the empirical model is a dummy variable that indicates whether the first birth is male and displays an average of .5127. Including or excluding multiple births has no significant impact on the results that follow.

Statistical Analysis

We study the association between maternal height and the likelihood of a male first-born. This approach avoids idiosyncratic stopping rules that bias the results when, for instance, the fraction of boys out of total births is correlated with maternal height.

Let *c* denote countries, *p* primary sampling units, *h* households, *m* mothers, and let *N* be sample size. Our basic specification is given by:

$$y_{cphm} = \alpha + height_{cphm}\beta + \varepsilon_{cphm}, \tag{1}$$

where y_{cphm} is a $N \times 1$ dummy variable that takes on the value of one for a male birth and zero otherwise and is an error term. In this setting the test of the gTWH boils down to a test of a significant and positive coefficient β associated with maternal height. We estimate the model both with a linear probability model and a logit specification due to the binary nature of the dependent variable. In a first instance, we estimated the models on the pooled dataset. Thereafter, we account for heterogeneity by estimating the models on each country separately and plot coefficients and associated confidence bands for between-country comparison purposes. It is possible that we find support for the gTWH within rather than across communities or countries.

We need to pay particular attention to the the multi-level structure of the data and unobservables that are correlated with maternal height in equation (1). To see this, decompose the error term into four nested components:

$$\varepsilon_{cphm} = \upsilon_c + \vartheta_{cp} + \mu_{cph} + \eta_{cphm}, \tag{2}$$

where μ_{cph} represents household, ϑ_{cp} primary sampling unit and υ_c country unobservables.

Three issues must be dealt with in any estimation of (1), given the structure of (2):

First, components of the error term given by (2) are likely to be correlated within certain groups (country, primary sampling unit and household). To account for intra-class correlation we check whether standard errors are sensitive to the level of clustering.

Second, unobservables at the country, primary sampling unit and household level may bias our estimate of β if they are correlated with maternal height. In addition, the gTWH could hold only within-communities. For both of these reasons, we provide a series of fixed effects estimates that consider within-country (only in the pooled model), within-primary sampling unit or within-household variation of height and the likelihood of a male first-born. In the latter case we achieve identification in the model by restricting our attention to households with more than one woman, which effectively purges $\upsilon_c + \vartheta_{cp} + \mu_{cph}$ from the disturbance term. In specifications in which we do not include fixed effects, observable covariates include country and primary sampling unit-specific variables such as national laws, local infrastructure, ethnic structure and customs, as well as household-specific variables such as wealth or family structure. Obviously, interactive effects cannot be ruled out.

Third, we provide random effects as well as nested random effects models in υ_c, ϑ_{cp} and μ_{cph} , under the assumption that unobservables are uncorrelated with maternal height. The nested model takes the non-independent hierarchical structure into account. We estimate various combinations of random intercept and random slope models across the nested-structure (random intercepts at country and/or primary sampling unit and/or household level, as well as random slopes at country and/or primary sampling unit level). For the fixed effects and multi-level models we report those which yielded the lowest Akaike Information Criterion. There are very few qualitative differences, however, between the results we explicitly report in what follows and those that we do not report in the interests of brevity.

RESULTS

Tables 2 and *3* present linear probability and logit model estimates, respectively. Regardless of the empirical specification, we fail to reject the null hypothesis of a zero impact of maternal height on the likelihood of a male first-born. In *Table 4*, we take into account the multi-level structure of the data by providing random intercept and random slope estimates. While all point estimates tend to be positive, they are never significant at the usual levels of confidence.

Table 2. Impact of maternal height on the likelihood of a male first-born in a linear probability model. CTY is the acronym for country, PSU for primary sampling unit, HH for household

	Male first birth			
Linear model	1	2	3	4
Maternal height	0.013167	0.021627	0.016081	−0.006999
	[0.010868]	[0.014583]	[0.013636]	[0.042675]
Fixed effects	–	CTY	PSU	HH
Clustering level of S.E.s	PSU	CTY	PSU	HH
AIC	580’816	580’722	550’515	−285’471
N	400’302	400’302	400’302	400’302

