The Effect of Female Education on Adolescent Reproductive Health in Bangladesh[1]

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Abstract Female education is presumed to influence fertility. The effect of female education, however, may be estimated incorrectly due to endogeneity of female education at the community and individual levels. This study estimates the causal effect of female education on adolescent reproductive health outcomes in Bangladesh and addresses the potential sources of endogeneity by applying instrumental variables (IV) constructed from education programs introduced nationwide in the 1990s in Bangladesh. We find that female education significantly delays marriage and childbearing, and reduces fertility during adolescence. The IV estimates of the effect of female education on ages at first marriage and first live birth are significantly different from the ordinary least squares (OLS) estimates. The education programs are found to have improved adolescent reproductive health by enhancing female education.

Female education is presumed to affect fertility through its influence on the proximate determinants of fertility, including exposure to intercourse, contraceptive use, and proportion of population married (Bongaarts 1978; Davis and Blake 1956), which reflect an individual's reproductive health behaviors. Two generations of research have broadly examined this presumption. The first generation of research, following the pioneering work of Caldwell and Cochrane on the role of female education, consists mainly of observational studies assessing the correlation between educational attainment and fertility and its proximate determinants (Caldwell 1979; Cochrane 1979). However, the proposed evidence based on the correlation has been questioned for its lack of interpretation as a causal effect of female education.

The estimated correlation based on observational studies, after controlling for observed covariates, may be subject to bias due to unobserved variables at the individual and community levels causing endogeneity (or what is called confounding in health science research). When individuals alter their reproductive behaviors in response to factors observed by them but not by researchers and when the factors are related to an individual's decision regarding schooling, the estimated correlation may not represent the true causal effect of female education. For example, educated women may have grown up in comparatively modern communities or households with implications not only for education but also for fertility (Desai and Alva 1998; Diamond et al.

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1999). Women in such a community may be more likely to work for income or assume responsibility, and may feel less pressured to marry upon puberty or initiate childbearing. Likewise, educated women are more likely to come from parents with a high educational attainment in high social strata. Such parents may have provided their daughter with education opportunities, as well as exposure to modern institutions which may have shaped her reproductive behaviors in her adulthood (Jeffery and Basu 1996). The magnitude and direction of the potential bias depend on the unobserved relationship between the omitted variables and reproductive and schooling decisions. The unobserved relationship cannot be determined as a priori knowledge and further complicates the interpretation of the estimated correlation.

The second generation of research emerged to address the lack of evidence of the causal effect, as opposed to correlation, as both the quantity and quality of available data improved and research methodology evolved. These studies were explicitly designed to address the potential endogeneity problem, for example, by using quasi-experimental methods, including the instrumental variable (IV) and regression discontinuity methods (Angeles et al. 2005; Black et al. 2004; Breierova and Duflo 2004; Royer and McCrary 2006). Most of these studies supported the direction of the conventionally observed correlation in observational studies between female education and fertility. At the same time, the magnitude of the estimated effect of female education was often found to change significantly when endogeneity was accounted for.

The second generation of research greatly enhanced the knowledge of the overall causal effect of female education on fertility; however, it has focused relatively little attention on the proximate determinants of fertility compared to the first generation of research. Consequently, there is a dearth of empirical evidence of the mechanisms through which female education influences fertility. Understanding the effect of female education on the proximate determinants as well as on fertility is essential. It is of particular interest to policy makers involved in developing effective strategies to reduce fertility in low-income countries, where female education programs are considered a priority in promoting reproductive health (Bledsoe et al. 1999).

This study aims to fill in the knowledge gap by estimating the causal effect of female education on an array of reproductive health outcomes, including fertility and its proximate determinants, using the IV method to specifically address the sources of endogeneity. In particular, this study estimates the causal effect of female education on adolescent reproductive health in Bangladesh, where persistently high adolescent fertility remains a major policy concern (Nahar and Min 2008; NIPORT et al. 2009). In 2007 the age-specific fertility rate among women ages 15–19 was 126 per 1,000 women, which was one of the highest worldwide. In addition, the age-specific fertility rate among women ages 15–19 decreased only by 31% (from 182 to 126 per 1,000 women) between 1989 and 2007, while that among women ages 25–29 almost halved (from 225 to 127 per 1,000 women) in the same time period (NIPORT et al. 2009). Teenage mothers are physiologically immature and more likely to experience pregnancy-related complications, which increase morbidity and mortality of both mothers and newborns (Nahar and Min 2008). Also, early initiation of childbearing can lead to a high fertility rate and a rapid population growth at the country level.

One of the major factors behind the persistently high adolescent fertility rate in Bangladesh is early exposure to intercourse (Field and Ambrus 2008; Nahar and Min 2008). In rural

Bangladesh premarital sex is taboo, therefore adolescent fertility is mainly determined by age at marriage and cohabitation with a husband (Agarwal 1994; Jeffery and Basu 1996). The high prevalence of adolescent marriage in Bangladesh may reflect traditional marriage practices and economic circumstances (Amin 1996; Davis and Blake 1956; Field and Ambrus 2008). In traditional Bangladeshi society, women may obtain their social status through marriage and childbearing (Mason 1987). Most marriages, especially in rural areas, are arranged by parents or guardians (Amin 1996; Barkat and Majid 2003). The purity or virginity of a bride is a critical condition of the marriage market (Begum 2003; Field and Ambrus 2008). Since younger girls are less likely to have been exposed to sexual encounters, they are often preferred by prospective husbands. Indeed, the latest Bangladesh Demographic and Health Survey reported that 78% of women ages 20–49 had married by age 18 (NIPORT et al. 2009).

At the same time, female educational attainment in Bangladesh was persistently lower than that of males, and the gap widened at the secondary education level (Liang 1996). To address the problem, a number of education programs were introduced during the 1990s at the primary and secondary education levels, especially in nonmunicipal areas (Ahmed and Sharmeen 2004; Ahmed et al. 2007; Arends-Kuenning and Amin 2000; Raynor and Wesson 2006). These programs include abolishing tuition at the primary education level, providing food rations to poor households at the primary education level, providing financial assistance to female students at the secondary education level, and constructing primary and secondary schools. At the primary and secondary education levels, the gap between boys and girls was eliminated. Between 1991 and 2000 the gross enrollment rates of both boys and girls at the primary education level increased to 97%, from 81% for boys and 70% for girls (Ahmed et al. 2007). Likewise at the secondary education level, the proportion of female students increased from 34% in 1990 to 52% in 2005, which suggests that more girls were enrolled than boys (BANBEIS 2006).

