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Women's Education in the Fertility Transition: The Reversal of the Relationship Between Women's Education and Birth Spacing in Indonesia

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Abstract

It is generally thought that both the demand for children and the cost of fertility control are major forces in fertility decline. Most researchers find that family planning programs in developing countries, which lower the cost of fertility control, play a small role in the fertility transition relative to other economic factors that affect the demand for children. This paper examines one aspect of fertility, namely the second birth interval in Indonesia over the period 1970 to 1993. It is observed that higher female education is associated with a shorter birth interval among earlier cohorts, but with a longer birth interval among later cohorts. The finding is that changes in the effect of education on birth hazard over time are primarily driven by changes in the cost of fertility control rather than through changes in the demand for children. Hence, family planning programs can have a big impact on fertility in a relatively low-educated population. In addition, in the context of contraceptive technology, the result can be interpreted as evidence for the hypothesis that education enhances the ability to decipher new information.

Keywords

Birth Spacing, Female Education, Family Planning Programs and Indonesia.

JEL Classification Number: J13, J24.

Authors

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1 Introduction

Although it is generally understood that both the demand for children and the cost of fertility control are major forces in fertility decline, there has been a long debate in demography and economics literature about the relative importance of these two factors in explaining fertility decline in developing countries.[1] In an attempt to contribute to this debate, I examine women's education and birth spacing in Indonesia.

There are two points to be made concerning this debate. First, while it has been widely documented that increases in women's education have a negative effect on fertility in developing countries, there are few studies that identify the mechanisms through which education affects fertility. I investigate this issue adopting the framework suggested by Bulatao & Lee (1983) and Easterlin & Crimmins (1985). Specifically, I consider how education might affect fertility through the demand for children, the supply of children or the cost of fertility control.[2][3]

Second, a striking feature of the fertility decline in Indonesia is its relatively fast pace. The birth rate of Indonesia decreased from 42.3 per 1,000 persons over the 1965–1970 period to 28.0 per 1,000 persons over the 1985–1990 period. A comparable change

[1] See, e.g., Bulatao & Lee 1983, Easterlin & Crimmins 1985, and Gertler & Molyneaux 1994.

[2] The demand for children means the number of surviving children parents would want if fertility control were costless (Easterlin & Crimmins 1985). Theories of the demand for children include Becker & Lewis (1973) and Willis (1973). Becker & Lewis (1973) suggest that the parents' education lowers the price of quality of children in the framework of quantity-quality trade-off. Willis (1973) suggests that women's education lowers fertility through an increase in the opportunity cost of women's time where the production technology for children is time intensive relative to the technology for parents' standard of living. The supply of children means the number of surviving children a couple would have if they made no deliberate attempt to limit family size (Easterlin & Crimmins 1985). Education may improve the health condition of women and infants by providing knowledge with regard to food care, personal hygiene and so on, which will increase the natural fertility of the women and survival prospects of infants. The cost of fertility control combines monetary and psychological cost related to contraception (Easterlin & Crimmins 1985). Education may enhance the ability to adopt new contraceptives, which, in turn, lowers the cost of fertility control (Schultz, T.W. 1975).

[3] Easterlin & Crimmins (1985) decompose the educational effect on fertility into those three mechanisms, but their proximate determinants analysis imposes the exogeneity of proximate determinants, which is unlikely to hold. Appleton (1996) examines the effect of female education on fertility through age at cohabitation and duration of breastfeeding (supply of children).

in birth rate took more than 60 years in the U.S. and more than 100 years in England and Wales.[4] In general, fertility decline is composed of a decrease in lifetime births, a delay in the age of the first birth and an increase in birth spacing. I examine one aspect of fertility: birth spacing. To be more specific, a single birth interval is used as a lens to view the dynamics of the rapid fertility transition in Indonesia. This study looks at individual birth history over the period 1970 to 1993. Focusing on a major part of the fertility decline in Indonesia (1970-1993) will provide a better understanding of the transitional dynamics.[5]

There are two other insights that can be gained by examining birth spacing behavior. One is that birth spacing has an impact on population growth even when family size is held constant. That is, longer birth intervals decrease the population growth rate, given the same number of children. The other implication of birth spacing is on the health outcomes of infants, which has been widely studied in the demography literature. Most studies have found that longer birth intervals have positive effects on infants' health (Forste 1994 and Miller et. al. 1992). Given these consequences of birth spacing, analyzing the determinants of birth intervals has important policy implications.

The purpose of this paper is to examine the effect of education on the second birth interval (the time from first birth to second birth) and to identify whether education affects birth spacing through the demand for children, the supply of children, or the cost of fertility regulation. In particular, I consider the effects of primary education employing a logit hazard model to implement the empirical analysis.

Using the 1993 Indonesian Family Life Survey (IFLS 93) I find that women who are more educated tend to have shorter birth intervals than women with less education among earlier cohorts. Interestingly, however, the opposite is found among later cohorts. That is, more educated women are likely to have longer birth intervals than their less educated counterparts. The latter is consistent with a finding that higher female education is associated with longer second birth intervals by other studies including Newman & McCulloch (1984), Heckman, Hotz & Walker (1985), Tasiran (1995), and Johnson-Hanks (2003).[6] The reversal of the relationship between women's schooling and birth intervals presents a particularly interesting characteristic of the Indonesian population.

To explain the initial relationship between education and birth spacing, I focus on supply effects. As an important factor in determining the exposure to the risk of

[4]Crude Birth Rate (in selected years)

year	Indonesia	year	U.S.	year	England and Wales
1965-1970	42.3	1850	43.3	1800	38.3
1985-1990	28.0	1910-1914	29.1	1900	27.6

Note: Crude birth rate is defined as the number of live births per 1,000 population in a year. Data source: Indonesia (UN 2000), U.S. (Andorka 1978, p.110), England and Wales (Andorka 1978, p.114).

[5]Among the studies on the Indonesian fertility decline using group birth hazard, Pitt, Rosenzweig & Gibbons (1994) covers the period from 1980 to 1985, Gertler & Molyneaux (1994) examines the period from 1982 to 1987, and Gertler & Molyneaux (2000) analyzes the period from 1986 to 1996.

[6]Newman & McCulloch (1984) used data from the 1976 Costa Rica National Fertility Survey. Heckman, Hotz & Walker (1985) used the 1981 Swedish Fertility Survey. Tasiran (1995) used the 1981 Swedish Fertility Survey, the Swedish 1984 and 1988 Household Market and Non-Market Activities, and the 1985-1988 PSID, finding a negative effect of female education on the second birth interval only in the U.S. sample. Johnson-Hanks (2003) used the 1998 Cameroon DHS.

pregnancy, I examine the breastfeeding pattern of educational groups over the sample period. I observe that more educated women tend to breastfeed for a shorter duration in the 1970s, and that the difference between educational groups disappears in the 1990s. With evidence from other studies including Sigle (1998) that the effect of breastfeeding on birth interval becomes less significant after six months of birth interval, I present the pattern of breastfeeding as a main determinant of birth intervals among earlier cohorts.

Two hypotheses are proposed regarding the reversal in sign of the effect of education on the second birth interval over time. The first hypothesis is based on the demand for children. One common implication of dynamic fertility models (e.g., Wolpin (1984) and Newman (1988)) is that a couple with a steeper income profile over their life cycle has an incentive to space births more widely in the absence of perfect capital markets.[7] If the return to schooling in the labor market increases as the economy becomes more industrialized, and if schooling and experience are complements in the formation of human capital, more educated women have an incentive to have longer birth intervals in a modern economy. I test this possibility using the occupational composition of a local economy as a measure of the level of industrialization. Since the slope of the wage profile may vary across occupations, I also test if the educational effect on the birth hazard can be explained by the occupational choices of individuals.

The second hypothesis centers on the cost of fertility regulation. The Indonesian government implemented a family planning program in Java and Bali in 1971, and extended the program to the national level over the following decade. In addition, the government has intensified the program by introducing subsidies for contraceptives and visits by field workers. With a family planning program in place, if more educated women are better able to learn new contraceptive technology as suggested by Schultz (1975), they will space births further apart because of the better access to modern birth control methods. Therefore, I test whether the family planning program has differential effects on educational groups.

Empirical results suggest that, although there is some evidence for the effects of industrialization and occupation on birth spacing, they are not significantly different across educational groups. On the other hand, the interaction between family planning programs and education turns out to be the key channel in which education affects birth spacing. The point estimate of the interaction between female education and family planning programs is significant at the 10 percent level, and 77 percent of the total change of the marginal effect of primary education on birth hazard over time can be explained by the differential effect of the family planning program on educational groups. The results imply that changes in the effect of education on birth spacing in Indonesia over the period from 1974 to 1990 were primarily driven by changes in the cost of fertility control rather than through changes in the demand for children. This, in turn, suggests that family planning programs can have a big impact on fertility in a relatively low-educated population. In addition, the result can be interpreted as evidence for the hypothesis that education enhances the ability to decipher new information in the context of contraceptive technology as discussed by Schultz (1975) and Rosenzweig (1994).

[7]The implication can be generalized when the capital markets are perfect and the market interest rate exceeds the time discount rate as demonstrated in Heckman & Willis (1976).

The rest of the paper is organized as follows. The next section describes the relationship between women's education and birth spacing in the fertility transition in Indonesia over the period 1970 to 1993. Section 3 explains in greater detail the two hypotheses for the reversal of the relationship. Section 4 provides the empirical results. Section 5 concludes.

2 Education and Birth Spacing in Indonesia

2.1 Descriptive Statistics

The data analyzed is found in the 1993 Indonesian Family Life Survey (IFLS 93), which provides data at the individual and household level on fertility, health, education, migration, and employment, as well as data at the community level on health facilities, schools and other community characteristics. The IFLS 93 consists of a sample of 7,224 households spread across 13 provinces on the islands of Java, Sumatra, Bali, West Nusa Tenggara, Kalimantan, and Sulawesi. The sample covers approximately 83 percent of the Indonesian population and much of its heterogeneity. One of the strengths of the IFLS 93 is the fact that it has extensive and reliable retrospective data at each level.[8] This is especially important for the purpose of this study because the fact to be explained involves the change of individual behavior over three decades.

There are three reasons why I choose the second birth interval as a measure of birth spacing. First, there is likely to be occurrence-dependence within birth intervals of an individual, as Heckman & Walker (1987) concluded using the goodness of fit tests in their study on Hutterite data.[9] Second, in a society like Indonesia where people expect a couple to have a baby soon after the marriage, the first birth interval is likely to be governed by incentives other than those based on economics considerations. In addition, first birth intervals are not a good measure due to arranged marriages or pregnancies before marriage. Finally, given the two reasons above, the second birth interval leaves the least amount of sample selection (married women without any children are discarded). As will be discussed next, the selection process appears to be minor from the sample used in this study.

Since I am examining the second birth interval, there is a potential sample selection issue due to removing married women without any children at the time of the survey year. There are 4,980 ever-married women in the pregnancy history section in the IFLS 93. After constructing the basic characteristics, there are 4,776 observations available. When limiting the sample to women who have at least one birth, this leaves 4,553 observations. The comparison of the total sample of married women and the subsample of women with at least one birth is presented in Table 1. Both samples have similar distributions regarding age, schooling, and age at marriage. The full sample has the average age of 34.1 years and the average schooling of 5.04 years, whereas women in the subsample are 34.5 years old and have 5.00 years of schooling on average. Age at first marriage is 18.09 years for the total sample and 18.02 years for the subsample.

[8]The IFLS 93 has a feature of cross-checking. The same question was asked to the respondent repeatedly (at individual level) and to different people (at community level).

[9]The Hutterites are a Mennonite population living in the Upper Midwest in the U.S. and Canada. They are considered a benchmark Western natural fertility population.

The percentage of women with complete primary education is 26.2 percent for the total sample and 25.8 percent for the subsample. The age distributions of the both samples are also similar to each other. Therefore, the selection from all the married women to the married women with at least one child seems to be small in terms of age and education.

Table 2 provides descriptive statistics of schooling groups and birth cohorts groups. On average, women who are more educated marry later and have their first baby later than the less educated women. The average age at marriage of women with complete primary education is 21.0 years whereas it is 16.3 years for women with less than primary education. The average age at first birth is 22.2 years for the primary education group and 17.0 years for less-than-primary education group. Birth cohorts do not differ substantially in terms of average age at marriage and age at first birth. However, there was an increase of average years of schooling over time. The early cohort (women in their 40s in 1993) completed on average 4.4 years of schooling, while the later cohort (women in their 20s in 1993) finished 5.8 years of schooling on average. Similarly, the percentage of women with complete primary education among each cohort increased from 22.6 percent for the early cohort to 31.7 percent for the later cohort.