Table 3. Impact of maternal height on the likelihood of a male first-born in a logit model. CTY is the acronym for country, PSU for primary sampling unit, HH for household

	Male first birth		
Logit model	1	2	3
Maternal height	0.052701	0.086583	0.073665
	[0.0435]	[0.058377]	[0.0470031]
Fixed effects	–	CTY	–
Random effects	–	–	CTY
Clustering level of S.E.s	PSU	CTY	–
N	400’302	400’302	400’302

Table 4. : Impact of maternal height on the likelihood of a male first-born in a multi-level linear probability model with random intercepts and/or random slopes. CTY is the acronym for country, PSU for primary sampling unit, HH for household

	Male first birth				
Linear mixed Model	1	2	3	4	5
Maternal height	0.018402 [0.011711]	0.018383 [0.011715]	0.018383 [0.011715]	0.018383 [0.011715]	0.018384 [0.011715]
CTY sd (Maternal height)				6.11E-07 [2.07E-06]	1.72E-07 [4.78E-07]
PSU sd (Maternal height)					0.0079698 [0.0043911]
sd (Residual)	0.4998106 [0.0005586]	0.499664 [0.000574]	0.4996637 [0.0005738]	0.4996645 [0.000574]	0.4996566 [0.0005739]
Random effects (Nested)	CTY	CTY-PSU	CTY-PSU-HH	CTY-PSU-HH	CTY-PSU-HH
Random intercept	Yes	Yes	Yes	Yes	Yes
Random slopes	No	No	No	CTY	CTY-CL
AIC	580’807	580’808	580’810	580’812	580’814
N	400’302	400’302	400’302	400’302	400’302

Figure 1 presents country-by-country estimates of the primary sampling unit fixed effect model in order to investigate heterogeneity. For the sake of brevity, we do not report the other country-by-country models, since they are qualitatively similar. They are however available upon request. Country codes used in the figures are provided in *Table 1*. The coefficient associated with maternal height ranges from negative to positive. In the overwhelming majority of cases, however, estimates are statistically insignificant. Note also that we have not considered maternal weight as an explanatory variable, as it is both time-variant and endogenous to the birth history. Moreover, it is measured at the time of the survey and not before the first birth. In any case, while we do not report results here for the sake of brevity, the impact of maternal weight is even weaker and more ambiguous than that of height.

While we fail to reject the null hypothesis of a zero effect of maternal height, this does not imply that we reject the absence of an effect. Consider the OLS model in column (1) of *Table 2*. A standard ANDREWS (1989) inverse power calculation indicates that an estimate of 0.0224 can be confidently rejected (i.e. the power against this alternative is high), while an estimate of 0.0112 cannot (i.e. the power against this alternative is low). The latter implies that a 10 cm increase in maternal height leads to a 0.1 % increase in the probability of having a male first-born. The question is then whether the size of the effect would be important from an evolutionary standpoint. Also note that if we estimate two simple OLS regressions clustering standard errors at the primary sampling unit level, one with and the other without maternal height, the Bayesian Information Criterion is lower when maternal height is included. This indicates that the model with maternal height provides a better fit.