The changes in the education system provide a quasi-experimental setting that allows estimation of the causal effect of female education on adolescent reproductive health. Employing IVs constructed from the aforementioned education programs, this study links the highest grade achieved and three reproductive health outcomes, including age at first marriage, age at first live birth, and adolescent fertility, of women ages 20-44 and residing in nonmunicipal areas in Bangladesh. The specific aims of the study are: (1) to estimate the effect of the education programs on educational attainment; (2) to estimate the effects of female education on adolescent fertility and its proximate determinants, including age at first marriage and age at first live birth; and (3) to compare the estimated coefficients of female education between the ordinary least squares (OLS) and the two-stage least squares (2SLS) methods. This study is novel in its examination of the mechanisms through which female education influences reproductive health and provides further insight into the potential impact of education programs on reproductive health.

1. Empirical Framework

Data

This study uses three datasets: (1) data on secondary schools, (2) data on population size, and (3) data on an individual's reproductive health and educational attainment. First, the data on secondary schools come from the database managed by the Bangladesh Bureau of Educational

Information and Statistics (BANBEIS). The database is based on a school census conducted in 2006 and includes each school's location and the year it opened. Second, the data on population size come from the 1981 Bangladesh population census (Bangladesh Bureau of Statistics 1983). The data provide the population size at the subdistrict level by age groups. We are interested in the population sizes of secondary school age groups, and the nearest found in the census data is the age group 10–17. Population sizes in other years are approximated under the assumption of an exponential growth at a rate r of .026 (UNICEF 2010). Let y_{ijt} denote the population size of age group 10–17 of woman i's subdistrict of residence j in year t. It is approximated as:

$$y_{ijt} = y_{ij,1981} \exp[0.026(t-1981)].$$

Finally, the data on an individual's reproductive health and educational attainment come from the 2007 Bangladesh Demographic and Health Survey (BDHS 2007), which is one of the largest demographic surveys in Bangladesh. In the full survey, 10,400 households and 10,996 ever-married women ages 10–49 were interviewed between March 2007 and August 2007 (NIPORT et al. 2009). The survey collected various information, including background characteristics, educational attainment, and reproductive history from eligible women. Our study analyzes 6,930 ever-married women ages 20–44 residing in nonmunicipal areas. Women 19 or younger are excluded from the study sample because their fertility rate is right-censored. The data on secondary schools and population size by year are aggregated at the subdistrict level to be matched with the data from the BDHS 2007. The final dataset includes 230 subdistricits, approximately one half of the total subdistricts in Bangladesh. Descriptive statistics of the individuals and the subdistricts are presented in Table 1.

Table 1. Descriptive statistics

Characteristics	Proportion
Individual Characteristics[a] (N=6,930)	
Age group	
20-24	26.4
25-29	22.9
30-34	19.6
35-39	17.9
40-44	13.2
Educational attainment	
No education	33.1
Incomplete primary	22.3
Complete primary	8.6
Incomplete secondary	22.7
Complete secondary or higher	13.3
Sub-district Characteristics (N=230)	
Women ages 5-29 attending school in 1981[b]	17.0

[a] Individuals are weighted by the sampling probability
[b] Data are from the 1981 population census

Education Programs
Educational attainment among girls was persistently lower than that among boys until the early 1990s. This situation likely reflects constraints of demand and supply (Ahmed and Sharmeen 2004), namely, affordability and accessibility. Parental perceptions of the returns on investments in daughters' education may be low in Bangladesh, where girls are expected to marry and subsequently belong to their husbands' households (Basu 1989; Das Gupta et al. 2003; Mason 1987). The dowry system, moreover, adds to the direct cost of raising daughters (Amin and Cain 1997) and may leave no financial resources for their schooling. Widespread poverty and limited job opportunities suitable for educated women also discourage parents from investing in a daughter's education (Liang 1996). In addition, adolescent girls in traditional Bangladeshi society are often allowed limited mobility, as their parents want to control their premarital sexual exposure because virginity is a critical condition of marriage (Ahmed and Sharmeen 2004; Begum 2003; Field and Ambrus 2008). Schooling of girls therefore is reported to be a concern, especially if it involves traveling a significant distance outside the community (Ahmed and Sharmeen 2004). Even though female education is considered a desirable attribute in the marriage market, the perceived risk of daughters' exposure to boys and men while traveling to school may outweigh the perceived benefits of education, resulting in parents withdrawing their daughters from school upon puberty.

During the 1990s the Bangladesh government launched a number of education programs to reduce the constraints to education on both the demand and the supply sides, especially in nonmunicipal areas at the primary and secondary education levels (Ahmed and Sharmeen 2004; Ahmed et al. 2007; Arends-Kuenning and Amin 2000; Raynor and Wesson 2006). In 1990 primary education (grades 1–5) became compulsory and free nationwide (Hossain 2004). In 1993 the government introduced the pilot program Food for Education (FFE) that provided food rations to poor households sending their children to primary school (Meng and Ryan 2007; Ryan and Meng 2004). In 1994, based on positive responses observed in the primary school enrollment rate, FFE expanded to all 460 nonmunicipal subdistricts in two stages: at the geographic and the individual levels (Ryan and Meng 2004). At the geographic level, two to three underdeveloped counties were selected in each of the nonmunicipal subdistricts based on their economic development and literacy rates. The program covered all of the registered primary schools and one religious school within each selected county (Ryan and Meng 2004). At the individual level, households sending their children to eligible primary schools were selected within each program county based on a set of four criteria.[2] Households meeting at least one of the criteria were entitled to food rations of 15 to 20 kilograms (kg) of wheat or 12 to 16 kg of rice per month (depending on the number of children attending primary school) on the condition that the children maintain an attendance rate of 85%. The estimated average monetary value of the food rations a household received was 120 taka (US$1.70)[3] per month (Ravallion and Wodon 2000). Nearly 27% of primary schools and 13% of pupils in the country were under FFE by 2000.

In 1999 FFE was supplemented by the Primary Education Stipend Project (PES Project), which was cash based and provided 25 taka (US$0.40) per month to eligible households in all rural

[2] The four criteria are (1) the household owns less than half an acre of land, (2) the household head is a day laborer, (3) the household head is female, or (4) the household has limited income.
[3] One U.S. dollar equaled 69.10 Bangladeshi taka on March 1, 2010.

non-FFE areas (Hossain 2004; Tietjen 2003). In 2002 both FFE and the PES Project were replaced by the Primary Education Stipend Program (PESP), which provided 100 to 125 taka (US$1.40–1.80) per month (depending on the number of children attending primary school) to qualifying households in all counties in nonmunicipal subdistricts (Ahmed and Sharmeen 2004). In 2003 the estimated average annual direct costs (fees and other payments) and indirect costs (textbooks, uniforms, private tutoring, and transportation) of primary education were 64 taka (US$0.90) and 892 taka (US$12.90), respectively. This suggests that the PES Project covered the direct costs and that FFE and PESP provided more than the total costs (Ahmed and Sharmeen 2004).

