Online Appendix

1 Breastfeeding and Maternal Fertility

Breastfeeding reduces fertility by interrupting the release of the Gonadotropin-releasing hormone that is necessary to begin ovulation. Breastfeeding also may increase the level of the hormone prolactin that is an ovulation inhibitor (Blackburn, 2007). This leads to a period of lactational amenorrhea in nursing mothers, during which they do not menstruate. Calories are diverted away from the mother while breastfeeding, and in developing countries this may lead to malnutrition that prevents ovulation. The caloric requirements of breastfeeding may conversely lead to children being weaned earlier in developing countries if their mothers become pregnant (Weis, 1993; Thapa, 1987, Huffman et. al., 1987). Mothers on average need to consume anywhere in the range of 300-670 calories extra over was needed to maintain pre-pregnancy weight to breastfeed exclusively without further weight loss (Riordan, 2004; Dewey, 1997; Institute of Medicine, 1991). Women in developing countries such as Egypt are unlikely to build up significant fat stores in pregnancy, and require additional calories close to the upper end of this range to exclusively breastfeed for an extended period. These nutritional and fecundity costs potentially lead to changes in maternal nursing in Egypt in response to child gender and number of offspring.

2 Son Preference in Egypt

There is a wealth of evidence in the medical and social science literatures documenting the presence of son preference in human capital investments and fertility choice in Egypt (E.g. see Obermeyer, 1995; El-Gilany and Shady, 2007; Yount, 2005). Women also have less autonomy than men in household decision-making, have fewer rights on owning and inheriting property, and face greater restrictions on working outside the home. The Egyptian female labour force participation rate in 2000 was 37%, which is much lower than the 46% participation rate for women in South Asia in the same year (World Bank, 2004). The roots of this gender bias are to a large extent attributable to conservative social norms that lead to gender inequality in marital and divorce rights, inheritance laws, and employment (Inhorn, 1996; Doumato and Posusney, 2003; Mir-Hosseini, 2000).[1] The combined effect of these institutions is to make sons more valuable than daughters, as sons are likelier to contribute earned income to the household and keep inherited property within the family. We can therefore expect that desired fertility outcomes will reflect this son preference, and in combination with existing sex-composition of offspring will affect birth spacing and indirectly breastfeeding duration as well.

3 Framework for Analysis

The theory in Jayachandran and Kuziemko (2011) predicts breastfeeding duration increases with birth order as mothers successively space their births more. The male advantage in breastfeeding is greatest in the middling birth orders, where girls are weaned faster to allow quicker conception of a following son before the mother's fertility ends. The gender gap then closes at higher birth orders, as mothers' desire for more children declines irrespective of offspring sex composition. Breastfeeding duration increases with more older male siblings, as a satiated desire for sons leads to fewer future births and more birth spacing. Finally, breastfeeding duration increases as mothers approach ideal total fertility, and increases further for children born beyond the ideal as they are not weaned to allow more pregnancies. This final prediction is important to isolate the fertility preference effect from other mechanisms that may lead to longer breastfeeding with increasing birth order, such as "learning-by-doing" or reduced maternal attachment to the labour market.

[1] Women were given the right to unilateral divorce in 2000 via the passage of the 'khula' law, but the new democratically elected government has attempted reversing the legislation.

4 Breastfeeding and Diarrheal Incidence

I investigate how the patterns of breastfeeding duration found in the results match diarrheal incidence patterns among children. I do this as diarrhea is the likeliest detrimental health effect on children of not being breastfed. I define a binary outcome variable taking the value 1 if the child has had diarrhea in the two weeks preceding the survey, and 0 otherwise. I then use probit models to estimate the same specifications as above with this new outcome of interest among children aged 0-36 months at the time of the survey, as these are ages during which I assume children are being breastfed.[2] To control for regional and seasonal variation in disease environment, I include region[3] and interview month fixed effects. I also use child birth year fixed effects to better control for time trends and annual variation.

The results on diarrheal incidence are reported in Table 1. Male children have a noticeably higher incidence of diarrhea compared to female children, reflecting the well-known greater health fragility of boys in childhood. The male disadvantage is estimated at 1.3-1.5 percentage points according to columns (1) and (2). Diarrheal incidence declines at a diminishing rate with increasing birth order according to the estimates in column (1), closely resembling the pattern of increasing breastfeeding duration with higher birth order found previously. In column (2) I find a 0.7 percentage point decline in diarrheal incidence for children born at each birth order that approaches mothers' ideal fertility. The effect is significant at the 5% level.[4] These results match the patterns of breastfeeding duration found in response to distance to ideal fertility in the previous section. The evidence therefore points to health consequences for children arising from breastfeeding behaviour that changes with fertility outcomes. There is however no sharp decline in diarrheal incidence for children born at mothers' ideal fertility

[2]The results using linear probability models are identical and available from the author on request. There are no discernible patterns in infant mortality as in the original study, most likely as Egypt is a wealthier country.