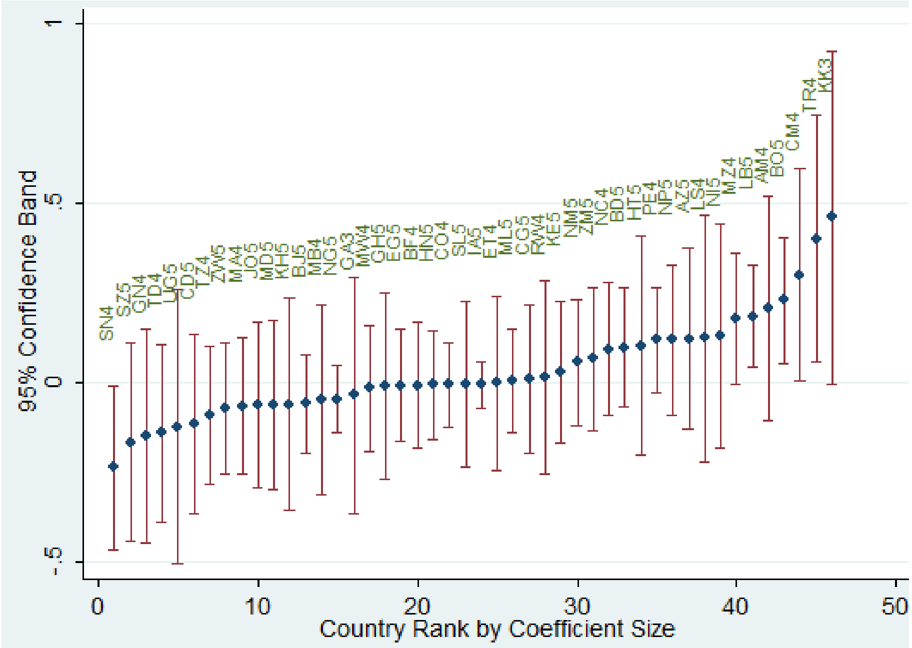

Figure 1. Impact of maternal height on the likelihood of a male first-born country-by-country. Primary sampling unit fixed effect estimates, with 95% confidence intervals obtained by clustering standard errors at the primary sampling unit level

DISCUSSION

In the spirit of KANAZAWA (2005) we tested whether taller mothers have a higher likelihood of having a male first-born using Demographic and Health Surveys (DHS) from 46 developing countries. In this particular sample we find no clear evidence in favor of the hypothesis despite employing a wide range of empirical models and robustness checks both across and within countries. This extends the findings of DENNY (2008) on British data, as well as POLLET and NETTLE (2010) on British and Guatemalan data, to many more human populations and ecologies.

If taller women are also more dominant or, as GRANT (1996) writes: “influential, ascendant, prevailing, authoritative, or high in control,” then we can also provide no support for Grant’s Maternal Dominance theory. In developed countries, height has been taken as a measure of “masculinisation”: taller women earn relatively more, care more about professional success, are more assertive as well as being more ambitious (DEADY and SMITH 2006). Similarly, NETTLE (2002) finds that in the UK taller women have both fewer children and a higher likelihood of having no children at all. In developing countries, in contrast, women rarely enjoy similar career opportunities, and therefore rarely "choose" not to have children.

Our study reveals an interesting variation in sex ratios and the correlation between maternal height and the likelihood of a male first-born across countries. We can investigate this further. In *Figure 2*, we plot the coefficient associated with maternal height from country-by-country primary-sampling unit fixed effect regressions against the annualized growth rate of Gross Domestic Product per capita (1990–2000) of the country in question. Visually (as well as statistically), there is no relationship between the magnitude of the estimated coefficient associated with maternal height and income growth at the country level. Similarly, we plot the sex ratio of first-borns against GDP growth in *Figure 3*. Here too, there is no systematic relationship that emerges. This is in contrast with WILLIAMS (1992), who report a positive Pearson correlation between food availability and human sex ratios at birth at the country level, as well as DAMA (2011) who finds a positive correlation between sex ratios at birth and life expectancies.