At the secondary level in 1994 girls in all nonmunicipal subdistricts were given free tuition and a stipend (Liang 1996). Under the program, female students were required to meet three conditions[4] (Asian Development Bank 1993; Asian Development Bank 1999; Asian Development Bank 2002; Asian Development Bank 2008; Uniconsult International Limited 2006; World Bank 2002a; World Bank 2002b; World Bank 2008). The yearly stipend increased as girls proceeded to higher grades: 300 taka (US$4.30) for 6th graders, 360 taka (US$5.20) for 7th graders, 420 taka (US$6.10) for 8th graders, and 720 taka (US$10.40) for 9th and 10th graders. Also 9th graders were provided a book allowance of 250 taka (US$3.60), and 10th graders were provided examination fees of 730 taka (US$10.50). In 2003 the estimated average annual direct costs (fees and other payments) and indirect costs (textbooks, uniforms, private tutoring, and transportation) of girls' secondary education were 346 taka (US$5.00) and 3,191 taka (US$46.00), respectively. This suggests that the stipend covered the direct costs but not the full indirect costs (Ahmed and Sharmeen 2004).

At the same time, both external donors and nongovernmental organizations supported school construction throughout the country (BANBEIS 2006). As a result, between 1990 and 2000 the number of secondary schools increased by approximately 47%, from 10,448 to 15,403 (BANBEIS 2010a). While the external donors and nongovernmental organizations did not establish priorities for school locations, the empirical evidence suggests that schools may have been constructed regressively with respect to the education levels of the subdistricts. The multiple regression model suggests that subdistricts with a lower female attendance rate (obtained from the 1981 population census) were allocated more secondary schools between 1990 and 1999 after controlling for the population size of ages 10–17 in 1981 (Table 2). While this evidence is crude, it suggests that the secondary school construction targeted the areas with greater needs.

[4] The conditions are (1) maintaining an attendance rate of 85% or higher, (2) passing the annual final exams with a score of 45% or higher, and (3) staying unmarried until the secondary school certificate examination (SSC) or age 18.

Table 2. Regression of the number of secondary schools

	Coefficients		(SE)
Female attendance rate[a]	-0.28	**	(.099)
Population aged 10-17 (10,000)[a]	1.04	**	(.369)
Constant	13.76	***	(2.250)
Adjusted R^2	0.05		
F-statistics	7.39		
N	230		

[a] Data are obtained from the 1981 population census
** p<.01; *** p<.001

Identification Strategy

This study estimates the causal effect of female education on reproductive health outcomes in Bangladesh in order to assess the mechanisms through which education influences adolescent fertility. The effect of female education is assessed within a framework relating background factors, proximate determinants of fertility, and fertility (Bongaarts 1978; Davis and Blake 1956; Jeffery and Basu 1996). As one of the background factors, female education is presumed to influence fertility through its effects on the proximate determinants of fertility. Bongaarts proposed a set of proximate determinants, including “exposure factors,” “deliberate marital fertility control factors,” and “natural marital fertility factors” (Bongaarts 1978).

This study focuses on exposure factors measured by two variables, age at first marriage and age at first live birth, to estimate the causal effect of female education on adolescent fertility among ever-married women ages 20–44 residing in nonmunicipal areas. In particular, we examine the probability of first marriage by age 15 and the probability of first live birth by age 16 to estimate the effect of female education on early marriage and early exposure to intercourse. The age at first marriage in the BDHS 2007 is defined as the age of cohabitation with a husband rather than formal marriage. Because cohabitation may occur sometime after formal marriage in Bangladesh (NIPORT et al. 2009), age at cohabitation is a desirable measure reflecting risk of pregnancy. As we assess ever-married women, the probability of first marriage by age 15 is conditional on marriage by the time of the survey interview. Specifically, 95% of all women ages 20–44 and located in the household questionnaire (but not necessarily interviewed for the women’s questionnaire) had ever married by age 20.

Education is presumed to be endogenous in the context of its causal effect on the proximate determinants of adolescent fertility due to omitted variables at the individual and community levels. To control for omitted variables at the individual level, this study employs IVs constructed from the education programs described in the preceding section. The programs were introduced at different times in different nonmunicipal areas, which suggests that variations in individuals’ exposures to the programs were determined by both the accessibility of a secondary school and an individual’s year of birth.

The first IV is intended to capture the rapid expansion in the accessibility of secondary education. Bangladeshi children normally attend primary school between the ages of 6 and 10 and enroll in secondary school at age 11. Therefore the first IV is the number of secondary schools in the subdistrict when an individual reaches age 11, standardized per 10,000 population of ages 10–17 (hereafter referred to as the number of schools). For instance, if woman i in subdistrict j was born in 1985, the measure is:

$$\frac{P_{j,1996} \times 10{,}000}{y_{ij,1981} \times \exp[0.026(1996-1981)]}$$

where $P_{j,1996}$ is the number of secondary schools in the subdistrict in 1996 (when the woman was 11).

The second IV is intended to capture the increasing exposure to the financial incentives among younger cohorts. Given that the programs were introduced in the 1990s, women born around the early 1980s, specifically between 1979 and 1982, are more likely to have been partially exposed to free and compulsory primary education and stipend assistance at the secondary level. Women born in or after the mid-1980s, specifically after 1983, are more likely to have been fully exposed to free and compulsory primary education and stipend assistance at the secondary level and partially exposed to FFE. On the other hand, women born before the 1980s, specifically before 1979, are less likely to have benefited from any of the programs, because they reached age 16 or older (i.e., at least 1 year older than the expected age at grade 10) before any of the programs were introduced. Therefore a woman's year of birth is the second IV.

To control for omitted variables at the community level specifically, a set of dummy variables of subdistricts is introduced in each model. The subdistricts where women received their educations, however, may be measured with errors in this study, as information on natal or childhood/adolescence subdistricts was not collected by the BDHS 2007; the data were matched based on where women were located at the time of the interview. The only relevant measure in assessing the potential measurement error is the duration in years lived in the current place, while the geographic boundary of "current place" was not defined to the interviewees. Specifically, 83.7% of women in the study sample have ever migrated, and about 50% of women who have ever migrated did so at age 16 or later, which is approximately the average age at first marriage (15.4) in the study sample. This may reflect migration upon marriage, because women in Bangladesh often move to their husbands' households upon marriage (Agarwal 1994).

The measurement error, if any, could be problematic in two ways, depending on the structure of the measurement error. The first potential problem is underestimation of the relation between educational attainment and the education programs, which results from a random measurement error. The underestimation in turn could invoke the weak instrument problem. We performed a partial-F test to assess the significance of the set of IVs; the results are presented in Table 4.

The second potential problem is the endogeneity of the education programs in estimating their effects on education and reproductive health, which result from a systematic measurement error (Duflo 2001). A systematic measurement error could arise from selective migration (Strauss and Thomas 1995), when the migration decision is a function of education and the destination of migration is based on factors related to the education programs and adolescent reproductive health (Cochrane 1979). For instance, women with higher educational attainment may be more likely to migrate to communities with a better set of characteristics, such as more schools and health facilities, and adolescent reproductive health may be affected by access to health facilities, which may be correlated with the number of schools.