A survival function analysis is used in order to deal with the right censoring due to open birth intervals. Figure 1 shows Kaplan-Meier survival function estimates for an event of second birth since first birth by education and by birth-cohort. In the first panel in Figure 1, which describes women in their 40s at the time of the survey, the survival function estimate is lower for the group of women with complete primary education than it is for the group with less-than-primary education. This implies that, among earlier cohorts, women who are more educated tend to have shorter second birth intervals. Interestingly, the opposite is seen in the third panel which covers women in their 20s in 1993. That is, the survival function estimate for women with complete primary education is higher than that of women with less than primary education. This means that more educated women are likely to have longer second birth intervals than less educated women among later cohorts.

The survival function estimates of second birth interval could be affected by the occurrence of birth stopping. That is, the survival function estimate becomes higher as there are more birth-stoppers. This case is demonstrated in Figure 2 using a hypothetical data in which there are 30 percent of birth stoppers in the sample. In this case, the survival function estimate for all women is a weighted average of two survival function estimates: one for non-stoppers and one for stoppers. Since the survival function curve for all women converges to the ratio of birth stoppers, I can calculate the survival function estimate for non-stoppers using the stoppers ratio. Hence, if there is a greater share of birth stoppers among the more educated women of the later cohorts, the third panel in Figure 1 is misleading. In order to check this possibility, I define the birth stoppers as women who have not given birth five years after their first birth and construct the survival function estimate for non-stoppers. As presented in Figure 3, taking into account the occurrence of birth stoppers does not change the previously observed relationship of education and birth intervals. The result does not change when I define birth-stoppers as women who have not had a second baby seven or ten years after their first birth. Hence, birth stopping behavior seems to be orthogonal to educational groups.

Table 1: Sample Statistics I

Variable	Ever Married Women	Women with at least one birth
No. Obs	4,776	4,553
Mean		
age	34.14	34.45
schooling	5.04	5.00
age at 1st marriage	18.09	18.02
age at 1st birth		19.96
Schooling Distribution		
less than primary	73.8	74.2
complete primary	26.2	25.8
Total	100.0	100.0
Age Distribution		
50s	1.3	1.3
40s	26.6	27.3
30s	41.4	42.3
20s	29.1	28.3
10s	1.7	0.9
Total	100.0	100.0

Notes: The data used are the 1993 Indonesian Family Life Survey.

Table 2: Sample Statistics II

	Age at marriage	Age at 1st birth		
Less than Primary	17.0	19.2		
Complete Primary	21.0	22.2		
	Age at marriage	Age at 1st birth	Schooling	% of Primary Education
40s	18.1	20.5	4.41	22.6
30s	18.2	20.1	4.87	24.1
20s	17.7	19.2	5.77	31.7

Notes: The data used are the 1993 Indonesian Family Life Survey.

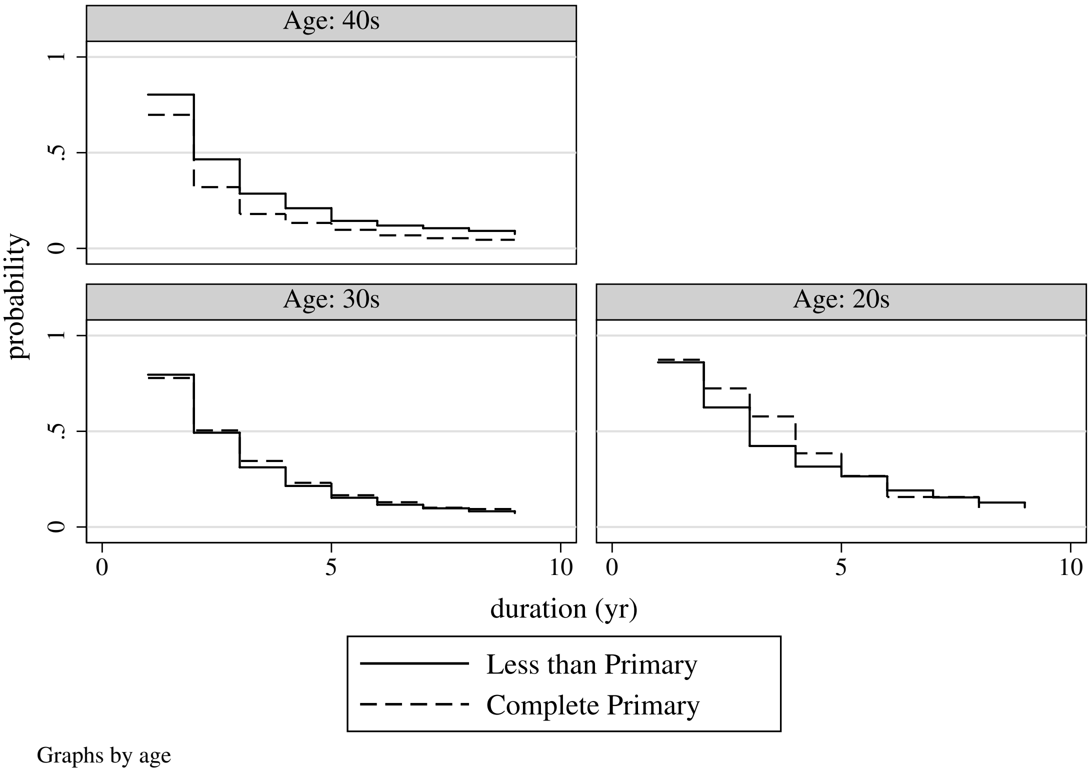

Figure 1: Survival Function Estimates of Second Birth Hazard

Notes: The data used are the 1993 Indonesian Family Life Survey.

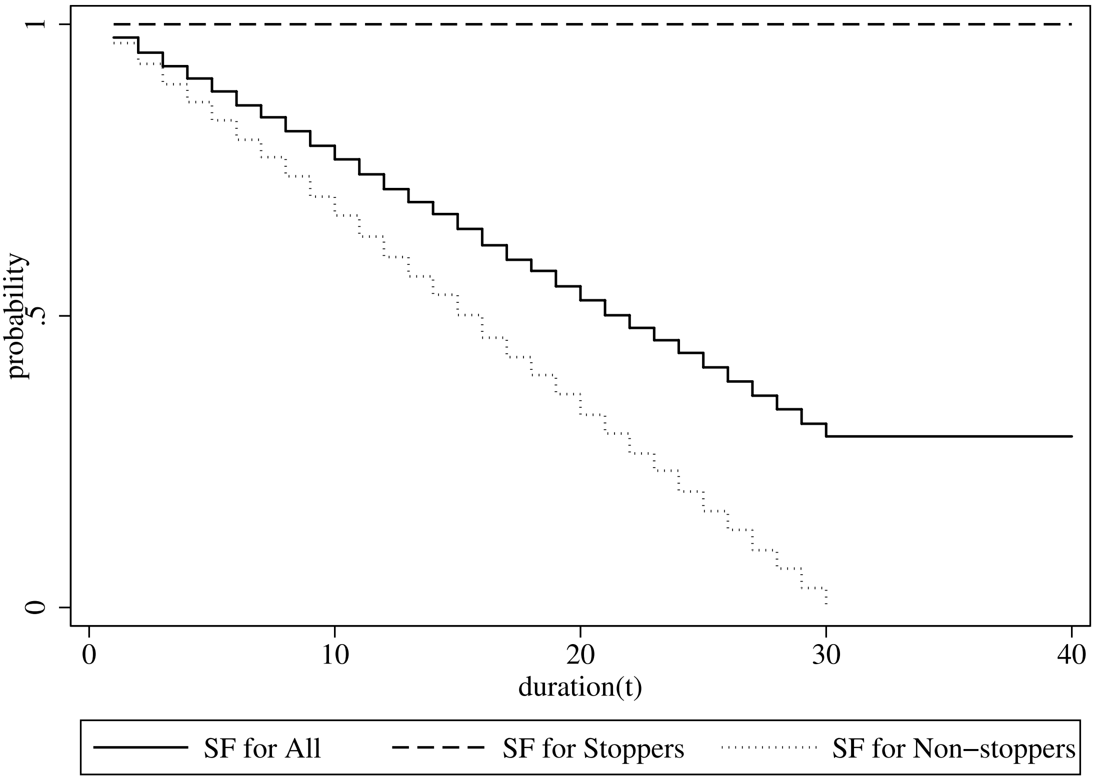

Figure 2: Survival Function Estimates in the case of 30 percent Birth-stoppers

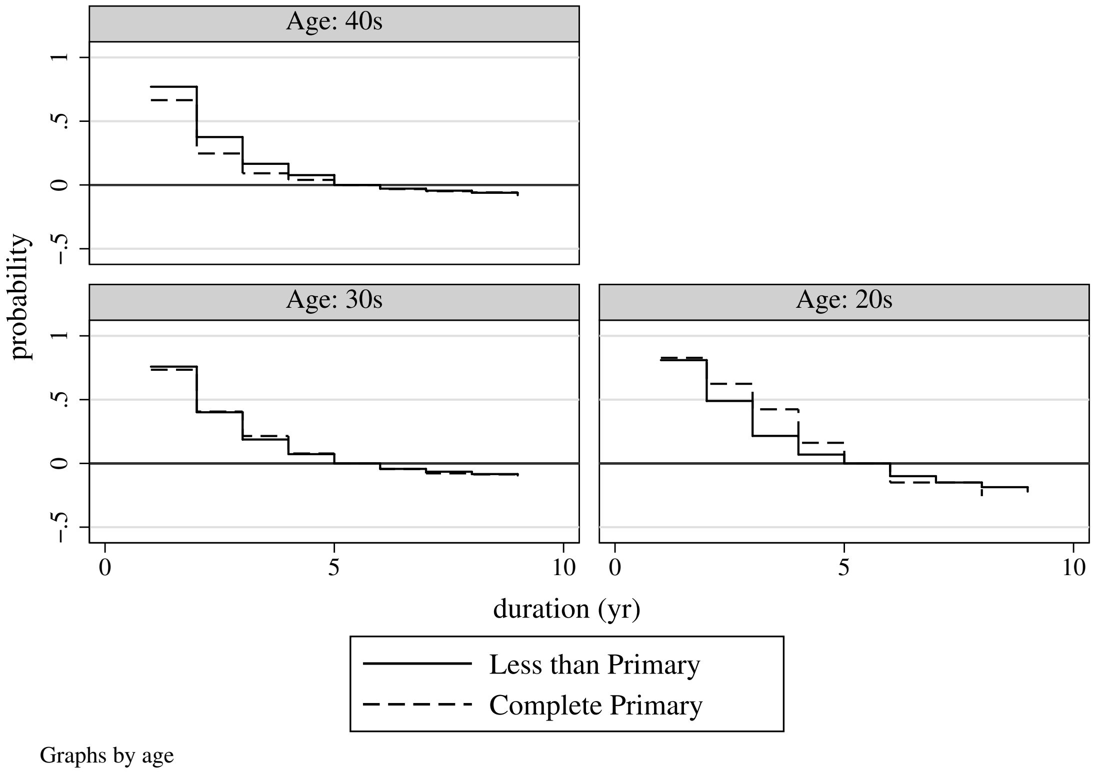

Figure 3: Survival Function Estimates assuming Birth-stoppers at t = 5

Notes: The data used are the 1993 Indonesian Family Life Survey.

Additionally, the proportion of women who have at least two children in their lifetime seems to be stable over the period 1971 to 1990. Figure 4 plots this statistic by age using the 1971, 1980 and 1990 Indonesian Population Census. Although there was a substantial delay in childbearing over time, most of the women seem to have at least two children during their lifetime.