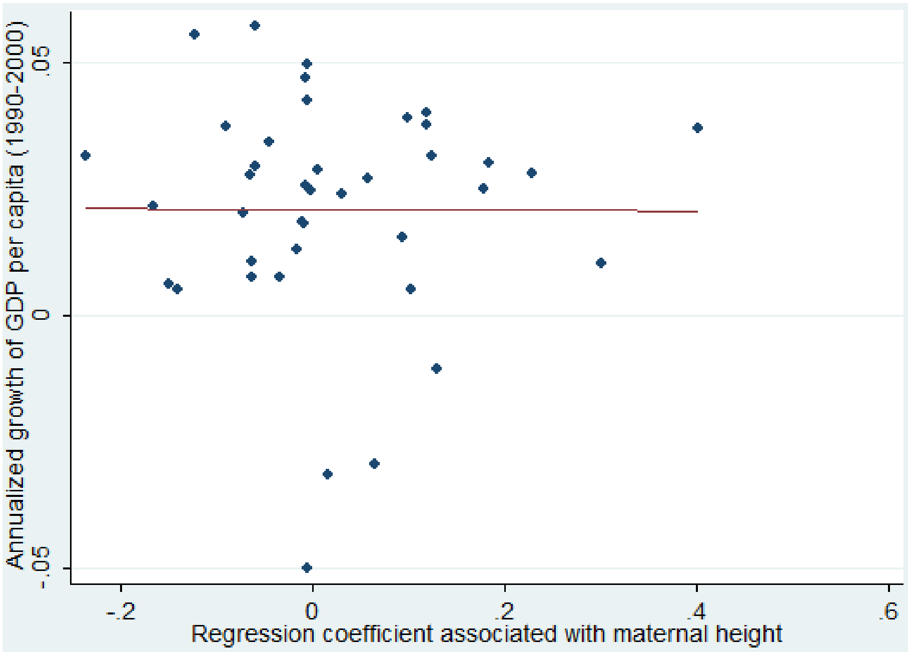

Figure 2: Coefficient associated with maternal height from primary sampling unit fixed effects models at the country level plotted against annualized growth rate of GDP per capita (1990–2000)

A few caveats to our study need to be flagged. First, a subset of mothers, in particular younger ones, might still be physically growing around the time of the first birth. Given that the gender of the first-born can differentially affect a mother's growth process, we cannot fully rule out reverse causality. Second, we cannot be sure that mother-level unobservables that affect both sex ratios and maternal body stature do not bias our parameter estimates. In order to truly detangle causal effects, it would be necessary to identify natural experiments or random shocks using instrumental variable methods (ROSENZWEIG and WOLPIN 2000 and WOOLDRIDGE and SEMYKINA 2005). To this end, one would need exogenous instruments that significantly impact maternal height, and therefore indirectly contribute to the likelihood of a first birth (e.g. timing of drought shocks in subsistence agriculture com munities), but that have no direct impact on sex ratios. Finding such instruments is likely to be a tall order of business. Third, one can imagine that misreporting voluntarily or not, as well as recall problems (in terms of the sex of children) may affect our results. The model proposed by HAUSMAN et al. (1998) might be one avenue to explore as it can provide a means of estimating such misclassification. Fourth, sex-selective abortion cannot be ruled out. While this is less of an issue in the poorer African countries included in the sample, this practice is widespread in countries such as India. Finally, while DHS surveys are standardized across countries and nationally representative, there is unobservable variation in the representativeness of each sample. It would be interesting to contrast our results with those that would be obtained using different data sources, such as the MICS from UNICEF or the LSMS from the World Bank. Another extension would be to analyze repeated cross sections for one country. This might reveal subtle time trends in human sex ratios.

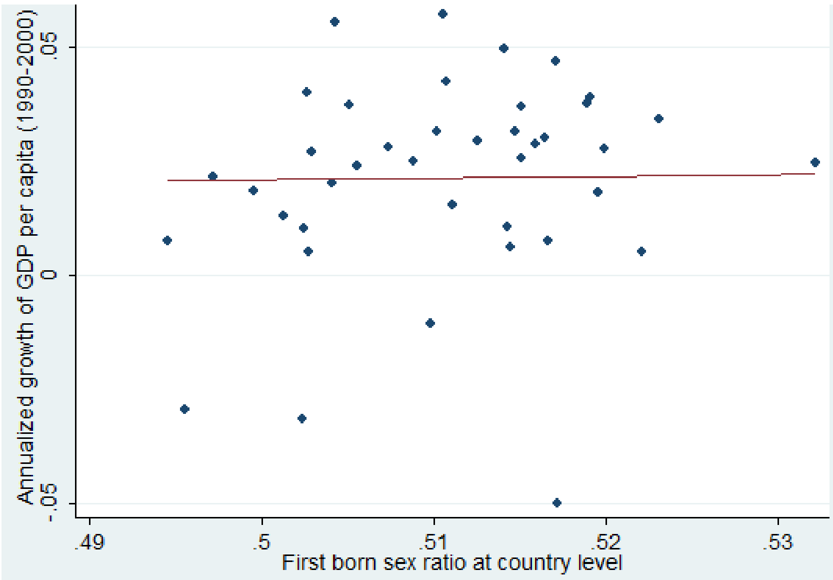

Figure 3: First-born sex ratio at the country level plotted against annualized growth rate of GDP per capita (1990–2000)

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