To address this potential problem, we employed the subdistrict as the unit of observation for the number of schools, based on previous studies reporting that the majority of marriages take place within the natal subdistricts in Bangladesh (Aziz 1979; Islam 1974; Kabeer 1985). In this situation, the number of schools in the resident subdistrict reflects that of the natal subdistrict. All the women in the study sample were born before any of the programs were introduced, which implies that the number of schools in the natal subdistrict is not endogenous (Duflo 2001). Also the set of dummies of subdistricts introduced in each model rids it of any time-invariant effect of unobserved factors at the subdistrict level. We performed the test of over-identifying restriction for each of the models; results are presented in Table 8. Overall, we are assured that the set of IVs is valid in terms of both its strength of correlation with female education and its collective exogeneity. This suggests that the measurement error in subdistricts, if any, does not pose a significant problem in this study.

The identification assumption is supported by preliminary evidence. Columns 1–4 of Table 3 show the average highest grade achieved stratified by quintiles of the number of schools and a woman's year of birth. They suggest that there is a substantial increase in the average highest grade achieved for women born in the 1980s, as hypothesized. Likewise, the average highest grade achieved is higher where there are more secondary schools, which again supports the hypothesis that women in subdistricts with more schools have a higher educational attainment on average.

Columns 5–16 of Table 3 present the average age at first marriage, the proportion of women who had their first live birth by age 16, and the average number of live births by age 20 stratified by quintiles of the number of schools and a woman's year of birth. While the pattern of reproductive health outcomes across the quintiles of the number of schools is less clear compared to that of female education in general, the higher quintiles are associated with a higher age at marriage, a smaller proportion of having the first live birth by age 16, and fewer live births by age 20 when compared to the middle quintiles. The lowest quintile, however, exhibits favorable outcomes compared to the middle quintiles. The associations between a woman's year of birth and the three reproductive health outcomes are more straightforward. As hypothesized, women born after 1982 have the highest average age at first marriage, the smallest proportion of first live birth by age 16, and the smallest number of live births by age 20, followed by women born between 1979 and 1982 and then by women born before 1979.

In the next section, based on the supportive preliminary evidence, we apply a regression framework to estimate the causal effects of female education and the education programs on the three adolescent reproductive health outcomes. Similar methods are applied by Duflo (2001) and Breierova and Duflo (2004).

Table 3. Female educational attainment and reproductive health outcomes by a woman's year of birth and the number of schools

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
	Highest grade achieved				Average age at first marriage				Proportion of women having the first live birth by age 16				Number of live births by age 20			
Quintile of number of schools	Year of birth				Year of birth				Year of birth				Year of birth			
	Before 1979	1979-1982	After 1982	Total	Before 1979	1979-1982	After 1982	Total	Before 1979	1979-1982	After 1982	Total	Before 1982	1982-1985	After 1985	Total
Lowest	3.16	5.05	6.24	3.42	15.38	15.87	16.08	15.44	0.33	0.35	0.18	0.33	1.12	1.04	0.81	1.11
	(.204)	(.404)	(.462)	(.205)	(.122)	(.252)	(.295)	(.121)	(.017)	(.039)	(.060)	(.016)	(.030)	(.076)	(.131)	(.031)
Low	3.47	6.05	6.68	3.95	15.27	16.09	15.74	15.40	0.37	0.27	0.33	0.36	1.15	0.94	1.01	1.12
	(.218)	(.489)	(.436)	(.224)	(.140)	(.229)	(.401)	(.128)	(.017)	(.044)	(.075)	(.017)	(.038)	(.082)	(.139)	(.035)
Middle	3.62	6.09	5.89	4.19	15.19	15.98	15.41	15.34	0.38	0.36	0.32	0.37	1.16	0.98	1.00	1.12
	(.218)	(.464)	(.402)	(.225)	(.123)	(.265)	(.214)	(.115)	(.018)	(.039)	(.043)	(.016)	(.038)	(.077)	(.062)	(.032)
High	4.26	5.21	6.24	4.70	15.18	15.63	15.82	15.36	0.40	0.37	0.27	0.38	1.17	1.07	0.89	1.12
	(.285)	(.331)	(.484)	(.253)	(.178)	(.174)	(.315)	(.151)	(.025)	(.029)	(.046)	(.021)	(.047)	(.046)	(.094)	(.039)
Highest	4.32	6.20	6.76	5.41	15.21	15.60	15.60	15.41	0.36	0.36	0.37	0.37	1.10	1.02	1.00	1.05
	(.289)	(.274)	(.303)	(.224)	(.172)	(.168)	(.164)	(.122)	(.031)	(.031)	(.039)	(.021)	(.050)	(.057)	(.053)	(.035)
Total	3.67	5.77	6.44	4.31	15.26	15.76	15.67	15.39	0.37	0.35	0.32	0.36	1.14	1.02	0.96	1.10
	(.118)	(.171)	(.195)	(.113)	(.074)	(.096)	(.123)	(.068)	(.010)	(.016)	(.024)	(.009)	(.020)	(.029)	(.039)	(.018)

Note: Standard errors are presented in parentheses.

2. Reduced Form Results: Effect of Education Programs

Effect of Education Programs on Female Education

Two variables, the number of schools and a woman's year of birth, are used as measures of school accessibility and exposure to the financial incentives to estimate the effect of the education programs on female educational attainment. To include the number of schools as a measure of the effect of school accessibility, we assume that the difference in educational attainment across the number of schools is due to different levels of school accessibility across subdistricts. The assumption is violated if there is any unobserved time-varying variable correlated with the number of schools specifically at the subdistrict level. This suggests running a model that includes interaction terms between subdistrict and birth cohort dummies. Due to the limited sample size, however, we are unable to fit the full set of interaction dummies. Instead, we use an available indicator of socioeconomic development at the subdistrict level, namely, the female attendance rate of ages 5–29, obtained from the 1981 population census, and interact that rate with birth cohort dummies. We assess the significance of the interaction terms in the following model to test the hypothesis:

$$(1) \qquad E_{ijt} = \alpha + \boldsymbol{\beta}'\mathbf{B} + \delta P_{jt'} + \varphi'\mathbf{u} + \boldsymbol{\gamma}'\mathbf{I} + \varepsilon_{ijt},$$

where E_{ijt} is the highest grade achieved by woman i in subdistrict j born in year t, $\mathbf{B}$ is a vector of dummies of woman's year of birth, $\mathbf{u}$ is a vector of dummies of subdistrict, $P_{jt'}$ is the number of schools in subdistrict j in year $t+11$, $\mathbf{I}$ is a vector of interactions between the female attendance rate of subdistrict j in 1981 and birth cohort dummies, and ε_{ijt} is the disturbance term. Specifically, we are interested in the collective significance of $\boldsymbol{\gamma}$, the coefficients of the interaction terms.