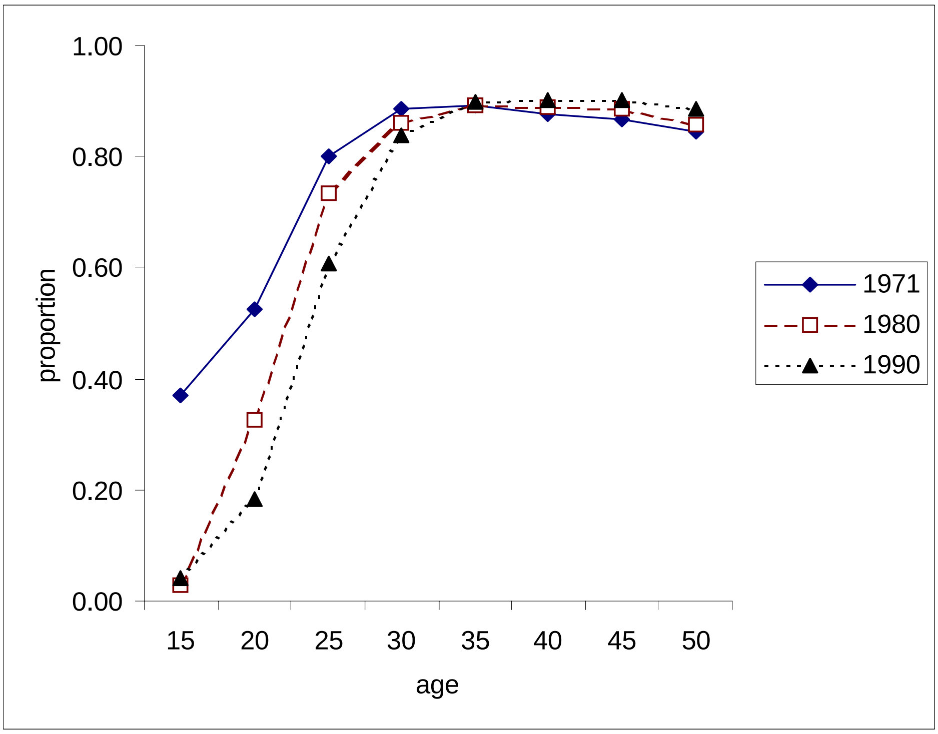

Figure 4: Proportion of Women with at least Two Children by Age

Notes: The data used are the 1971, 1980 and 1990 Indonesian Population Censuses

2.2 Duration Analysis

It is assumed that a woman's pregnancy history can be described in the following manner.[10] The woman is exposed to the risk of the first birth at a calendar time $\tau = 0$. This is the age of marriage. Transitions occur on or after $\tau = 0$. A finite-state discrete time birth process is defined as $\{Y(\tau), \tau = 0, 1, 2, ...\}, Y(\tau) \in \Gamma$, where the set of possible states (parities) are finite ($\Gamma = \{0, 1, 2, ..., C\}, C < \infty$). Each element of Γ indicates the parity attained at time τ. In constructing the conditional hazard, $H(\tau)$ is defined as the relevant conditioning set at time τ.

The potential durations are denoted by $T_1, ..., T_C$. If a woman is at risk for the jth birth at time $\tau(j-1)$, the conditional hazard at duration t_j is defined to be

$$h_j(t_j|H(\tau(j-1)+t_j)) = \Pr(T_j = t_j | T_j \geq t_j, H(\tau(j-1)+t_j)). \qquad (1)$$

[10] The setup adopted is essentially the discrete version of that used in Heckman & Walker (1990a, 1990b, and 1991).

The survival function of duration t_j is

$$S_j(t_j|H(\tau(j-1)+t_j)) = \prod_{t=0}^{t_j} \left[1 - h_j(t|H(\tau(j-1)+t))\right]. \tag{2}$$

It is likely that there is an unobserved individual variation in fecundity that is unknown to both the agent under study and the econometrician. It is assumed that this variation can be summarized as a scalar random variable Θ. It is also assumed that the distribution of Θ, $M(\theta)$, is time-invariant, and that Θ is independent of $H(0)$. Then, the conditional hazard becomes a function of Θ,

$$h_j(t_j|H(\tau(j-1)+t_j),\theta) = \Pr(T_j = t_j|T_j \geq t_j, H(\tau(j-1)+t_j),\theta). \tag{3}$$

In general, in the presence of unobserved heterogeneity, the separate estimation of each transition produces inconsistent estimates of coefficients of interest because the relevant distribution of Θ for the agent is $M(\theta|H(\tau))$, not $M(\theta)$ in the density of T_j. Heckman & Walker (1990a, 1990b, and 1991) provide a set of sufficient conditions for using the separate estimation of each birth hazard: (1) no defective distribution, (2) no censoring, and (3) $H(0) = H(\tau)$, for all τ. For the purpose of this study, these conditions are reasonably well satisfied in estimating the transition to a second birth separately. No defective distribution means that all the women eventually have a second child. As discussed in the previous section, most women in Indonesia seem to have more than one child eventually, and the distribution of women with at least two children does not differ across educational groups. The second condition means that there is no censoring from all the married women to married women with at least one child. As discussed in the data description, the selection does not seem to be severe. The third condition implies that the conditioning set is time-invariant. a set of time-invariant regressors is used in this section, and the result is not sensitive to the inclusion of new information acquired after the first birth (age at first birth). However, time-varying covariates are used in hypothesis testing in section 4. Therefore, further robustness tests will estimate the first birth interval and the second birth interval together to see if the main result of this paper is sensitive to the third condition.

In this study, the second birth interval is estimated in order to examine the effect of women's education on the second birth interval separately. Further, a logit hazard model is used with a year as a unit of time,

$$h_2(t|H(\tau(2)+t),\theta) = \Lambda(\alpha_t + x_t'\beta + \theta) = \frac{\exp(\alpha_t + x_t'\beta + \theta)}{1 + \exp(\alpha_t + x_t'\beta + \theta)}. \tag{4}$$

The effect of duration on hazard, α_t, implies the baseline hazard, and x_t gives the observed characteristics of individuals.[11] As long as unobserved heterogeneity, θ, is orthogonal to the observed characteristics, the existence of θ affects only inferences regarding time variation (Lancaster 1979).[12] It is not possible to identify time variation without further assumptions about functional form of time variation and distribution of

[11]As the interval width becomes smaller, the logit hazard model converges to the proportional hazard model (Thompson, 1977).

[12]It is unlikely that there is a systematic correlation between fecundity and women's primary education.

θ in the case of estimating a single birth interval. For correct inference when calculating standard errors of coefficients, the correlation between observations from the same woman due to unobserved heterogeneity is allowed.[13]

The result of the logit model in a hazard framework presented in Table 3 confirms the finding from the survival function analysis. With time-invariant basic variables as in column (1), the coefficients of education and the interaction term of education and year shows that the marginal effect of schooling on the probability of having a second birth is positive among earlier cohorts (implying that more educated women tend to have shorter birth intervals), and that it is negative among later cohorts (implying the opposite). Figure 5 plots the marginal effect of schooling evaluated at each year of the sample period, which shows the reversal of the sign in the effect of education on birth hazard. In Table 3 a duration of less than five years has a positive effect on birth hazard, and a negative effect after five years. Additionally a higher level of husbands' education is associated with higher hazard (shorter birth interval). The effect of women's education on birth hazard does not change when other characteristics are controlled. In column (2), having a girl as the first child does not have a significant effect on birth hazard. Women who marry later tend to have significantly shorter birth intervals. Being Muslim which is the case for around 85 percent of Indonesian population is associated with significantly longer birth intervals. Living with the woman's parents after marriage lowers the birth hazard significantly, while living with the woman's parents-in-law increases the birth hazard significantly. This may be because a woman living with her husband's family receives more pressure to have children. When duration specific effects (duration dummies) are allowed as in column (3), the result of the effect of education on birth hazard remains qualitatively the same.

It is possible that unobserved fecundity biases the coefficient on education. If more fertile women are likely to have their first baby earlier, this will bias the coefficient of education through the relationship between education and age at first birth.[14] The major results are not likely to be sensitive to this possibility because what is observed is the reversal of the effect of education on birth hazard, whereas the bias is based on the correlation between education and age at first birth, which seems to be stable over time.

Another potential bias comes from the endogeneity of educational choice. That is, women who choose to go to college are likely to enter the marriage market later, and women who marry later may have a certain preference for pregnancy, for example, a desire to have a shorter birth interval. This will bias the coefficient on education through the relationship between education and age at marriage. The descriptive statistics suggest that this is not likely to be the case. The average years of schooling in the sample changed from 4.4 years to 5.8 years when comparing the early cohort and the later cohort. In addition, less than 15 percent of women in the sample have more than 9 years of schooling, and less than 3 percent women have more than 12 years of schooling. Given that the average age at marriage is around 18 years, the educational

[13]The cluster (woman id) option is utilized in the logit estimation in the STATA program.

[14]Note that fecundity does not bias the coefficient of education if age at first birth is perfectly correlated with fecundity. To check this point, consider a linear probability model for simplicity. Then, $y = \beta_1 AGE_{birth} + \beta_2 EDU + \mu + \varepsilon$. Since age at first birth is perfectly correlated with fecundity (μ), $\mu = AGE_{birth} - \tilde{A}$. Then, the original equation becomes $y = -\tilde{A} + (\beta_1 + 1) AGE_{birth} + \beta_2 EDU + \varepsilon$. Since $\tilde{A}$ is a constant, β_2 can be estimated without a bias.

Table 3: Reversal of Relationship between Education and Birth Spacing

	(1) Logit	(2) Logit	(3) Logit
age at 1st birth		−0.0186	−0.0187
		(2.52)	(2.49)
duration	0.3054	0.3641	
	(10.93)	(12.23)	
duration^2	−0.0330	−0.0369	
	(11.82)	(12.39)	
year	−0.0261	−0.0245	−0.0250
	(8.76)	(7.33)	(7.36)
primary edu	2.6335	3.0367	3.2447
	(5.01)	(5.51)	(5.76)
primary edu*year	−0.0318	−0.0374	−0.0401
	(5.02)	(5.63)	(5.89)
husband's edu	0.0187	0.0158	0.0165
	(3.33)	(2.64)	(2.72)
1st baby female		−0.0157	−0.0163
		(0.37)	(0.38)
age at marriage		0.0145	0.0158
		(1.98)	(2.12)
Muslim		−0.1258	−0.1388
		(1.94)	(2.10)
living w/ own parents		−0.0944	−0.1003
		(1.99)	(2.08)
living w/ parents-in-law		0.1682	0.1696
		(3.02)	(3.00)
d1			1.4589
			(3.75)
d2			2.5299
			(6.52)
d3			2.5548
			(6.57)
d4			2.4009
			(6.13)
d5			2.2082
			(5.59)
d6			2.1014
			(5.25)
d7			1.5878
			(3.83)
d8-d10			1.5719
			(3.92)
d11-d15			0.9470
			(2.21)
constant	0.5952	0.5551	−0.9536
	(2.53)	(2.16)	(1.99)
no. of observations	13,206	12,286	12,286

Notes: The dependant variable is the index that takes 1 if a woman experienced a second birth (in that year). Note that a positive coefficient is associated with a higher hazard (shorter birth interval), and that a negative coefficient is associated with a lower hazard (longer birth interval). Absolute values of t-statistics are in parentheses. Omitted duration is "d15 above" in column (3).

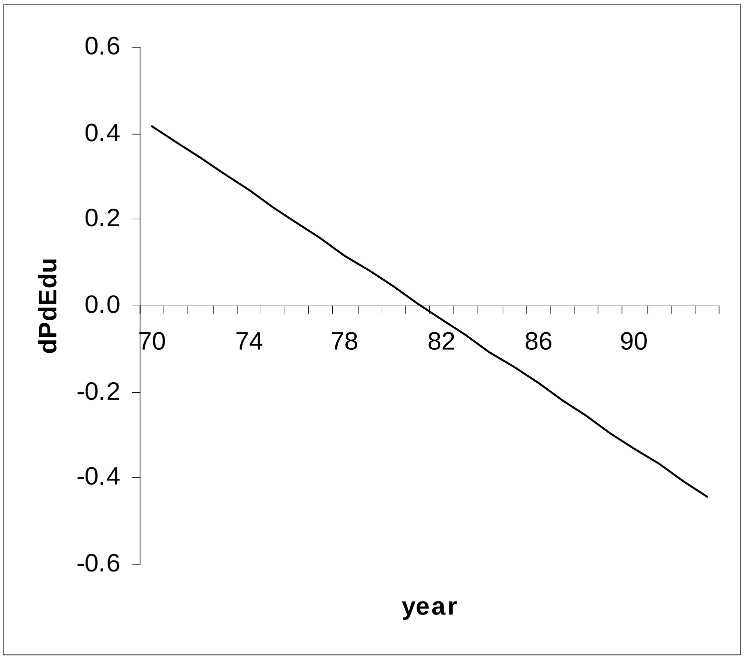

Figure 5: Marginal Effect of Schooling on Birth Hazard

choice seems to be predetermined to the fertility choice.[15]

3 Education and the Second Birth Interval: Theoretical Consideration

The framework adopted was developed in the demography literature Bulatao & Lee (1983), Easterlin & Crimmins (1985), and ascribes realized fertility to three factors: the demand for children, the supply of children, and the cost of fertility control. Since what is observed is the differential change in the birth spacing behavior of female educational groups, the exogenous variations that could have affected the educational groups differentially will be focused on.

