Joint estimates of Education, Marriage, First Birth, and Labor Market Participation in Senegal [1]

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ABSTRACT

This paper jointly estimates the determinants of education, age at marriage, age at first birth, and labor market participation for young women in Senegal using a rich individual-level survey conducted in 2003. We use a multiple-equation framework that allows us to account for the endogeneity that arises from the simultaneity of the decisions that we model. Differences in the characteristics of the dependent variable informed the choice of the models that are used to estimate each equation: an ordered probit model is used to analyze the number of completed years of schooling, and a generalized Gompertz hazard model for the other three decisions. Results indicate that access to primary and secondary schools, along with school quality increases grade attainment; and each additional year of schooling delays marriage and the age at first birth by 0.5 and 0.4 years, respectively. Parents' education also reduces the hazard of marriage and age of first birth, while the death of parents has just the opposite effect, with the magnitudes of effects being larger for mothers. Delaying marriage also leads to an increase in the hazard of entering the labor market, as does the education and death of the women's parents.

Keywords: *Multiple equations; duration models; unobserved heterogeneity; Senegal.*

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1. Introduction

"Woman's growing independence, as a result of better education and improved career opportunities, is one of the main factors in the rise of delaying marriage, increasing marital instability and decline in fertility" (Blossfeld and Huinink, 1991). This quotation summarizes the common thinking about the relationships between these key outcomes in a woman's life course. The importance and complexity of the interactions between education, marital choices, childbearing, and labor market outcomes are widely acknowledged to be among the most critical set of life-course decisions of women. Understanding these relationships, however, is complex and difficult to quantify because of the need to address the issues of the endogeneity of decisions such as labor supply and marriage. In addition, it is necessary to deal with the related issue of unobserved heterogeneity; for example, heterogeneous preferences for market work and for children influence schooling and other investments in human capital. Tackling these problems is a formidable analytical challenge. At a minimum, it requires not only detailed event history data, rarely available in developing countries, but also a rich enough data set to allow us to exploit creative sources of identification.

In a recent article, Ganguli, Haussmann, and Viarengo (2011) present some stylized facts on education, marriage, fertility, and labor market outcomes using census data from a number of countries. They report evidence of a weak link between the increase in the average level of education and the increase in labor market participation. They also show that married women are less likely to participate in the labor market.

There is also considerable evidence of a robust negative association between female education and fertility (Schultz, 1997); although for African countries, it is common to find such a result only for high levels of education (Younger, 2006; Appleton, 1996; Thomas and Maluccio, 1996). Several explanations can be found for this negative relationship: woman's higher education increases the opportunity cost of childbearing (Becker, 1981), improves child health and reduces child mortality (Schultz, 1994), improves the knowledge of contraceptive methods (Rosenzweig and Schultz, 1985), or increases the female bargaining power in fertility decisions (Mason, 1986). Even if the correlation between education and fertility is robust, doubts can be expressed on the causality of this relationship: omitted variables, like unobserved preferences or household or community resources, can affect both schooling and fertility choices. Addressing this problem is a major challenge, with a good example of doing so being the work of Osili and Long (2008) who use the exposure to a program that involved investment in local schools in Nigeria as an instrument that is not related to fertility outcome.[3]

While education is often seen as crucial to success in the labor market, the challenge that pervades the literature on the interrelation between work and schooling is that the level of education is endogenous; this could lead to a bias in the estimated impact of education on employment, if not accounted for. This challenge has only been partially recognized, let alone adequately addressed. This applies to various strains of the literature, including that on child labor where school and work are often seen as alternative choices; and the literature that analyzes the relationship between education level and labor market participation for young women, who are supposed to make their participation decision after having acquired the desired schooling level. In the case of the child labor literature, Basu and Van (1998) argue that parents send children to school when wages are high enough for them to earn a living without resorting to children's labor to contribute to family needs. Dessy (2000) argues that there is a critical level of adult wages under which child labor is supplied. Empirical studies of

[3] Childbearing can also induce school interruption (Angrist and Evans, 1998), but we will not consider this hypothesis in this paper.

child labor, however, do not always support the implications of these theoretical models.[4] More relevant to our paper, there are a few studies that succeed in simultaneously modeling the schooling and labor decisions. Ray (2002) estimates a three-stage least squared model where child labor, child schooling, and household poverty status are jointly endogenous. Maitra and Ray (2002) jointly estimate child participation in schooling and employment using an ordered probit based on a ranking of the various outcomes: child schooling, employment, non-schooling, and non-employment. Kruger, Soares, and Berthelon (2007) model the parents' choice about children's work and schooling through a generalized ordered discrete choice model where schooling is preferred to the combination of work and schooling, and the latter is preferred to work alone. Levison, Moe, and Knaul (2001) employ a multinomial logit model to simultaneously estimate the determinants of participation in school, work, both, or neither. The use of such a model assumes the somewhat implausible IIA (independence of irrelevant alternatives); that is, the decision to work is not affected by the presence of a schooling option, and vice versa. Wahba (2006) strongly criticizes the use of both models: those that attempt to give an order to alternative choices, asin Kruger, Soares, and Berthelon (2007) and Maitra and Ray (2002), and the use of models that assume the IIA. She instead uses a bivariate probit model regression where schooling and work choices are correlated, as in Zapata, Contreras, and Kruger (2011).[5]