The results are presented in column 1 of Table 4. None of the interaction terms is significant at the 5% level. While the model captures only limited characteristics at the subdistrict level, it is reassuring that there is no time-varying effect of a major socioeconomic development indicator, which is most likely to be correlated with female educational attainment.

Next, to include a woman's year of birth as a measure of the effect of the financial incentives, we assume that differences in educational attainment across cohorts are due to different levels of exposure to the financial incentives provided by the education programs. The assumption is violated if there is any systematic difference across cohorts that affects an individual's schooling decision. We examine the extent to which this assumption is supported by assessing educational attainment by birth cohorts. Because women who had reached age 16 or older in 1994 had left grade 10 before any of the programs were introduced, they are least likely to have benefited from any of the financial incentives. If female educational attainment differs significantly within this group of women, it may imply that there is a significant cohort effect besides exposure to the financial incentives, in which case the estimated effect of the financial incentives may be biased. This suggests running the following model:

$$(2) \qquad E_{ijt} = \alpha + \boldsymbol{\beta}'\mathbf{B} + \delta P_{jt'} + \varphi'\mathbf{u} + \varepsilon_{ijt}.$$

We are interested in $\boldsymbol{\beta}$, the coefficients of the vector of dummies of woman's year of birth, especially for women 16 or older in 1994, that is, for $t \leq 1978$.

Column 2 of Table 4 presents the results. The coefficients of birth cohorts are insignificant for women 19 or older in 1994, as hypothesized. However, the coefficients of birth cohorts for women 16, 17, or 18 in 1994 are significant, which contradicts our assumption. This may reflect exposure to the financial incentives due to grade repetition or delayed entry into school. Indeed, the reported repetition rate among girls was 9.6% for grades 1–5 and ranged from 6.5% to 18.0% for grades 6–10 in 2005 (BANBEIS 2010b; BANBEIS 2010c). In addition, about 9.4% of girls in the first grade were 7 or older in 2004 (Ahmed et al. 2007), which is substantially older than the expected age 6 in the first grade. Although corresponding figures for the 1980s are not available, it could be argued that girls older than expected were exposed to the financial incentives due to grade repetition or delayed entry. On the other hand, coefficients of birth cohorts for women 15 or younger in 1994 are significantly positive, as expected. Overall, the results support our assumptions that there is no substantial difference in educational attainment across birth cohorts among women who are least likely to have been exposed to the financial incentives and that educational attainment gradually increases for younger women who are likely to have been exposed to the financial incentives.

The results obtained from models (1) and (2) indicate that both the number of schools and a woman's year of birth are unlikely to be confounded by omitted variables. This suggests running the following model in estimating the effect of the education programs on a woman's highest grade achieved:

$$(3) \qquad E_{ijt} = \alpha + \beta_1 B_{1t} + \beta_2 B_{2t} + \delta P_{jt'} + \varphi' \mathbf{u} + \varepsilon_{ijt},$$

where B_{1t} and B_{2t} are dummies of woman's year of birth. The dummies indicate two cohort groups, women born between 1979 and 1982 and those born after 1982, respectively, so that they capture the effects of partial and full exposures relative to no exposure to the financial incentives. While some of the older women may have benefited from the financial incentives due to grade repetition or delayed entry into school, as shown in the previous analysis, they may differ in unobserved characteristics from women of the same birth cohorts who completed education at the expected age. Therefore we categorize birth cohorts by expected exposure to the financial incentives without any grade repetition or delayed entry into school.

The results presented in Table 4 suggest that a one-school increase per 10,000 population of ages 10–17 significantly increases the highest grade achieved by .165 years. Likewise, partial and full exposures to the financial incentives significantly increase the highest grade achieved by 1.666 and 2.652 years, respectively.

Table 4. Reduced form estimates of the effect of the education programs on female education

	Model (1)			Model (2)			Model (3)		
	Coefficients		(SE)	Coefficients		(SE)	Coefficients		(SE)
Number of schools	0.079		(.049)	0.095	*	(.049)	0.165	***	(.048)
Age in 1994									
12 to 15	-			-			1.666	***	(.147)
11 or younger	-			-			2.652	***	(.139)
7	4.856	***	(1.070)	3.382	***	(.346)	-		
8	5.080	***	(1.145)	3.251	***	(.361)	-		
9	3.717	**	(1.074)	3.031	***	(.363)	-		
10	4.137	***	(1.163)	2.776	***	(.373)	-		
11	3.320	**	(1.182)	2.682	***	(.372)	-		
12	2.560	*	(1.125)	2.278	***	(.369)	-		
13	0.867		(1.138)	1.914	***	(.375)	-		
14	2.810	*	(1.130)	2.090	***	(.395)	-		
15	1.998		(1.119)	1.913	***	(.381)	-		
16	2.335		(1.197)	1.626	***	(.420)	-		
17	1.417		(1.186)	1.411	***	(.391)	-		
18	1.331		(1.060)	0.932	*	(.402)	-		
19	1.772		(1.136)	0.723		(.394)	-		
20	2.586		(1.340)	0.688		(.407)	-		
21	-0.412		(1.116)	0.322		(.390)	-		
22	1.830		(1.226)	0.687		(.392)	-		
23	1.252		(1.157)	-0.025		(.373)	-		
24	1.841		(1.072)	Reference			-		
25	1.047		(1.441)	0.391		(.421)	-		
26	0.115		(1.126)	-0.207		(.382)	-		
27	0.589		(1.151)	-0.712		(.366)	-		
28	0.222		(1.107)	-0.577		(.370)	-		
29	1.429		(1.154)	-0.364		(.392)	-		
30	0.144		(1.218)	-0.221		(.405)	-		
31	Reference			-0.405		(.420)	-		
Constant	-0.157		(1.261)	0.171		(1.209)	-0.152		(1.186)
Year of birth*female attendance rate	Yes			No			No		
F-statistics[a]	1.33			6.13			10.82		
Adjusted R^2	0.238			0.234			0.219		
N	6,930			6,930			6,930		

Note: Subdistricts of residence are controlled for in all the models.

[a] F-tests assess the collective significance of the interaction terms, the year of birth (1978 or earlier), and all the independent variables for models (1), (2), and (3), respectively.

* p<.05; ** p<.01; *** p<.001

Effect of Education Programs on Adolescent Reproductive Health

The effect of the education programs on adolescent reproductive health as measured by the number of live births by age 20 and the probabilities of first marriage by age 15 and first live birth by age 16 can be assessed in the same manner. To assess omitted time-varying variables at the subdistrict level, we again examine the coefficients of the interactions between birth cohorts and the female attendance rate in the following model:

(4) $$Y_{hijt} = \alpha_h + \boldsymbol{\beta}'_h \mathbf{B} + \delta_h P_{jt'} + \varphi'_h \mathbf{u} + \boldsymbol{\gamma}'_h \mathbf{I} + \varepsilon_{hijt}, \quad h = 1,2,3,$$

where Y_{1ij} and Y_{2ij} are binary variables and indicate whether woman i in subdistrict of residence j was married by age 15 and whether she had her first live birth by age 16, respectively. Y_{3ij} is the measure of adolescent fertility and is the number of live births by age 20. A linear probability model is applied to the first and second equations.