3.1 Education and the Supply of Children

Although the relationship is imprecise, the time interval when a woman is exposed to the risk of pregnancy after the first birth is closely related to the duration of breastfeeding due to postpartum amenorrhea. Figure 6 presents the survival function estimates of the event of weaning the first child. The estimates indicate that more educated women tend to breastfeed for a shorter period of time than less educated women among earlier cohort, although the difference between education groups gets smaller among later

[15] As a simple way of checking out the possibility of marriage market selection, I include the fraction of women who are married in the same cohort as a measure of the probability of being married in the estimation. The coefficients of education and its interaction with year are not sensitive to the inclusion of this variable.

cohorts. The results from the duration analysis are presented in Table 4. However, the pattern becomes clearer in Figure 7, which plots the marginal effect of primary education on the duration of breastfeeding the first child at each year. The effect of primary education is positive (shorter period of breastfeeding) in the 1970s, and it is closer to zero in the 1990s. The literature on breastfeeding suggests that highly educated women tend to breastfeed for a shorter duration because they have a higher opportunity cost of time, they could afford to buy substitutes for mother's milk, or they might think of breastfeeding as undignified.[16]

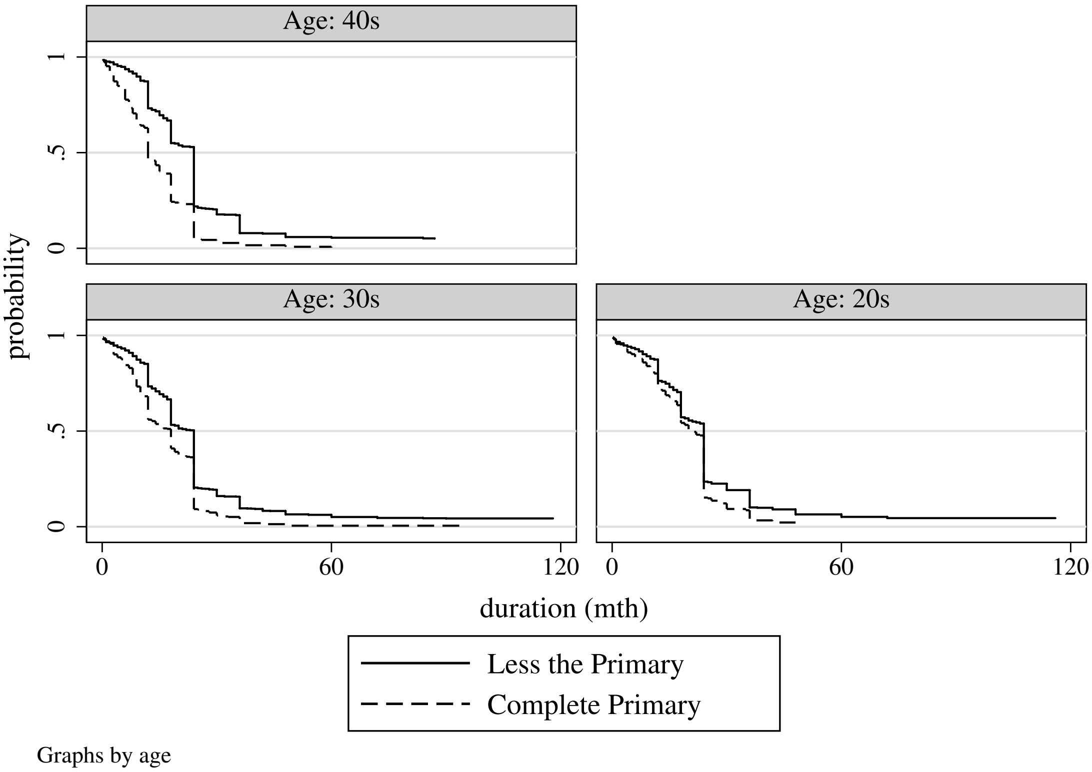

Figure 6: Duration of Breastfeeding

In any case, in the absence of a strong demand for children or with imperfect birth control, more educated women will have shorter birth intervals on average due to the earlier exposure to risk of pregnancy. Sigle (1998) suggests that the effect of breastfeeding on birth intervals is strongest within first six months of breastfeeding, but diminishing thereafter.[17] This study observes the average duration of breastfeeding first child among women with complete primary education is 11 months in the 1970s and 15 months in the 1990s, which implies that the effect of breastfeeding on birth interval is likely to become small over time. Therefore, the pattern of breastfeeding is presented as an explanation for the negative relationship between education and second

[16] See, e.g., Appleton (1996), Barrera (1991), Oni (1985), and Wolfe & Behrman (1982).

[17] Meredith, Menken & Chowdhury (1987) also find that the contraceptive effect of lactation diminishes over the period of partial breastfeeding.

Table 4: Education and Breastfeeding

	(1) Logit	(2) Logit	(3) Logit
age at 1st birth	0.0148	0.0101	0.0083
	(3.06)	(1.39)	(1.13)
duration	0.0939	0.0948	
	(21.85)	(21.74)	
duration^2	−0.0012	−0.0012	
	(14.04)	(13.95)	
year	−0.0129	−0.0116	−0.0112
	(4.53)	(3.96)	(3.80)
primary edu	1.1375	1.2024	1.4614
	(2.50)	(2.60)	(3.13)
primary edu*year	−0.0118	−0.0129	−0.0161
	(2.15)	(2.31)	(2.85)
husband's edu	0.0251	0.0230	0.0196
	(4.84)	(4.29)	(3.62)
1st baby female		−0.0549	−0.0463
		(1.47)	(1.23)
age at marriage		0.0022	0.0032
		(0.31)	(0.44)
Muslim		−0.3105	−0.3093
		(5.38)	(5.30)
living w/ own parents		−0.0977	−0.1006
		(2.32)	(2.36)
living w/ parents-in-law		−0.0773	−0.0651
		(1.57)	(1.31)
d3			0.9693
			(2.84)
d6			0.2724
			(0.79)
d9			0.7146
			(2.08)
d12			1.4969
			(4.40)
d15			1.0411
			(3.03)
d18			1.8966
			(5.57)
d21			1.0941
			(3.15)
d24			2.8506
			(8.38)
d27			1.0825
			(2.92)
d30			1.7457
			(4.87)
d33			0.3870
			(0.91)
d36			2.6938
			(7.64)
d39			1.3293
			(3.10)
d42			1.4978
			(3.49)
d48			1.8720
			(4.91)
d54			0.3651
			(0.60)
d60			1.7912
			(4.04)

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continued

	(1) Logit	(2) Logit	(3) Logit
constant	−3.1825	−2.8699	−3.2369
	(14.65)	(12.59)	(7.76)
no. of observations	57,141	55,446	55,446

Notes: The dependant variable is the index that takes 1 if a woman weaned her first child (in that month). Notice that a positve coefficient is associated with a higher hazard (shorter duration of breastfeeding), and that a negative is associated with a lower hazard (longer duration of breastfeeding). Absolute values of t-statistics are in parentheses. In column (3), d3 is a duration dummy that indicates first three months after first birth. Other duration dummies are defined in the same fashion. Omitted duration is "d60 above".

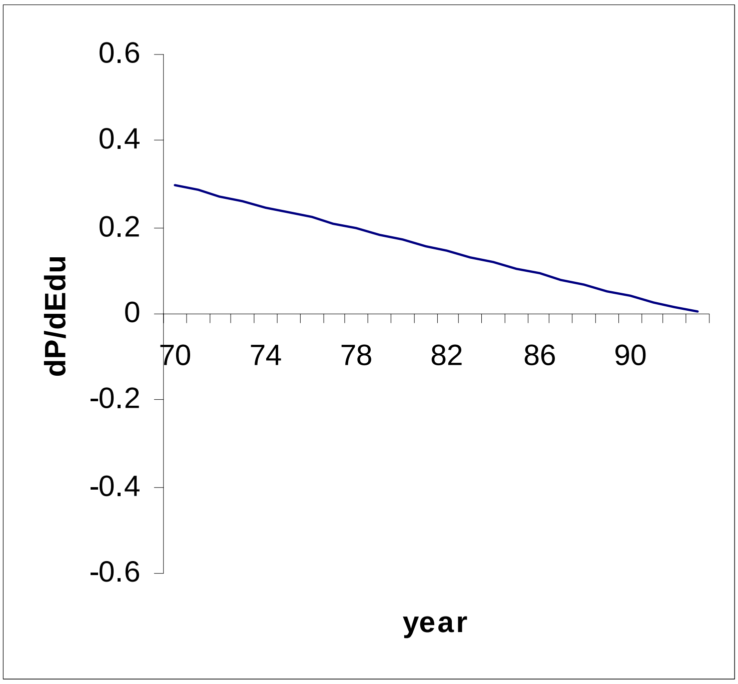

Figure 7: Marginal Effect of Schooling on Breastfeeding

birth interval among earlier cohorts.

3.2 Education and the Demand for Children

Indonesia experienced a rapid economic growth over the period from the early 1960s to the mid 1990s. Its GDP per capita increased from \$190 in 1961 to \$610 in 1991 (in 1991 U.S. dollars). In terms of shares of GDP, agriculture decreased from 54 percent in 1960 to 22 percent in 1990. The share of manufacturing and construction grew from 11 percent to 26 percent, while service industry grew from 32 percent to 39 percent in 1960 in 1990 (See Figure 8). The increase in industrialization is likely to be associated with the increase in the return to schooling. Cross sectional wage regressions from the IFLS 93 provide some evidence. The results presented in Table 5 suggest that the return to schooling is higher in areas where the share of manufacturing industry is higher and the share of service industry is lower. The point estimates of schooling interacted with manufacture and service industry dummies are significant at 10 percent and at 5 percent, respectively. This leads to one hypothesis for the reversal of the relationship between women's education and birth spacing in Indonesia.[18]

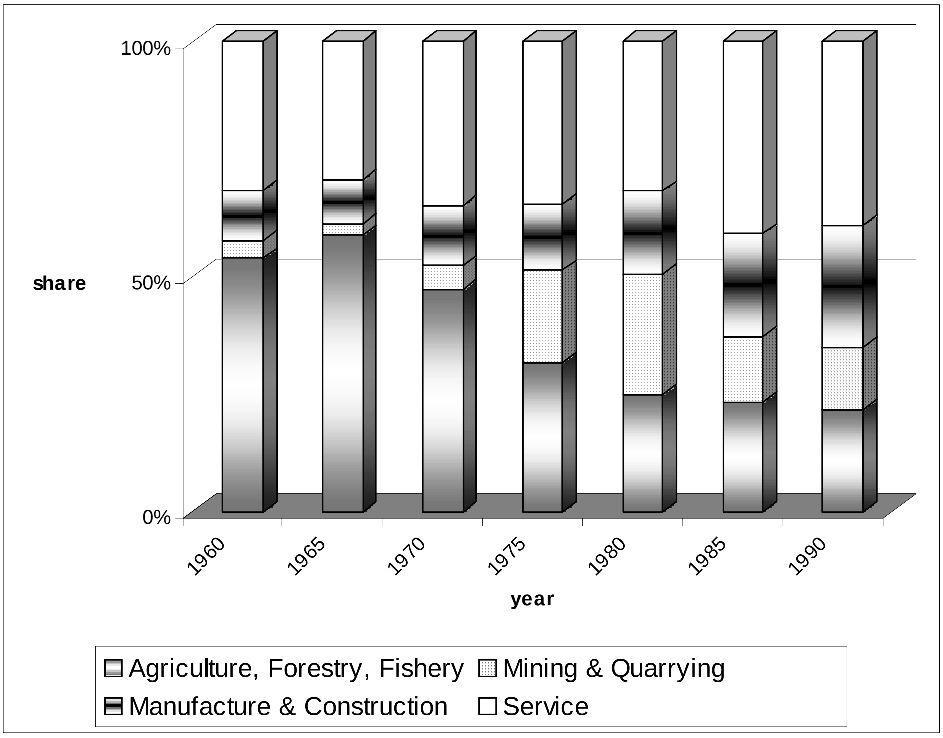

Figure 8: GDP Share by Industry in Indonesia

[18]The descriptive statistics from Indonesian population census suggest that the decrease of the average number of children ever-born over time is parallel over educational groups (See Figure 10 in Appendix A.). They also suggest that there is no differential change across educational groups regarding the inverse U-shape relationship between education and female labor market participation rate over the period 1970 to 1990 (See Figure 11 in Appendix A.). Therefore, no consideration is taken for the change in the number of children ever-born or the female labor market participation rate as potential explanations for the fact addressed.