[3]In my data Egypt's governorates are divided into six broad regions.

[4]I find no gender differences in the estimated diarrheal incidence in columns (1) and (2), probably due to the large male health disadvantage in early childhood that breastfeeding cannot compensate for.

ceiling to match the corresponding increase in breastfeeding at ideal fertility found in the previous results. I find small declines in diarrheal incidence in the presence of a higher fraction of older male siblings in columns (3) and (4), but these are statistically insignificant.

5 Robustness Checks

I implement the same estimations from the previous section with vaccinations as the dependent variable. Vaccinations are not a fertility-related health investment, so I should not find the same pattern of results with respect to the regressors of interest as I do for breastfeeding.[5] The results of these estimations are in Table 2. In column (1) I find that there is a statistically significant male advantage of 0.029 vaccinations. However vaccinations received do not vary significantly with birth order unlike breastfeeding duration. In column (2) I find that vaccinations increase as birth order approaches ideal fertility, as with breastfeeding duration. However vaccinations received decline discontinuously once mothers achieve ideal fertility, and decline further still with each birth beyond the ideal total. This is the opposite of what I find for breastfeeding duration, and suggests a reduced maternal desire to protect children beyond the ideal fertility total even if breastfeeding increases. These results collectively point to son-biased fertility preferences driving the results on breastfeeding.

I further examine whether I find a supportive pattern of coefficients from the regressors of interest where the outcome of interest is the existence of a younger sibling. The results are in Table 3. In column (1) I find that children are significantly less likely to have younger siblings if they are male, providing support for the argument that boys receive more breastfeeding due to reduced maternal desire to conceive again after having sons. This effect remains in columns (2) and (3). The coefficients on the linear and quadratic birth order terms also support the desired fertility argument, with probability of having a younger sibling declining at a diminishing rate with birth order. In column (2) I find more supportive evidence, with

[5] The vaccines recorded in the dataset for each child are the BCG and measles vaccines, and three rounds each of the polio and DPT vaccines.

the probability of having a younger sibling declining as child birth order approaches ideal fertility, and discontinuously declining at birth orders thereafter. In column (3) I find having a younger sibling is less likely if the child has an older brother, but this effect is smaller if the child is male. Again this mirrors what I find for breastfeeding duration, which also increases in the presence of an older male sibling, but less so if the child is male. These estimates also point to desired fertility being the mechanism behind the main results.

6 Weaning and Exclusive Breastfeeding

The probit marginal effects of fertility outcomes on weaning rates for children aged 0-6 months at the time of the survey are shown in Table 4. In column (1) I find that the probability of being weaned during these ages declines by 1.6 percentage points with each unit increase in child birth order, and the effect diminishes negligibly at higher birth orders. These results match those found for mean breastfeeding duration among children of all ages, showing that birth order affects breastfeeding behaviour at early life as well as beyond infancy. Column (2) shows that the effect of birth order on weaning is gender-neutral, implying that the male advantage in breastfeeding duration occurs beyond ages 0-6 months. In column (3) there is a one-time decline in the probability of being weaned of 1.4 percentage points for children born at or beyond their mother's ideal total fertility. This effect constitutes 36.75% of the baseline probability of being weaned in the sample (which is 3.81%). It is also in line with the previous results showing longer breastfeeding duration among children of all ages who are born in excess of their mother's ideal. There is no significant effect of having at least one older brother on weaning in column (4), but in column (5) I find a decline in weaning as the fraction of males among older siblings increases. This decline peaks at 1.4 percentage points when all older siblings are male. Again, this fits the pattern of results seen previously with breastfeeding duration increasing with more older brothers present. However unlike the previous results, the effect of the male fraction of older siblings on weaning is gender neutral.

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Table 1: Child Had Diarrhea in Past Two Weeks

	Had Diarrhea in Past Two Weeks			
	(1)	(2)	(3)	(4)
$Male$	0.015***	0.013***	0.010	0.011*
	(0.005)	(0.004)	(0.006)	(0.006)
$Birth\,Order$	-0.015***	-	-	-
	(0.005)			
$Birth\,Order^{2}$	0.002***	-	-	-
	(0.001)			
$\Delta Ideal$	-	-0.007**	-	-
		(0.003)		
$\mathbb{1}(\Delta Ideal \geq 0)$	-	0.000	-	-
		(0.008)		
$\mathbb{1}(\Delta Ideal \geq 0) * \Delta Ideal$	-	0.017***	-	-
		(0.004)		
$\mathbb{1}(Male\,Fraction > 0)$	-	-	-0.007	-
			(0.007)	
$Male\,Fraction$	-	-	-	-0.010
				(0.008)
Observations	28,758	25,532	28,758	28,758
Diarrhea %	18.7%	19.0%	18.7%	18.7%
Covariates	Yes	Yes	Yes	Yes

Notes: The regression sample consists of children aged 0-36 months at the time of the survey. Coefficients are marginal effects from Probit estimations. Robust standard errors clustered by mother in parentheses. Additional regressors include the standard covariates, dummy variables for Governorate and month of interview, and child age-in-months fixed effects. Columns (3)-(4) also include linear birth order controls, and child gender-sibling sex interactions which are insignificant. *** Significant at 1% ; ** Significant at 5% ; * Significant at 10%.