The results are presented in column 1 of Tables 5, 6, and 7. Overall, the interaction terms are jointly insignificant in each of the three equations at the 5% level. Again, it is reassuring that there is no time-varying effect of a major socioeconomic development indicator, which is likely to be correlated with reproductive health.

Next, to assess if there is any cohort effect among women who are not exposed or who are least exposed to the financial incentives, we examine the coefficients of cohort dummies in the following model:

(5) $$Y_{hijt} = \alpha_h + \boldsymbol{\beta}'_h \mathbf{B} + \delta_h P_{jt'} + \varphi'_h \mathbf{u} + \varepsilon_{hijt}, \quad h = 1,2,3.$$

Again, we are interested in $\boldsymbol{\beta}$, the coefficients of the vector of dummies of woman's year of birth, especially for women 16 or older in 1994, that is, for $t \le 1978$.

The results are presented in column 2 of Tables 5, 6, and 7. The cohort fixed effects for women 16 or older in 1994 are insignificant except for the birth year 1972 (age 22) in the model regressing the probability of first marriage by age 15.

The results again indicate that both the number of schools and a woman's year of birth are unlikely to be confounded by omitted variables. This suggests running the following model in estimating the effect of the education programs on the three adolescent reproductive health outcomes:

(6) $$Y_{hijt} = \alpha_h + \beta_{h1} B_{1t} + \beta_{h2} B_{2t} + \delta_h P_{jt'} + \varphi'_h \mathbf{u} + \varepsilon_{hijt}, \quad h = 1,2,3.$$

The results are presented in column 3 of Tables 5, 6, and 7. It is suggested that a one-school increase per 10,000 population of ages 10–17 significantly reduces the probability of first marriage by age 15 by .014 and the number of live births by age 20 by .027 births. However, the number of schools does not have a significant effect on the probability of first live birth by age 16. Partial exposure to the financial incentives, as captured by a woman's birth between 1979 and 1982, significantly reduces the probability of first marriage by age 15 by .108 and the number of live births by age 20 by .118 births. However, its effect on the probability of first live birth by age 16 is insignificant. Full exposure to the financial incentives, as captured by a woman's birth after 1983, has a larger effect on average than partial exposure. It reduces the probability of first marriage by age 15 by .103, the probability of first live birth by age 16 by .057, and the number of live births by age 20 by .184 births.

Table 5. Reduced form estimates of the effect of the education programs on the probability of first marriage by age 15

	Model (4)			Model (5)			Model (6)		
	Coefficients		(SE)	Coefficients		(SE)	Coefficients		(SE)
Number of schools	-0.010		(.006)	-0.010		(.006)	-0.014	*	(.006)
Age in 1994									
12 to 15	-			-			-0.108	***	(.017)
11 or younger	-			-			-0.103	***	(.027)
7	-0.290	*	(.140)	-0.150	**	(.046)	-		
8	-0.259		(.142)	-0.148	**	(.046)	-		
9	-0.227		(.136)	-0.176	***	(.045)	-		
10	-0.204		(.142)	-0.150	**	(.046)	-		
11	-0.239		(.141)	-0.123	**	(.046)	-		
12	-0.170		(.140)	-0.155	**	(.046)	-		
13	-0.218		(.139)	-0.111	*	(.045)	-		
14	-0.220		(.137)	-0.109	*	(.047)	-		
15	-0.096		(.135)	-0.096	*	(.046)	-		
16	-0.229		(.134)	-0.052		(.046)	-		
17	-0.069		(.138)	-0.057		(.047)	-		
18	-0.141		(.134)	-0.027		(.047)	-		
19	-0.150		(.136)	-0.039		(.048)	-		
20	-0.033		(.146)	-0.065		(.047)	-		
21	-0.117		(.141)	-0.025		(.049)	-		
22	-0.241		(.146)	-0.112	*	(.048)	-		
23	-0.168		(.140)	-0.058		(.045)	-		
24	-0.072		(.142)	Reference			-		
25	-0.134		(.168)	-0.025		(.050)	-		
26	-0.004		(.134)	0.022		(.049)	-		
27	-0.274		(.160)	0.001		(.050)	-		
28	0.088		(.144)	0.038		(.048)	-		
29	-0.134		(.145)	0.026		(.050)	-		
30	-0.132		(.150)	0.015		(.053)	-		
31	Reference			0.008		(.053)	-		
Constant	1.181	***	(.105)	1.146	***	(.094)	1.142	***	(.087)
Year of birth*female attendance rate	Yes			No			No		
F-statistics[a]	0.72			1.39			7.99		
Adjusted R^2	0.136			0.134			0.128		
N	6,930			6,930			6,930		

Note: Subdistricts of residence are controlled for in all the models.
[a] F-tests assess the collective significance of the interaction terms, the year of birth (1978 or earlier), and all the independent variables for models (1), (2), and (3), respectively.
* p<.05; ** p<.01; *** p<.001

Table 6. Reduced form estimates of the effect of the education programs on the probability of first live birth by age 16

	Model (4)			Model (5)			Model (6)		
	Coefficients		(SE)	Coefficients		(SE)	Coefficients		(SE)
Number of schools	-0.004		(.006)	-0.004		(.006)	-0.002		(.006)
Age in 1994									
12 to 15	-			-			-0.016		(.017)
11 or younger	-			-			-0.057	*	(.026)
7	-0.182		(.150)	-0.037		(.048)	-		
8	0.011		(.145)	-0.003		(.048)	-		
9	0.003		(.140)	-0.028		(.047)	-		
10	-0.036		(.149)	0.007		(.048)	-		
11	0.002		(.152)	0.021		(.049)	-		
12	0.123		(.151)	0.024		(.049)	-		
13	0.063		(.150)	0.040		(.049)	-		
14	-0.028		(.146)	0.026		(.049)	-		
15	0.022		(.145)	0.019		(.049)	-		
16	-0.006		(.148)	-0.007		(.050)	-		
17	-0.059		(.148)	-0.046		(.050)	-		
18	-0.008		(.144)	0.038		(.051)	-		
19	0.010		(.146)	0.032		(.051)	-		
20	-0.005		(.156)	0.012		(.052)	-		
21	0.110		(.156)	0.069		(.053)	-		
22	-0.084		(.146)	-0.079		(.049)	-		
23	-0.119		(.149)	0.031		(.049)	-		
24	-0.044		(.150)	Reference			-		
25	-0.088		(.165)	-0.041		(.052)	-		
26	0.055		(.148)	0.026		(.055)	-		
27	0.008		(.172)	0.096		(.055)	-		
28	0.380	*	(.163)	0.078		(.056)	-		
29	-0.073		(.163)	-0.052		(.054)	-		
30	-0.199		(.168)	-0.032		(.057)	-		
31	Reference			0.015		(.057)	-		
Constant	0.466	**	(.155)	0.446	**	(.144)	0.451	**	(.139)
Year of birth*female attendance rate	Yes			No			No		
F-statistics[a]	0.98			1.91			15.54		
Adjusted R^2	0.088			0.084			0.079		
N	6,930			6,930			6,930		