Table 5: Effect of Industrialization on Returns to Schooling

	(1) OLS	(2) OLS
age	0.0549	0.0557
	(8.86)	(8.74)
age^2	−0.0006	−0.0006
	(8.41)	(8.54)
schooling	0.1182	0.1106
	(37.06)	(13.66)
female	−0.1690	−0.2113
	(5.84)	(7.04)
dummy for wage	−0.0007	−0.1091
	(0.02)	(2.43)
dummy for net profit	−0.1501	−0.1775
	(3.51)	(4.02)
manufacture 90		0.2070
		(0.67)
manufacture 90*sch		0.0722
		(1.75)
service 90		1.3872
		(10.12)
service 90*sch		−0.0409
		(2.42)
constant	−2.4404	−2.8613
	(17.21)	(18.66)
no. of observations	7,096	6,194
R^2	0.22	0.25

Notes: Dependent variable is log of hourly wage. There are three kinds of wage: wage, net profit, and gross income. Absolute values of t-statistics are in parentheses. Omitted category for industry dummies is agriculture and mining in column (2).

As noted earlier, one common implication of dynamic fertility models (e.g., Wolpin (1984) and Newman (1988)) is that a couple with a steeper income profile over their life cycle has an incentive to space births more widely in the absence of perfect capital markets. In the following model, I derive a set of conditions under which a steeper wage profile leads to a shorter or longer birth interval. Then, I discuss how the implication of the model can explain the reversal of the relationship between education and birth spacing in Indonesia.

A simple model of birth spacing is considered where all women work, and each woman plans to have exactly two children during her life time. At time zero, a woman has a choice over timing of two events: first birth (T_1) and second birth (T_2). Her wage profile, $y(t)$, is an increasing function of experience ($\frac{\partial y}{\partial t} > 0$). No time discounting is assumed. The result is not sensitive to discounting rate as long as the growth rate of wage profile is greater than discounting rate which is likely to be the case in a rapid growing economy like Indonesia. An imperfect capital market is assumed so that the budget constraint is binding at every period. Her instant utility has two arguments, child composite good, $c(t)$, and consumption of other goods, $x(t)$. Then, her lifetime utility is defined as

$$U = \int_0^L u(c(t), x(t))dt$$

Child composite good is the product of number of children and quality of children, q, where the latter is another choice variable. The price per unit quality of a child is denoted by p.[19] Since $c(t) = n(t)q$, and $x(t) = y(t) - n(t)p$, the maximization problem is

$$\max_{T_1,T_2,q} U(T_1, T_2, q) = \int_0^{T_1} u(0, y(t))dt + \int_{T_1}^{T_2} u(q, y(t) - pq)dt + \int_{T_2}^{L} u(2q, y(t) - 2pq)dt. \tag{5}$$

The first order condition amounts to

$$\frac{\partial U(\cdot)}{\partial T_2} = u(q, y(T_2) - pq) - u(2q, y(T_2) - 2pq) = 0 \tag{6}$$

$$\frac{\partial U(\cdot)}{\partial T_1} = u(0, y(T_1)) - u(q, y(T_1) - pq) = 0 \tag{7}$$

$$\begin{aligned} \frac{\partial U(\cdot)}{\partial q} &= \int_{T_1}^{T_2} \frac{\partial u(q, y(t) - pq)}{\partial c} - p\frac{\partial u(q, y(t) - pq)}{\partial x}dt \\ &+ \int_{T_2}^{L} \frac{\partial u(2q, y(t) - 2pq)}{\partial c} - 2p\frac{\partial u(2q, y(t) - 2pq)}{\partial x}dt = 0. \end{aligned} \tag{8}$$

In order to comprehend the implication of the model, the following specification of the utility function and wage function is taken.

$$u(n, x) = (n + \alpha)(x + \beta) \tag{9}$$

$$y(t) = \theta t \tag{10}$$

Then, the first order condition becomes

$$\frac{\partial U(\cdot)}{\partial T_2} = (q + \alpha)(\theta T_2 - p + \beta) - (2q + \alpha)(\theta T_2 - 2p + \beta) = 0 \tag{11}$$

$$\frac{\partial U(\cdot)}{\partial T_1} = \alpha\theta T_1 - (q + \alpha)(\theta T_1 - p + \beta) = 0 \tag{12}$$

$$\begin{aligned} \frac{\partial U(\cdot)}{\partial q} &= \frac{1}{2}\theta(T_2^2 - T_1^2) - (2pq + p\alpha - \beta)(T_2 - T_1) + \theta(L^2 - T_2^2) \\ &\quad -(8pq + 2p\alpha - \beta)(L - T_2) = 0. \end{aligned} \tag{13}$$

Assuming that unit price per quality is less than wage at the end of her life ($\theta L > p$) generates the following interior solution.

$$q^* = \frac{\theta L - \alpha p + \beta}{5p} \tag{14}$$

$$T_1^* = \frac{\theta L + 4\alpha p - 4\beta}{5\theta} \tag{15}$$

$$T_2^* = \frac{3\theta L + 2\alpha p - 2\beta}{5\theta} \tag{16}$$

[19] It is certainly restrictive to assume that the quality of a child is fixed over her lifetime. However, it can be argued that the education in early childhood is as important as that in later stage of a child's mental and physical development.

Consequently, the second birth interval and its derivative with respect to the slope of wage profile are

$$BI = T_2^* - T_1^* = \frac{2(\theta L - \alpha p + \beta)}{5\theta} \tag{17}$$

$$\frac{\partial BI}{\partial \theta} = \frac{2(\alpha p - \beta)}{5\theta^2}. \tag{18}$$

As can be seen above, the effect of increase in the wage growth rate on the second birth interval depends on the three parameters, α, β,and p. When marginal utility of child good is small relatively to that of consumption good ($\alpha p > \beta$), increase in the wage growth rate generates an incentive to have a longer birth interval. On the other hand, the marginal utility of child good is large relatively to that of consumption good ($\alpha p < \beta$), steeper income profile increases an incentive to have a shorter birth interval.

In general increase in wage growth rate generates income effect and substitution effect. Income effect means that increase in wage rate makes the timing of a birth earlier because a woman can afford to have a child earlier given the quality. The substitution effect is that with a higher income she demands more quality of a child, which increases the shadow price of a child, and consequently delays the childbearing. The total effect of the growth rate of wage profile on birth interval can be in either direction. Therefore, it may be that, among early cohorts, substitution effect dominates when highly educated women face steeper wage profile than the less educated counterparts: women with higher education have shorter birth intervals. On the other hand, among later cohorts, income effect dominates: women with higher education have longer birth intervals than the less educated counterparts. Given the return to schooling seems to increase with a level of industrialization as shown in Table 5, the differential effect of the industrialization on birth spacing across educational groups might explain the reversal of the relationship between education and birth spacing . I explore this possibility by looking at the effect of the level of industrialization of a local economy on birth hazard.

3.3 Education and the Cost of Fertility Control

Although a voluntary organization, the Indonesian Planned Parenthood Association (PKBI) was formed in 1957, and promoted family planning programs through the sales of contraceptives throughout the 1960s. A serious national level program was not implemented until the Indonesian government invited a group of foreign experts, sponsored by the UN, World Bank, and WHO, to evaluate the country's family planning program in 1969. With the detailed recommendations of the group, the Indonesian government initiated a Five-year Family Planning Program (1971-5) for Java and Bali (stage I). In the fiscal year 1970/71, the family planning program received an equivalent of US$1.3 million from the government and over US$3 million from foreign donors. The funding for the family planning program increased dramatically over time. In 1984, it was estimated that the funding from the government was about US$65 million, that US$25 million was from foreign donors (USAID, 1984:15).

With the beginning of the Second Five-Year Development Plan, the family planning program expanded beyond Java and Bali to ten large provinces in 1974 (stage II). At the same time, the National Family Planning Coordinating Board (BKKBN) increased the intensity of the program in Java and Bali through a village family planning system.

In 1977, all the remaining provinces were included in the family planning program, and the village family planning program began to be extended beyond Java and Bali (stage III). The development of the family planning program is well illustrated by the contraceptive-use rates in Table 6. The estimated proportion of married women of reproduction age using contraception in stage I provinces increased from two to seven percent in 1971/72 to between 39 and 60 percent in 1985. Stage II provinces experienced an increase from one to four percent in 1974/1975 to between 22 and 60 percent in 1985. The contraceptive-use rates for stage III provinces also increased from one to nine percent in 1979/1980 to between 10 to 42 percent in 1985.

Table 6: Estimated Contraceptive Prevalence in Indonesia in Selected Years under Successive Development Plans

	Province	1971/2 Repelita Third Year	1974/5 Repelita II First Year	1979/80 Repelita III First Year	1984/5 Repelita IV Revised Figures July 1985	1985 Supas
	Indonesia	3	13	29	51	38
Stage I						
1	DKI Jakarta	4	10	20	46	44
2	Jawa Barat	2	11	21	54	44
3	Jawa Tengah	2	13	43	57	39
4	Daerah Istimewa Yogyakarta	4	16	57	57	53
5	Jawa Timur	4	27	51	58	40
6	Bali	7	28	50	75	60
Stage II						
1	Daerah Istimewa Aceh		2	7	44	22
2	Sumatera Utara		2	14	45	30
3	Sumatera Barat		1	15	41	26
4	Sumatera Selatan		2	8	49	29
5	Lampung		1	18	41	42
6	Nusa Tenggara Barat		1	13	45	25
7	Kalimantan Barat		1	7	42	22
8	Kalimantan Selatan		2	17	48	39
9	Sulawesi Utara		4	32	45	60
10	Sulawesi Selatan		2	14	41	23
Stage III						
1	Riau			1	23	21
2	Jambi			4	32	38
3	Bengkulu			9	40	42
4	Nusa Tenggara Timur			1	20	29
5	Kalimantan Tengah			4	27	29
6	Kalimantan Timur			5	35	37
7	Sulawesi Tengah			3	34	38
8	Sulawesi Tenggara			3	31	24
9	Maluku			2	21	17
10	Irian Jaya			1	17	17
11	Timor Timur				6	10

Notes: The numbers indicate the estimated proportion of married women of reproductive age using contraception. The data source includes BKKBN Monthly Service Statistics, July 1985 and BPS, 1986. This table is requoted from Hugo et. al. (1995), p.145.

The dramatic increase in the availability of contraceptives over last three decades provides one possible explanation for the effect of education on the birth hazard discussed earlier. When a family planning program is introduced, the more educated women will have better access to modern methods of birth control if they are better at adopting or at making use of new methods as suggested by Schultz (1975). This implies that highly educated women are able to space their births even if it is still the case that they tend to breastfeed for a shorter period of time than the less educated counterparts. The hypothesis that the effect of the family planning program is stronger for more educated women has been supported by a number of studies. For example, Rosenzweig & Schultz (1987) conclude that the mother's education mitigates the consequences of exogenous variations in the supply of births.[20] Therefore, this hypothesis is tested by examining whether the presence of a family planning program has *differential* effects on educational groups over the sample period.

4 Empirical Results

4.1 Key Variables

There are 4,980 married women in the sample. A birth history is constructed for each woman using her total live births. Twins are counted as one birth. The first set of key variables is the measure of industrialization in the economy and the occupational choice of the individuals. The occupational composition of a local labor market is constructed as a measure of industrialization. Indonesia has 27 provinces and approximately 300 districts (kabupatan). The kabupatan is considered a local labor market, and the 1971, 1980, and 1990 Indonesian Population Censuses are used in order to make a linear projection of the occupational composition over the years between them. The IFLS 93 provides extensive data on work history, which allows the occupational choice of married women in the year closest to the year of her first birth to be identified. Since the history of occupational choice is incomparable, I do not include women for whom there is more than a five-year difference between the year of their documented occupational choice and the year of their first birth.

The second set of key variables is the measure of availability of a family planning program by community and by year. Five variables are proposed. The first one is an index variable, which takes value only if there was a family planning clinic at the community in a particular year. The question of when the first family planning clinic (PPKBD) was built was asked to the head of the village and the chairman of the Association of Family Activities (PKK). These two answers do not always coincide, and where they disagree, the average of the answers is taken. The second variable indicates if there was an Integrated Service Post (Posyandu) at the village in each year. This is constructed in the same way as the family planning clinic variable. The third variable is the number of community health centers (Puskesmas) by year. In the early stage of a family planning program, community health centers were used as posts for distributing contraceptives in the absence of family planning clinics, and since the beginning of the program, they have been complementary to the family planning

[20] The fact that women with primary education tend to breastfeed more over time discussed in the previous section can also be interpreted as evidence for the return to schooling in adopting new information.

clinics. The number of Puskesmas in a year includes those in the village, as well as those outside the village, but used by people in the community. It is constructed as the total number of Puskemas mentioned by the head of village, the chairman of PKK, and the staff of the Puskesmas. It should be noted that the maximum number of Puskesmas answered by heads of village and PKK was five. Although there are few communities in which a respondent provided information about five community health centers, it is possible that this variable is top–coded. An additional concern is that the information from the staff of Puskesmas is restricted to three community health centers per village. However, these three sources together provide a reasonable measure of the access to community health centers. The final two availability measures are the number of family planning clinics per 1,000 women in the age between 15 and 49 at both the provincial and the national level in each year from 1971 to 1990. The population of target women is projected using the 1971, 1980, and 1990 Indonesian Population Censuses. Biro Pusat Statistik (BPS) provides the number of family planning clinics by province through the Statistical Yearbook of Indonesia series.