Table 2: Vaccinations and Fertility Outcomes

	Number of Vaccinations		
	(1)	(2)	(3)
$Male$	0.029*** (0.010)	0.015 (0.010)	0.015 (0.013)
$Birth\,Order$	-0.018 (0.015)	-	-
$Birth\,Order^2$	-0.004* (0.002)	-	-
$\Delta Ideal$	-	0.054*** (0.012)	-
$\mathbb{1}(\Delta Ideal \geq 0)$	-	-0.085*** (0.021)	-
$\mathbb{1}(\Delta Ideal \geq 0) * \Delta Ideal$	-	-0.079*** (0.015)	-
$\mathbb{1}(Male\,Fraction > 0)$	-	-	-0.032 (0.020)
$Male * \mathbb{1}(Male\,Fraction > 0)$	-	-	0.029 (0.021)
Observations	45,329	40,095	45,329
Covariates	Yes	Yes	Yes
R-Squared	0.702	0.729	0.702

Notes: Robust standard errors clustered by mother in parentheses. Additional regressors include linear and quadratic terms of mother's current age, dummy variables for mother's educational attainment, linear and quadratic terms of the child's birth year, a dummy variable for whether the mother lives in a rural area, and child age-in-months fixed effects. Columns (3) also includes a linear birth order term. *** Significant at 1% ; ** Significant at 5% ; * Significant at 10%.

Table 3: Younger Sibling and Fertility Outcomes

	Younger Sibling		
	(1)	(2)	(3)
$Male$	-0.040***	-0.036***	-0.057***
	(0.003)	(0.004)	(0.005)
$Birth\,Order$	-0.137***	-	-
	(0.004)		
$Birth\,Order^{2}$	0.013***	-	-
	(0.000)		
$\Delta Ideal$	-	-0.026***	-
		(0.003)	
$\mathbb{1}(\Delta Ideal \geq 0)$	-	-0.131***	-
		(0.006)	
$\mathbb{1}(\Delta Ideal \geq 0) * \Delta Ideal$	-	0.030***	-
		(0.004)	
$\mathbb{1}(Male\,Fraction > 0)$	-	-	-0.126***
			(0.005)
$Male * \mathbb{1}(Male\,Fraction > 0)$	-	-	0.034***
			(0.007)
Observations	45,329	40,095	45,329
Covariates	Yes	Yes	Yes
R-Squared	0.360	0.368	0.359

Notes: Robust standard errors clustered by mother in parentheses. Additional regressors include linear and quadratic terms of mother's current age, dummy variables for mother's educational attainment, linear and quadratic terms of the child's birth year, a dummy variable for whether the mother lives in a rural area, and child age-in-months fixed effects. Column (3) also includes a birth order term. *** Significant at 1% ; ** Significant at 5% ; * Significant at 10%.

Table 4: Weaning During Ages 0-6 Months

	Weaned Aged 0-6 Months				
	(1)	(2)	(3)	(4)	(5)
$Male$	-0.001 (0.005)	-0.003 (0.014)	-0.002 (0.005)	-0.001 (0.004)	-0.001 (0.004)
$Birth\,Order$	-0.013** (0.006)	-0.015** (0.007)	-	-	-
$Birth\,Order^2$	0.002*** (0.001)	0.002*** (0.001)	-	-	-
$Male^*\,Birth\,Order$	-	0.005 (0.009)	-	-	-
$Male^*\,Birth\,Order^2$	-	-0.001 (0.001)	-	-	-
$\Delta Ideal$	-	-	0.001 (0.003)	-	-
$\mathbb{1}(\Delta Ideal\ \geq\ 0)$	-	-	-0.014** (0.007)	-	-
$\mathbb{1}(\Delta Ideal\ \geq\ 0)^*\,\Delta Ideal$	-	-	0.002 (0.004)	-	-
$\mathbb{1}(Male\,Fraction\ >\ 0)$	-	-	-	-0.006 (0.005)	-
$Male\,Fraction$	-	-	-	-	-0.014** (0.006)
Observations	5,224	5,222	4,664	5,222	5,222
Weaned %	3.81%	3.81%	3.81%	3.81%	3.81%
Covariates	Yes	Yes	Yes	Yes	Yes

Notes: The regression sample consists of children born within six months of the survey. Coefficients are marginal effects from Probit estimations. Robust standard errors clustered by mother in parentheses. Additional regressors include the standard covariates, dummy variables for Governorate and month of interview, and child age-in-months fixed effects. Columns (4) and (5) also include linear birth order controls. *** Significant at 1% ; ** Significant at 5% ; * Significant at 10%.