Note: Subdistricts of residence are controlled for in all the models.
[a] F-tests assess the collective significance of the interaction terms, the year of birth (1978 or earlier), and all the independent variables for models (1), (2), and (3), respectively.
* p<.05; ** p<.01; *** p<.001

Table 7. Reduced form estimates of the effect of the education programs on the number of live births by age 20

	Model (4)			Model (5)			Model (6)		
	Coefficients		(SE)	Coefficients		(SE)	Coefficients		(SE)
Number of schools	-0.023	*	(.011)	-0.023	*	(.011)	-0.027	*	(.011)
Age in 1994									
12 to 15	-			-			-0.118	***	(.031)
11 or younger	-			-			-0.184	***	(.044)
7	-0.255		(.292)	-0.210	*	(.086)	-		
8	0.034		(.290)	-0.089		(.086)	-		
9	-0.027		(.279)	-0.236	**	(.084)	-		
10	-0.066		(.289)	-0.109		(.087)	-		
11	0.042		(.294)	-0.089		(.091)	-		
12	0.178		(.298)	-0.084		(.088)	-		
13	0.208		(.300)	-0.040		(.091)	-		
14	0.086		(.290)	-0.070		(.092)	-		
15	0.232		(.293)	-0.054		(.089)	-		
16	0.014		(.304)	-0.127		(.095)	-		
17	-0.236		(.295)	-0.154		(.090)	-		
18	0.059		(.288)	0.027		(.096)	-		
19	0.172		(.303)	-0.007		(.095)	-		
20	0.378		(.317)	-0.001		(.099)	-		
21	0.296		(.321)	0.055		(.102)	-		
22	-0.260		(.313)	-0.142		(.095)	-		
23	-0.213		(.311)	0.031		(.095)	-		
24	0.121		(.303)	Reference			-		
25	-0.034		(.346)	-0.039		(.103)	-		
26	0.233		(.327)	0.105		(.116)	-		
27	0.003		(.350)	0.062		(.108)	-		
28	0.718	*	(.316)	0.185		(.102)	-		
29	0.083		(.330)	0.000		(.104)	-		
30	-0.030		(.324)	-0.009		(.113)	-		
31	Reference			0.011		(.109)	-		
Constant	1.826	***	(.267)	1.857	***	(.240)	1.902	***	(.229)
Year of birth*female attendance rate	Yes			No			No		
F-statistics[a]	1.16			1.68			2.82		
Adjusted R^2	0.092			0.088			0.082		
N	6,930			6,930			6,930		

Note: Subdistricts of residence are controlled for in all the models.
[a] F-tests assess the collective significance of the interaction terms, the year of birth (1978 or earlier), and all the independent variables for models (1), (2), and (3), respectively.
* p<.05; ** p<.01; *** p<.001

3. Instrumental Variable Method Results: Effect of Education

Effect of Education on Adolescent Reproductive Health

We employ the 2SLS method to address the potential endogeneity of female education and to estimate the causal effect of female education on the adolescent reproductive health outcomes. Using model (3) as the first-stage equation to obtain the estimated highest grade achieved of woman i ($\hat{E}_{ijt}$), the second-stage equations of the 2SLS model are specified to estimate the effect of female education on the three adolescent reproductive outcomes as:

$$(7) \qquad Y_{hijt} = \lambda_h + \theta_h \hat{E}_{ijt} + \boldsymbol{\tau}_h^{'}\mathbf{u} + \varepsilon_{hijt}, \quad h = 1,2,3.$$

We are interested in θ_h, the coefficients of estimated highest grade achieved. Again, a linear probability model is applied to the first and second equations.

The results are presented in columns 1, 3, and 5 of Table 8. The test of over-identifying restriction for each of the three equations does not reject the collective orthogonality of the IVs (p=.343, p=.480, and p=.409, respectively). It is suggested that a one-year increase in the highest grade achieved significantly reduces the probability of first marriage by age 15 by .050, the probability of first live birth by age 16 by .013, and the number of live births by age 20 by .072 births.

The estimates are consistent with the reduced form results. Note that increasing the number of schools by one increases the highest grade achieved by .165 years. Then the direct effect of the number of schools on the probability of first marriage by age 15, the probability of first live birth by age 16, and the number of live births by age 20 should be -.008 (=.165*-.05), -.002(=.165*-.013), and -.012(=.165*-.072), respectively. These estimates are approximately equal to the results shown in Tables 5-7. Also partial and full exposures to the financial incentives are estimated to increase the highest grade achieved by 1.666 and 2.652 years, respectively. Then the direct effect of partial exposure on the three reproductive health outcomes should be -.083 (=1.666*-.05), -.022(=1.666*-.013), and -.120(=1.666*-.072), respectively. Similarly, the direct effect of full exposure should be -.133(=2.652*-.05), -.035(=2.652*-.013), and -.191(=2.652*-.072), respectively. Again, these estimates are approximately equal to the results shown in Tables 5-7.

Difference between 2SLS and OLS Estimates

Finally, the estimated coefficients of female education are compared between 2SLS and OLS, the latter replacing $\hat{E}_{ijt}$ with E_{ijt} in model (7). The OLS estimates are presented in columns 2, 4, and 6 of Table 8. The Durbin-Wu-Hausman test suggests that the 2SLS estimates are significantly different from the OLS estimates in the first two equations, those regressing the probability of first marriage by age 15 and the probability of first live birth by age 16, but not in the equation regressing the number of live births by age 20. The 2SLS estimate is larger in the absolute term than the corresponding OLS estimate in the equation regressing the probability of first marriage by age 15 but is smaller in the equation regressing the probability of first live birth by age 16.