4.2 Differential Change in the Demand for Children

To assess the demand-side influence of education proposed above, two variables are considered that are likely to be associated with the return to schooling. The first is the level of industrialization of a local economy, for which the occupational composition is used as a measure. In the specification, I include the set of dummy variables representing occupational compositions of labor force ($Ind_{kab,t}$), in addition to the basic characteristics of individual from Table 3 (x_{it}),

$$h_2(t|H(\tau(2)+t),\theta) = \Lambda(\alpha_t+\beta' x_{it}+\delta_1 Ind_{kab,t}+\delta_2 Ind_{kab,t}\times Yr_t+\delta_3 Ind_{kab,t}\times Edu_i+\theta). \tag{19}$$

It is predicted that more educated women tend to have longer birth intervals in the face of industrialization ($\delta_3 < 0$ in equation (19)). To address the concern that the location of the industry is endogenous to the unobserved propensity to conceive at the local economy level, the kabupatan fixed-effects are allowed for.

The second return-to-schooling variable is the occupational choice at an individual level. A set of occupation dummies (Occ_{it}) is included the hazard estimation. Occupations that exhibit an increasing depreciation of human capital over time will be associated with longer birth intervals ($\gamma_2 < 0$ in equation (20)),

$$h_2(t|H(\tau(2)+t),\theta) = \Lambda(\alpha_t+\beta' x_{it}+\gamma_1 Occ_{it}+\gamma_2 Occ_{it}\times Yr_t+\gamma_3 Occ_{it}\times Edu_i+\theta). \tag{20}$$

Using the community-level industrialization measure causes the sample to be restricted due to the availability of occupational composition series. Table 7, column (1) shows that the subsample preserves the same relationship between education and birth spacing as in the total sample. For the estimation in Table 7, the ten industries are categorized into three groups. Agriculture and mining form the first group, manufacturing, water and construction form the second group, and trade, transportation, finance, and government are put into the third group.

Based on the hypothesis described above, women who are more educated should space their births more widely in the face of higher level of industrialization. This

Table 7: Effect of Occupational Composition of Local Labor Market I

	(1) Logit	(2) FE Logit	(3) FE Logit	(4) FE Logit
age at 1st birth	−0.0254	−0.0266	−0.0253	−0.0265
	(2.69)	(2.77)	(2.66)	(2.74)
duration	0.3387	0.3612	0.3403	0.3633
	(9.74)	(10.30)	(9.77)	(10.33)
duration^2	−0.0355	−0.0368	−0.0356	−0.0370
	(10.15)	(10.46)	(10.17)	(10.49)
year	−0.0249	−0.0220	−0.0137	−0.0109
	(4.56)	(3.91)	(1.27)	(0.98)
primary edu	3.0716	3.1841	2.8800	3.0392
	(3.55)	(3.65)	(3.01)	(3.14)
primary edu*yr	−0.0382	−0.0399	−0.0359	−0.0379
	(3.62)	(3.75)	(3.12)	(3.24)
husband's edu	0.0160	0.0136	0.0161	0.0138
	(2.24)	(1.86)	(2.25)	(1.89)
second	−0.8715	−1.2553	8.0941	9.0782
	(2.07)	(2.52)	(1.36)	(1.49)
second*year			−0.1093	−0.1257
			(1.50)	(1.68)
second*primary edu			0.0061	−0.1337
			(0.01)	(0.14)
third	0.1836	−0.3202	−0.4248	−1.3913
	(1.25)	(0.78)	(0.21)	(0.66)
third*year			0.0072	0.0132
			(0.28)	(0.51)
third*primary edu			0.0182	−0.0044
			(0.06)	(0.01)
1st baby female	−0.0218	−0.0211	−0.0230	−0.0227
	(0.44)	(0.42)	(0.46)	(0.45)
age at marriage	0.0262	0.0254	0.0260	0.0251
	(2.82)	(2.68)	(2.78)	(2.64)
Muslim	−0.1248	−0.1342	−0.1255	−0.1364
	(1.61)	(1.68)	(1.62)	(1.70)
living w/ own parents	−0.0797	−0.0723	−0.0798	−0.0725
	(1.41)	(1.24)	(1.41)	(1.24)
living w/ parents-in-law	0.1795	0.1429	0.1778	0.1411
	(2.76)	(2.17)	(2.73)	(2.14)
constant	0.6872		−0.2139	
	(1.56)		(0.24)	
no. of observations	8,547	8,547	8,547	8,547
no. of Kabupatans		35		35

Notes: Absolute values of t-statistics are in parentheses. Manufacturing, water and construction form the second group, and trade, transportation, finance, and government are put into the third group. Omitted industry is first industry, which include agriculture and mining.

implies that including industry interacted with year and education will make insignificant the coefficients of education and its interaction with year. As a first step, the estimation without the interaction terms is implemented. The second industry group has a significantly negative effect on birth hazard, and the third industry group has a significantly positive effect without the kabupatan fixed effect as seen in column (2) of Table 7. The kabupatan fixed effect estimation suggests that there is some correlation between the location of industries and unobserved propensity to conceive at kabupatan level. After this correlation is taken into account by introducing fixed-effects, it appears that women in locations where more people work in the second industry tend to have significantly longer birth intervals. The effect of the individual characteristics other than women's education does not change in the fixed effect estimation with the exception of parents-in-law's co-residence, which becomes insignificant at 5 percent level. The reversal of the relationship between education and birth intervals is still present with the same magnitude after industry variables are included.

Including industries variables interacted with year and education does not give any significant change in the coefficients of education and its interaction, nor provides any evidence for the differential effect of industrialization on educational groups. This can be seen by comparing columns (2) and (4) of Table 7. The coefficients on education and its interaction with year become slightly smaller in terms of magnitude and significance, but both of them remain strongly significant. However, the interaction terms between education and industry groups have insignificant coefficients. Even so, the fact that industry-related variables are jointly significant in both specifications suggests that the second industry group has a significantly negative effect on birth hazard of all women as compared to the first industry, and that the third industry group does not have any significant effect on birth hazard.

Disaggregating the three groups does not change the result qualitatively, as seen in Table 8. Including the time-education-industry interactions does not change the coefficients of education and its interaction with year as in column (4). The manufacturing industry share has a significantly negative effect (longer birth interval) on birth hazard, and the transportation industry share has a positive effect (shorter birth interval) given that industry variables are jointly significant in column (2). The manufacturing industry share interacted with education has a negative coefficient, which suggests that the manufacture industry has more effect on more educated women. However, it is not significant. Further, there is no single industry whose interaction with education is significant in column (4). Therefore, there is no evidence for the differential effect of industrialization on birth hazard across educational groups.

The extensive retrospective questionnaire in IFLS 93 enables us to track the occupational choice of a woman at the year close to the year of her first birth. Due to the endogeneity of the labor market participation choice, the following analysis is restricted to women who have ever worked. This restriction and the availability of data reduce the sample size to 40 percent of the original sample. As presented in column (1) of Table 9, this subsample still exhibits the reversal of relationship between education and birth spacing over generations. One hundred occupation categories have been simplified into four major groups, which are agriculture, professional, manufacture and service. Including occupational groups in the hazard model does not change the coefficients on education and its interaction with year. Column (2) shows that

Table 8: Effect of Occupational Composition of Local Labor Market II

	(1) Logit	(2) FE Logit	(3) Logit	(4) FE Logit
age at 1st birth	−0.0260	−0.0286	−0.0270	−0.0294
	(2.74)	(2.95)	(2.82)	(3.02)
duration	0.3464	0.3681	0.3557	0.3761
	(9.92)	(10.45)	(10.11)	(10.60)
duration^2	−0.0360	−0.0372	−0.0368	−0.0379
	(10.27)	(10.54)	(10.40)	(10.64)
year	−0.0288	−0.0287	0.0008	−0.0026
	(4.26)	(4.09)	(0.05)	(0.18)
primary edu	2.7697	3.0885	3.7610	3.8643
	(3.14)	(3.46)	(2.70)	(2.73)
primary edu*yr	−0.0345	−0.0387	−0.0468	−0.0478
	(3.21)	(3.54)	(2.81)	(2.83)
husband's edu	0.0149	0.0146	0.0144	0.0148
	(2.07)	(1.99)	(1.99)	(2.01)
mine	−0.7451	−0.5936	42.4398	40.9787
	(0.45)	(0.34)	(1.71)	(1.64)
mine*year			−0.5023	−0.4791
			(1.66)	(1.58)
mine*primary edu			−4.8348	−5.4821
			(1.34)	(1.48)
manu	−0.5942	−1.0999	0.1969	0.9975
	(1.15)	(1.78)	(0.03)	(0.12)
manu*year			−0.0060	−0.0236
			(0.06)	(0.24)
manu*primary edu			−1.1366	−0.9243
			(0.90)	(0.71)
water	−11.6067	17.0066	−698.0151	−627.0656
	(0.53)	(0.71)	(2.25)	(1.90)
water*year			8.4863	8.0069
			(2.15)	(1.91)
water*primary edu			20.0499	0.4961
			(0.38)	(0.01)
cons	1.3168	−0.8641	73.1451	69.8296
	(0.74)	(0.41)	(2.64)	(2.43)
cons*year			−0.8751	−0.8660
			(2.58)	(2.47)
cons*primary edu			0.7437	0.0124
			(0.17)	(0.00)
trade	−2.7451	−1.8913	6.4210	0.3873
	(3.85)	(2.20)	(0.61)	(0.03)
trade*year			−0.1222	−0.0320
			(0.93)	(0.23)
trade*primary edu			2.6724	1.6148
			(1.57)	(0.91)
trans	4.6303	6.3507	−45.4866	−50.1236
	(3.18)	(3.23)	(2.01)	(2.15)
trans*year			0.6235	0.7149
			(2.18)	(2.43)
trans*primary edu			−0.6201	−0.5364
			(0.19)	(0.16)
finan	0.8177	−0.3094	38.9778	49.7993
	(0.30)	(0.10)	(0.80)	(0.97)
finan*year			−0.4344	−0.5657
			(0.75)	(0.94)
finan*primary edu			−5.2878	−3.2365
			(0.92)	(0.56)
gov	0.7468	−0.0616	5.6329	10.5441
	(1.53)	(0.09)	(0.83)	(1.52)
gov*year			−0.0622	−0.1357
			(0.72)	(1.52)
gov*primary edu			−0.9184	−0.0627
			(0.83)	(0.06)

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continued

	(1) Logit	(2) FE Logit	(3) Logit	(4) FE Logit
uncla	−3.0475	−4.7774	35.9703	15.6056
	(1.31)	(1.83)	(0.61)	(0.25)
uncla*year			−0.4808	−0.2310
			(0.59)	(0.27)
uncla*primary edu			−11.3507	−12.2530
			(1.66)	(1.76)
1st baby female	−0.0176	−0.0161	−0.0198	−0.0178
	(0.35)	(0.32)	(0.40)	(0.35)
age at marriage	0.0259	0.0272	0.0260	0.0278
	(2.77)	(2.85)	(2.75)	(2.90)
Muslim	−0.1316	−0.1490	−0.1256	−0.1507
	(1.69)	(1.85)	(1.60)	(1.85)
living w/ own parents	−0.0726	−0.0671	−0.0771	−0.0688
	(1.28)	(1.15)	(1.35)	(1.17)
living w/ parents-in-law	0.1759	0.1462	0.1712	0.1456
	(2.69)	(2.21)	(2.60)	(2.19)
constant	1.1442		−1.2247	
	(2.05)		(1.08)	
no. of observations	8,547	8,547	8,547	8,547
no. of Kabupatans		35		35

Notes: Absolute values of *t*-statistics are in parentheses. Omitted industry is agriculture.

they become slightly larger in terms of both magnitude and significance. Given that the occupational-group variables are jointly significant, women with professional or managerial occupations tend to have significantly shorter birth intervals. Including occupations interacted with year and education does not change the relation between education and birth spacing over time as shown in column (3).[21]

The strong significance of the coefficients on service and its interaction with year and education suggests that women in service-related occupations experienced a dramatic change in birth spacing behavior over time and across education groups. Specifically, women who had service-related jobs tend to have had shorter birth intervals in the 1970s, and women of the same occupational group in the 1990s have longer birth intervals. This provides some evidence for change in the growth rate of a wage profile in the service industry over time. In addition, more educated women in the service industry tend to have shorter birth intervals net of the calendar effect. However, these effects of occupation are not enough to explain the change in the relationship between education and birth spacing behavior over time, given the continued significance of education and its interaction with year.