Table 8. 2SLS and OLS estimates of the effect of female education on reproductive health

	Probability of first marriage by age 15				Probability of first live birth by age 16				Number of live births by age 20			
	2SLS		OLS		2SLS		OLS		2SLS		OLS	
Education	-0.050	***	-0.040	***	-0.013	*	-0.025	***	-0.072	***	-0.063	***
	(.006)		(.002)		(.006)		(.002)		(.010)		(.003)	
Constant	1.087	***	1.055	***	0.451	***	0.483	***	1.738	***	1.712	***
	(.063)		(.004)		(.122)		(.004)		(.214)		(.009)	
R^2	0.213		0.220		0.111		0.120		0.152		0.154	
Over-identifying restriction												
Chi-square	2.14		-		1.47		-		1.79		-	
p-value	0.343		-		0.480		-		0.409		-	
Durbin-Wu-Hausman test												
Chi-square	4.83		-		5.48		-		1.24		-	
p-value	0.028		-		0.019		-		0.266		-	
N	6,930		6,930		6,930		6,930		6,930		6,930	

Note: Standard errors are presented in parentheses.
Note: Subdistricts of residence are controlled for in all the models.
* p<.05; ** p<.01; *** p<.001

4. Conclusion

In this paper we examined the causal effect of female education on adolescent fertility and the exposure factors as proximate determinants of adolescent fertility. We assumed that female education is endogenous due to unobserved variables at the individual and community levels. We employed IVs generated through the education programs to estimate the causal effect of female education on reproductive health outcomes and to understand the mechanisms through which female education influences fertility. Our finding suggests that female education significantly influences all of the adolescent reproductive health outcomes assessed. Specifically, a one-year increase in the highest grade achieved reduced significantly the probabilities of first marriage by age 15 and first live birth by age 16, on average by .050 and .013, respectively. Correspondingly, a one-year increase in the highest grade achieved reduced the number of live births by age 20 by .072 births. The set of specification tests supports our assumptions and yields a causal interpretation from these estimates. Female education therefore significantly delays exposure to the risk of pregnancy and reduces fertility during adolescence.

The difference between the 2SLS and OLS estimates varied across the adolescent reproductive health outcomes. We did not find a significant difference between 2SLS and OLS estimates for the effect of female education on the number of live births by age 20. However, the 2SLS estimates differed significantly from the corresponding OLS estimates for the effect on the exposure factors, namely the probabilities of first marriage by age 15 and first live birth by age 16, suggesting that female education is endogenous in the context of the exposure factors of fertility. While the 2SLS estimate for the effect of female education on the probability of first marriage by age 15 was significantly larger than the corresponding OLS estimate in the absolute term, it was significantly smaller for the effect on the probability of first live birth by age 16. These results therefore do not support the hypothesis that OLS estimates for the effect of female education are uniformly biased upward or downward in the context of adolescent reproductive health. This suggests that the magnitude or direction of the omitted variable bias cannot be determined as a priori knowledge. The conclusion is consistent with the study by Breierova and Duflo, finding that differences in 2SLS and OLS estimates for the effect of education vary across reproductive health outcomes of their interests (Breierova and Duflo 2004). Therefore, studies need to address the endogeneity of female education for each outcome of interests to examine the mechanisms through which female education influences adolescent fertility.

We also examined the causal effect of the education programs on female education. A one-school increase per 10,000 population of ages 10–17 when a woman was age 11 increased the highest grade achieved by .165 years. Likewise, partial and full exposures to the financial incentives, as captured by a woman's year of birth, significantly increased the highest grade achieved by 1.666 and 2.652 years, respectively. The finding suggests that the education programs have been effective in enhancing female educational attainment. Correspondingly, a one-school increase per 10,000 population of ages 10–17 significantly reduced the probability of first marriage by age 15 by .014 and the number of live births by age 20 by .027 births. However, the number of schools did not have a significant effect on the probability of first live birth by age 16. Partial exposure to the financial incentives significantly reduced the probability of first marriage by age 15 by .108 and the number of live births by age 20 by .118 births; its effect on the probability of first live birth by age 16 was insignificant. Full exposure to the financial incentives had a larger effect on average than partial exposure. Full exposure reduced the probabilities of first marriage by age 15 by .103, first live birth by age 16 by .057, and the number of live births by age 20 by .184 births. The results suggest that education programs can serve as a means to improve adolescent reproductive health by enhancing female educational attainment.

However, the study results require a cautious generalization to other settings or populations for several reasons. First, the effect of female education estimated in this study may not be generalized to other settings if the relationship between female education and reproductive health is not immune to differences in those contexts. In particular, socioeconomic characteristics at the societal level are suggested to influence the relationship between female education and reproductive health. Cochrane reported differences in the expected inverse relationship between female education and fertility in her review of studies in several countries (Cochrane 1979). She found that in general countries at the middle level of development exhibited the expected inverse relationship between female education and fertility, but she found it insignificant in a few countries of other levels of development. Although these studies are not entirely comparable to ours due to differences in research methods and potentially confounding factors, the findings call for a cautious generalization of our study results.

Why may the effect of female education vary across socioeconomic characteristics at the societal level? One potential explanation is a difference in the opportunity costs of childbearing and child rearing (Cleland 2009; Diamond et al. 1999). Female education is presumed to increase these opportunity costs because of enhanced female participation in society outside the home through formal employment and other activities (Diamond et al. 1999). When such opportunities are not readily available to women, educated women may not face increased opportunity costs, and higher education may not reduce fertility substantially. Bangladesh has experienced a rapid increase in formal employment opportunities generated especially by the growth of the garment manufacturing industry since the mid-1980s (Amin et al. 1998; Raynor and Wesson 2006). By 1995 approximately 2,400 registered factories provided formal employment to more than one million women (Amin et al. 1998). Many young women have gained opportunities to earn independent incomes.

On the other hand, in the traditional society of Bangladesh the institution of *purdah* (seclusion of women) is widespread (Amin 1996). Educated women may have less mobility, as enhanced social prestige and higher status are often accompanied by a greater degree of seclusion (Amin 1996). Their social prestige could be undermined by working outside the home, especially in rural areas (Amin 1996). Bangladesh therefore can be characterized by the coexistence of a rapid increase in job opportunities and traditional *purdah*, both of which potentially moderate the effect of female education. The effect of female education on reproductive health should be interpreted in the context unique to the country.

In addition, the estimated effect of female education on reproductive health may depend on the types of education programs introduced. This is true when the effect of female education on reproductive health varies among women. Consequently, the 2SLS estimates reflect the weighted average effect of education of women affected by the education programs (Imbens and Angrist 1994). For instance, it could be argued that the stipend provided at the secondary education level affects those who have completed a primary education and can afford the partial cost of a secondary education. The very poor, who are less likely to complete a primary education or to afford any costs associated with education, may not be affected by the stipend. On the other hand, the food rations may have enticed poor households to send their children to primary school. Multiple programs targeting different populations were introduced simultaneously in Bangladesh, complicating efforts to identify a group of women who have been affected by the programs since the 1990s.

Finally, this study does not identify the components of the programs that have been most effective in bringing about the improvements in female education or adolescent reproductive health due to a lack of variations in the components across the country. For instance, women born after 1982, who are presumed to have been exposed to more financial incentives than those born earlier, were found to have higher educational attainment. While this may indicate that larger financial incentives have increased female educational attainment, it could instead reflect a lagged response to the programs introduced previously. Because the program components have become more varied since the mid-2000s, further research on adolescent reproductive health that addresses these variations will provide insight into education program designs that effectively promote female education and reproductive health.

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