[21] We could observe no change in the coefficient on education after including interaction terms in column (3) of Table 9, even though the occupational choice had an impact on the coefficient of education. That is, the omitted variable bias in column (1) and the bias in column (3) due to the endogeneity of occupational choice can be similar in terms of magnitude, in which case we observe the coefficients of education in both specifications are close to each other. Therefore, it is only claimed that these biases are likely to be small, given that coefficients of education are fairly stable throughout other specifications.

Table 9: Effect of Occupations

	(1) Logit	(2) Logit	(3) Logit
age at 1st birth	−0.0384	−0.0407	−0.0419
	(2.93)	(3.05)	(3.08)
duration	0.5144	0.5138	0.5146
	(10.02)	(10.01)	(10.02)
duration^2	−0.0522	−0.0518	−0.0514
	(9.26)	(9.21)	(9.14)
year	−0.0406	−0.0389	−0.0262
	(7.68)	(7.32)	(3.80)
primary edu	2.4732	2.2666	1.9065
	(2.78)	(2.52)	(1.75)
primary edu*yr	−0.0311	−0.0297	−0.0288
	(2.94)	(2.79)	(2.22)
husband's edu	0.0247	0.0265	0.0260
	(2.68)	(2.78)	(2.71)
professional		0.2633	1.0871
		(1.77)	(0.63)
professional*year			−0.0078
			(0.39)
professional*primary edu			0.0413
			(0.09)
manufacture		−0.1752	1.1149
		(1.73)	(1.13)
manufacture*year			−0.0161
			(1.34)
manufacture*primary edu			0.2399
			(0.91)
service		−0.1450	2.5645
		(1.55)	(2.92)
service*year			−0.0343
			(3.17)
service*primary edu			0.5222
			(2.20)
family worker		0.0021	0.0102
		(0.02)	(0.12)
1st baby female	−0.1682	−0.1711	−0.1705
	(2.65)	(2.68)	(2.67)
age at marriage	0.0369	0.0373	0.0385
	(2.87)	(2.85)	(2.91)
Muslim	−0.1107	−0.1047	−0.1241
	(1.22)	(1.15)	(1.35)
living w/ own parents	−0.0896	−0.0901	−0.0875
	(1.23)	(1.23)	(1.19)
living w/ parents-in-law	0.2024	0.2067	0.2071
	(2.49)	(2.52)	(2.52)
constant	1.6483	1.6138	0.6235
	(3.99)	(3.84)	(1.13)
no. of observations	5,572	5,572	5,572

Notes: Absolute values of t-statistics are in parentheses. Omitted occupation is agriculture.

4.3 Differential Change in the Cost of Fertility Control

The nature of a family planning program is to reduce the price of delaying births. The price of delaying births is specified, in addition to the basic individual characteristics in the hazard model,

$$h_2(t|H(\tau(2)+t),\theta) = \Lambda(\alpha_t + \beta' x_{it} + \rho P_Z + \theta). \tag{21}$$

The price of delaying births is considered as a function of the family planning program, calendar time, and individual education as in equation (22),

$$P_z = g(FP, Year, Edu). \tag{22}$$

One testable implication derived from the discussion of family planning programs, education, and birth spacing is that family planning programs have potentially differential effects on educational groups.

The availability of the family planning program is measured at the village, provincial, and national level, and the village-level family planning program interacted with calendar year and women's education are also included as in equation (23),

$$\begin{aligned} & h_2(t|H(\tau(2)+t),\theta) \\ = \; & \Lambda(\alpha_t + \beta' x_{it} + \rho_1 FP_{vil,t} + \rho_2 Yr_t + \rho_3 FP_{vil,t} \times Yr_t + \rho_4 FP_{vil,t} \times Edu_i \\ & + \rho_5 FP_{prov,t} + \rho_6 FP_{nation,t} + \theta). \end{aligned} \tag{23}$$

The framework predicts that the marginal effect of a family planning program on birth hazard increases as woman's education increases ($\rho_4 < 0$). As discussed in previous studies (e.g., Pitt, Rosenzweig & Gibbons (1993), Gertler & Molyneaux (1994), and Gertler & Molyneaux (2000)), the Indonesian family planning program has not been expanded over time in a random manner. Rather, it has reflected the local demand through an allocation of the budget at each administrative level. Therefore, it is likely that unobserved heterogeneity in the propensity to conceive at the village level is correlated with the level of the family planning program. This correlation can be removed by using a village fixed-effects estimation if the policy rule reflects the local unobserved propensity.

The second set of estimation results is presented in Table 10, which shows the results when only the existence of a family planning (FP) clinic is used as a measure of the FP program. Since the data on the history of the FP program are available for the villages covered in IFLS 93, this analysis is restricted to the women who have not migrated after their first birth. This restriction, combined with the availability of the measure of the FP program, leaves us with 40 percent of the total sample. As can be seen from column (1) in Table 10, this subsample shows the same reversal of the relationship between education and birth spacing. Without the village fixed-effects, the effect of the FP clinic is positive and significant as shown in column (2) in Table 10. With a community fixed effect, however, the effect of the FP clinic becomes insignificant. Although the coefficient of the FP clinic is still not significant in column (3) in Table 10, this suggests that the placement of the Indonesian Family Planning program is not random given that the Hausman test rejects the null hypothesis that the fixed-effects specification is not different from the one without it. The correlation of the FP program and the unobserved propensity to conceive at the village level appears to be negative, which implies that there are more FP program inputs in areas where women have a higher propensity to delay births. This is counterintuitive, but women with a higher propensity to delay births have a higher demand for contraceptives. Therefore, the FP program was placed more intensively in areas with higher demand for contraceptives, which implies an efficiency criterion. The comparison between column (2) and column (3) of Table 10 also suggests that the variation in birth hazards due to the gender of first baby, age at marriage, being Muslim, and living with own parents come partly from the local variation, which disappears in the fixed-effects estimation. Similar results are found in the comparison of column (4) and column (5) of Table 10. When the FP clinic variable is included, the marginal effect of primary education on birth hazard over time is still significant (the coefficients on primary education and its interaction with year are jointly significant at the 1 percent level).

Table 10: Effect of Family Planning Programs I

	(1) Logit	(2) Logit	(3) FE Logit	(4) Logit	(5) FE Logit
age at 1st birth	−0.0439	−0.0438	−0.0496	−0.0437	−0.0492
	(3.60)	(3.60)	(3.38)	(3.58)	(3.35)
duration	0.3430	0.3439	0.5335	0.3443	0.5370
	(8.26)	(8.26)	(11.78)	(8.28)	(11.85)
duration^2	−0.0345	−0.0345	−0.0433	−0.0345	−0.0436
	(8.44)	(8.43)	(9.95)	(8.43)	(10.00)
year	−0.0405	−0.0394	−0.0646	−0.0645	−0.0367
	(5.46)	(1.03)	(1.56)	(1.37)	(0.71)
primary edu	1.9809	1.9576	2.3182	1.1579	0.2741
	(1.52)	(1.49)	(1.54)	(0.74)	(0.15)
primary edu*year	−0.0246	−0.0242	−0.0309	−0.0133	−0.0036
	(1.60)	(1.55)	(1.74)	(0.69)	(0.16)
husband's edu	0.0107	0.0106	−0.0004	0.0105	−0.0012
	(1.22)	(1.20)	(0.03)	(1.19)	(0.11)
FP clinic		−0.1505	0.0199	−1.2964	2.1955
		(2.00)	(0.17)	(0.87)	(1.21)
FP clinic*year				0.0141	−0.0249
				(0.79)	(1.16)
FP clinic*primary edu				−0.2013	−0.4467
				(1.03)	(1.94)
FP province		−0.0706	0.7254	−0.0280	0.6746
		(0.11)	(0.71)	(0.05)	(0.66)
FP nation		0.8644	1.6496	2.7512	−0.8320
		(0.21)	(0.37)	(0.59)	(0.16)
1st baby female	−0.0919	−0.0914	−0.0956	−0.0883	−0.0901
	(1.51)	(1.50)	(1.37)	(1.45)	(1.29)
age at marriage	0.0398	0.0408	0.0396	0.0409	0.0396
	(3.27)	(3.34)	(2.72)	(3.35)	(2.72)
Muslim	−0.1261	−0.1042	−0.0308	−0.1109	−0.0481
	(1.30)	(1.01)	(0.16)	(1.07)	(0.25)
living w/ own parents	−0.0610	−0.0658	0.0223	−0.0655	0.0194
	(0.87)	(0.94)	(0.26)	(0.93)	(0.22)
living w/ parents-in-law	0.1913	0.1826	0.1852	0.1860	0.1821
	(2.37)	(2.26)	(1.93)	(2.29)	(1.90)
constant	2.0067	1.8221		3.5839	
	(3.27)	(0.71)		(1.12)	
no. of observations	6,127	6,127	6,114	6,127	6,114
no. of communities			274		274

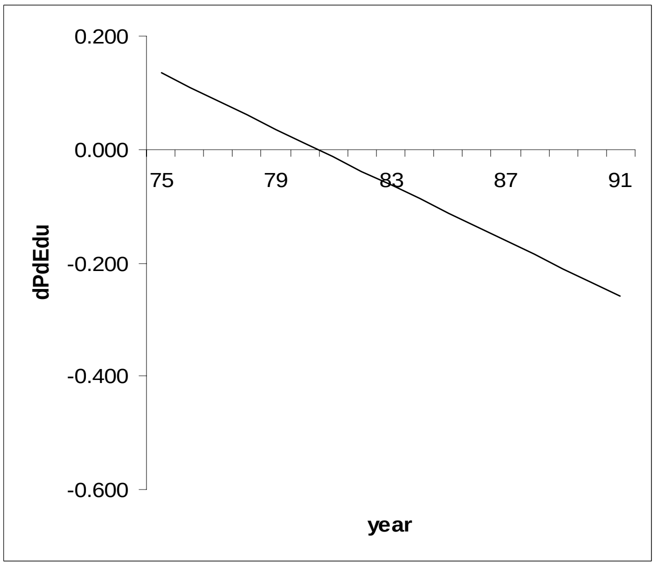
Notes: Absolute values of t-statistics are in parentheses. The data used are the 1993 Indonesian Family Life Survey.

Since a nonrandom placement of FP clinics is implied, a comparison of the results is made from two fixed-effects estimations, one without interaction terms of FP clinic (column (3) in Table 10) and the other with interaction terms (column (5) in Table 10), in order to consider the differential effect of FP program on educational groups. When the interaction terms are included, the coefficient on primary education and its interaction with year, indeed, become approximately 80 percent smaller in absolute value and are insignificant, both individually and jointly. The coefficient on FP clinic interacted with primary education is negative and significant at the 10 percent level, which is consistent with the hypothesis that more educated women are better at adopting new contraceptive methods. The decomposition of the marginal effect of women's primary education on birth hazard is demonstrated in Figure 9. Panel (a) shows the marginal effect of primary education on birth hazard corresponding to column (1) in Table 10. Using the coefficients in column (5) in Table 10, panel (b) plots the marginal effect over time of primary education evaluated at three different values of the FP clinic variable: 0, 1, and the average of the FP clinic variable in each year. The decomposition shows that the differential effect of the FP clinic on women with primary education can explain 77 percent of the total change of the marginal return to primary education in terms of the birth hazard over the period between 1974 and 1990.

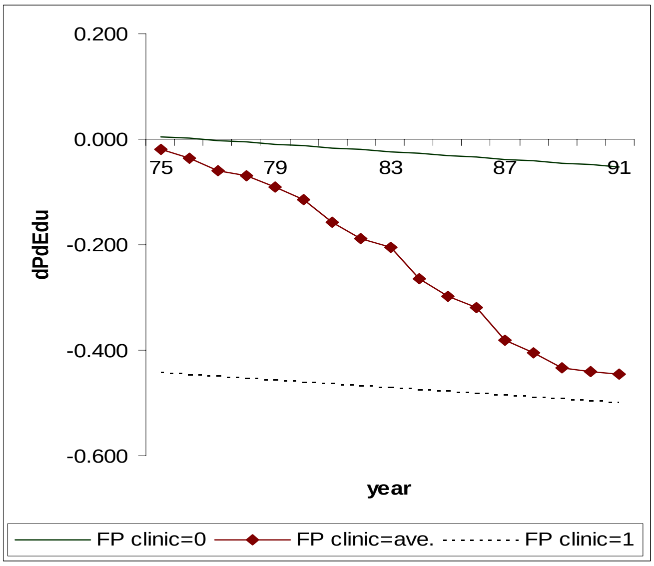
Although it is not clear if education enables women to have a better access to new information (as suggested by Thomas, Strauss & Henriques (1991)), or whether it actually enhances the ability to learn new technology (as suggested by T.W. Schultz (1975)), the finding is consistent with the hypothesis that more educated women are better at adopting modern methods of birth control. The policy implication derived from this result is that investment in women's education complements the Family Planning program.

However, Rosenzweig & Schultz (1989) conclude the opposite when examining the contraceptive use and its effectiveness in a sample of the United States. That is, they found that schooling does not have much return in the acquisition of information on contraceptives when new methods are easy to use, which implies that promoting female education is a substitute for a birth control information-dissemination program. However, their result is not necessarily contradictory to the one presented in this study. While the average years of schooling of women in their study, which used the 1973 National Survey of Family Growth, was 12.7 years, it is only 5.0 years for women in the IFLS 1993 in this study. Hence, their result can be interpreted as the marginal effect of schooling at the college level, whereas the marginal return to primary education is examined here. Therefore, the result in Table 8 together with the one in Rosenzweig & Schultz (1989) suggests that the marginal returns to schooling diminish as schooling increases.

Pitt, Rosenzweig & Gibbons (1993) suggest that, in the evaluation of FP program on fertility, the effects of other government programs should also be considered because these programs may come as a package. Therefore, two other health institutions are taken into account. They are the Integrated Health Post (Posyandu) and the community health center (Puskesmas). Although these two institutions are run by the Department of Health clinics, they function as a complementary to FP clinic on many occasions. In practice, around 75 percent of contraceptives are distributed



(a) Without interaction of education and FP clinic



(b) With interaction of education and FP clinic (FE)

Figure 9: Differential Effect of Family Planning Programs on Educational Groups

through Puskesmas in Indonesia. The same analysis is conducted, including the measures of Posyandu and Puskesmas. The results, presented in Table 11, are qualitatively the same as in Table 10. Comparing the specifications with and without the village fixed-effects (column (2) and column (3) in Table 11) suggests that the correlation between the village-specific propensity to conceive and placement of the FP clinic and Puskesmas is negative, and that the presence of Posyandu is associated with a higher propensity to conceive. With the inclusion of FP program variables interacted with education, the coefficients on education and its interaction with year become insignificant as shown in column (5) of Table 11. The FP clinic variable interacted with education is still negative and significant at the 5 percent level, while Posyandu interacted with education and Puskesmas interacted with education are not significant. Hence, the effects of these two health institutions seem to reinforce the differential effect of FP clinic on educational groups.

5 Conclusion

This study examines the effect of education on the birth spacing behavior of Indonesian women by looking at second birth intervals. Descriptive statistics and a duration analysis show that, among earlier cohorts, women who are more educated tend to have shorter birth intervals than the less educated and that the opposite is true among later cohorts.

There are three mechanisms through which a women's education affects birth spacing: the demand for children, the supply of children, and the cost of fertility control. On the supply side, the association between women's education and breastfeeding is proposed as an explanation for the relationship between women's education and second birth interval among earlier cohorts. For the reversal of the effect of education on birth spacing, two explanations are proposed. The first hypothesis, which involves the demand for children, relies on an increase in the return to schooling due to technological advance. The second hypothesis, which involves the cost of fertility control, is that there is a potentially superior ability to adopt the new technology of birth control due to education.

The empirical results based on first hypothesis shows that the level of industrialization of the local economy does not have a differential effect on birth spacing across educational groups. It also shows that occupational choice does not play a crucial role in explaining the change in the birth spacing behavior of educational groups over time, although women in the service industry experienced a change from shorter birth intervals to longer birth intervals over time. On the other hand, empirical results using measures of the availability of a family planning program show a differential effect on delaying births across educational groups. The point estimate is significant at the 10 percent level, and 77 percent of the total change of the marginal effect of primary education on birth hazard over time can be explained by the differential effect of the FP program on educational groups. The result still holds when the effects of two other health institutions are considered. Therefore, women's education had an impact on second birth interval mainly through changing the cost of fertility regulation rather than changing the demand for children over the period 1974 to 1990. This implies that family planning programs can actually have a big impact on fertility in a relatively

Table 11: Effect of Family Planning Programs II

	(1) Logit	(2) Logit	(3) FE Logit	(4) Logit	(5) FE Logit
age at 1st birth	−0.0445	−0.0433	−0.0496	−0.0436	−0.0502
	(3.62)	(3.53)	(3.36)	(3.54)	(3.39)
duration	0.3411	0.3460	0.5297	0.3518	0.5359
	(8.17)	(8.25)	(11.65)	(8.36)	(11.75)
duration^2	−0.0341	−0.0343	−0.0429	−0.0348	−0.0433
	(8.33)	(8.35)	(9.84)	(8.44)	(9.91)
year	−0.0392	−0.0333	−0.0527	0.0626	0.1057
	(5.24)	(0.85)	(1.23)	(1.01)	(1.53)
primary edu	1.9532	1.8439	2.3663	1.6911	1.4263
	(1.47)	(1.38)	(1.55)	(0.84)	(0.62)
primary edu*year	−0.0243	−0.0227	−0.0311	−0.0189	−0.0185
	(1.55)	(1.44)	(1.73)	(0.73)	(0.63)
husband's edu	0.0095	0.0107	−0.0033	0.0107	−0.0044
	(1.06)	(1.19)	(0.29)	(1.19)	(0.39)
FP clinic		−0.1402	0.0601	−3.3645	−0.2173
		(1.70)	(0.48)	(1.95)	(0.10)
FP clinic*year				0.0392	0.0046
				(1.90)	(0.18)
FP clinic*primary edu				−0.3118	−0.5646
				(1.43)	(2.20)
Posya		−0.0765	−0.1601	6.9355	6.1413
		(0.78)	(1.26)	(2.99)	(2.23)
Posya*year				−0.0844	−0.0755
				(3.05)	(2.30)
Posya*primary edu				0.3557	0.2956
				(1.40)	(1.03)
nPusk		−0.0371	−0.0017	0.1790	0.7963
		(1.46)	(0.03)	(0.40)	(1.40)
nPusk*year				−0.0023	−0.0092
				(0.44)	(1.40)
nPusk*primary edu				−0.0762	0.0145
				(1.27)	(0.20)
FP province		−0.0641	0.7470	−0.0990	0.5926
		(0.10)	(0.72)	(0.16)	(0.57)
FP nation		1.4305	1.5953	−6.7179	−11.2626
		(0.34)	(0.36)	(1.18)	(1.80)
1st baby female	−0.0979	−0.0949	−0.1023	−0.1045	−0.1030
	(1.59)	(1.54)	(1.45)	(1.69)	(1.46)
age at marriage	0.0403	0.0415	0.0399	0.0437	0.0416
	(3.28)	(3.39)	(2.73)	(3.55)	(2.83)
Muslim	−0.1577	−0.1160	−0.1312	−0.1404	−0.1689
	(1.58)	(1.10)	(0.66)	(1.32)	(0.84)
living w/ own parents	−0.0417	−0.0524	0.0376	−0.0508	0.0368
	(0.59)	(0.74)	(0.43)	(0.71)	(0.42)
living w/ parents-in-law	0.2027	0.1862	0.1802	0.1903	0.1825
	(2.48)	(2.27)	(1.86)	(2.31)	(1.88)
constant	1.9214	1.3104		−5.3916	
	(3.11)	(0.50)		(1.25)	
no. of observations	6,004	6,004	5,991	6,004	5,991
no. of communities			267		267

Notes: Absolute values of t-statistics are in parentheses. The data used are the 1993 Indonesian Family Life Survey.

low-educated population.

The finding that the FP program has a differential influence on educational groups is consistent with the hypothesis that more educated women are better at adopting new contraceptive technology. A policy implication is that investment in women's education complements the FP program. Rosenzweig (1995) discusses the cases where the returns to schooling are small for the acquisition of the simple new technologies. However, the result in this paper suggests that the returns to education at the primary level is significant, even if the (modern) methods of birth control are considered to have little scope for misuse. Further, together with the study on the United States sample by Rosenzweig & Schultz (1989), this result implies that the marginal returns to schooling may diminish as the level of schooling increases.

A natural extension of this study is to look at the first birth interval and to examine the effect of education on birth spacing with a structure of unobserved heterogeneity that associates first and second birth intervals. It would also be interesting to investigate higher order-birth intervals in order to see if own-experience complements education or not in the context of contraceptive technology. Although it did not turn out to be significant in the Indonesian case, assessing the effects of an increase in the return to schooling on birth spacing behavior in other countries remains a task for future research.

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Appendix

A Changes over Time by Women's Education

The descriptive statistics from Indonesian population census suggest that the decrease of the average number of children ever-born over time is parallel over educational groups as in Figure 10.

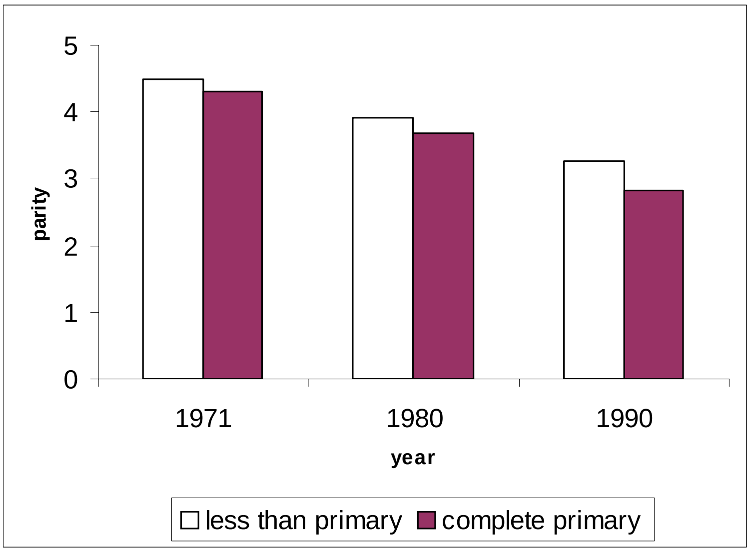
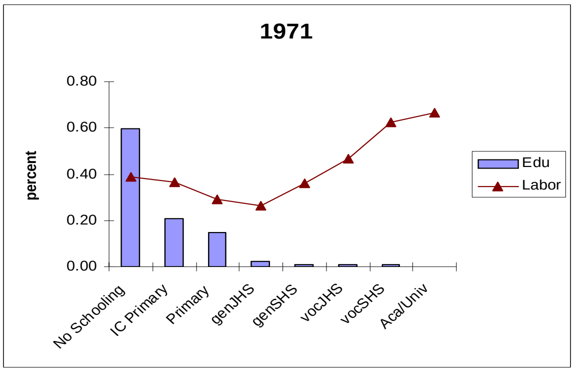


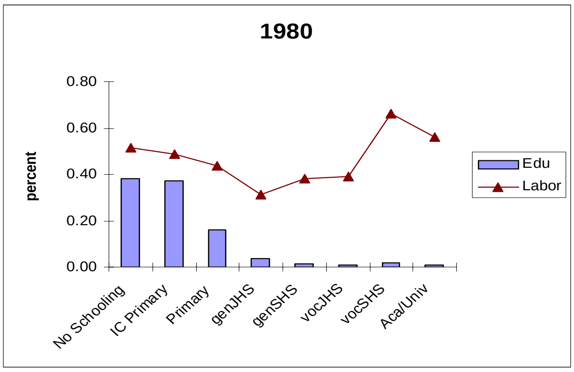
Figure 10: Average Children Ever-born of Married Women at Age 30

Notes: The data used are the 1971, 1980 and 1990 Indonesian Population Censuses.

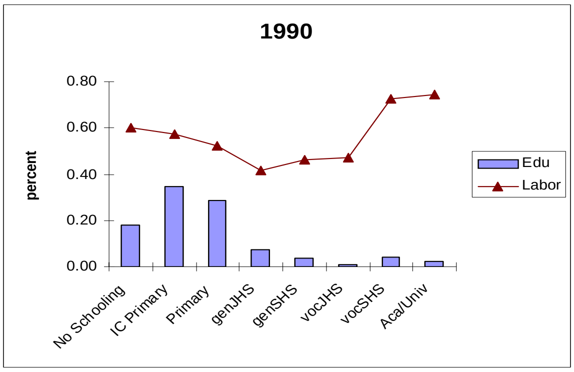
They also suggest that there is no differential change across educational groups regarding the inverse U-shape relationship between education and female labor market participation rate over the period 1970 to 1990 as in Figure 11.



(a) 1971



(b) 1980



(c) 1990

Figure 11: Female Education Distribution and Labor Market Participation Rate by Education for All Women at Age 30

Notes: The data used are the 1971, 1980 and 1990 Indonesian Population Censuses.)

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