In the strand of literature that shares our focus on the participation of young women, the impact of education is often viewed as dependent on a constellation of factors that go beyond schooling, and instead reflect the cultural context and social norms regarding gender roles. There are two main pathways through which schooling positively affects participation:

[4] Duryea and Arends-Kuenning (2003) show that the incidence of child labor is higher and educational outcomes are lower in areas characterized by higher average wages. Kruger (2007) finds that, in coffee-producing regions, children are more likely to work (and less likely to go to school) during periods of coffee booms.

[5] Wahba (2006) finds that the correlation coefficient ρ is significant, implying that working and schooling are not independent, and thus, the choice of this estimation technique is appropriate. In addition, the correlation coefficient ρ is negative, indicating that there is a trade-off between child labor and child schooling choices.

through increasing wage offers, and through the expectation that education increases the bargaining power in the household (see, for example, Cameron, Dowling, and Worwick (2001)). Like the child labor literature, however, the empirical work on the impact of education on young women's entries into the labor market has failed to convincingly deal with the joint nature of these decisions. Thus, we are motivated in this paper to address this deficiency in the literature and the relationship between education and work by accounting for the endogeneity that arises from the simultaneity of the schooling and the labor market participation decisions that we model.

Beyond that, however, we are also interested in the links between marriage, fertility, and labor market participation, a literature that heavily relies on the seminal works by Becker (1973, 1981). The main strain of this literature is that the production of children is costly, especially for the mother; the increase in the value of mother's time as a result of increases in investments in education and career opportunities will affect the relative cost of children, thus reducing the demand for children. Given the complexity of estimating the shadow value of the time of a woman, many of the studies that have empirically tested Becker's theory use her education level as a proxy; these studies generally find that higher education levels are associated with lower fertility rates (Schultz, 1997). Other papers find more mixed results, including Liefbroer and Corijn (1999) who report mixed results of the effects of education and women's labor participation on marriage and childbearing in developed countries: in general, they report no effect on marriage timing on fertility, while observing a negative effect of education on the age of entry into motherhood. The main problem is that education (and labor participation) can be endogenous to the fertility choice: strong preferences for market work may induce women to invest more in education and to have fewer children.

Among the early work that attempts to explicitly address this problem is the study by Waite and Stolzenberg (1976), who hypothesize the existence of a simultaneous reciprocal causation between fertility expectations and labor force participation. According to the authors, there are background factors that completely account for this relationship. Browning (1992) subsequently reviewed the papers, written up to 1992, that try to access the issue of the endogeneity of fertility in the labor market supply and affirms, ". . . although we have a number of robust correlations, there are very few credible inferences that can be drawn from them," as with the literature on the impact of education on labor market outcomes (p. 1435). Another innovative approach to deal with this endogeneity problem is that of Horz and Miller (1988) who hypothesize that, in each period of time, parents choose a level of contraception and the amount of time a mother allocates to the competing needs of child care, homemaking, and work in the labor market. They use a four-equation system to model child care, labor participation, wages, and the use of contraceptives, and they find that there is a trade-off between the time spent for child care and the time spent in the labor market.

Life-cycle models have also been extensively used to model the relation between life-cycle fertility and female labor supply by others as well, including Heckman and Willis (1975) and Moffit (1984). In these models, the labor market decision is often modeled as conditional upon marital status. Van der Klaauw (1996) builds a dynamic utility maximization model using longitudinal data from a US panel survey, where he predicts changes in the life-cycle pattern of employment, marriage, and divorce due to differences in education, race, earnings, and husband's earnings. In the model, the sequential choices are interdependent, in the sense that all the previous choices have an effect on current choices and preferences; for instance, wages depend on previous work experiences. Similarly, Ma (2010) builds a structural dynamic model where, in each period over her life cycle, a woman maximizes the expected discounted value of her utility by simultaneously determining what category of occupation to enter (professional, non-professional, or housework), whether to be married, and whether to use contraception. These decisions depend on her previous career and family choices, and education is included as a control variable.

There are several papers which have tried to assess the endogeneity of fertility in the participation decision by adopting an instrumental variable approach. The main challenge is to find instruments that influence the fertility decision without directly affecting the participation decision. Rosenzweig and Wolpin (1980) rely on the use of sources of unplanned birth, like the presence of twins, while Rosenzweig and Shultz (1985) use the availability and cost of contraceptive technology. Similarly, Bailey (2006) uses the variation in the state-level legislation on access to the contraceptive pill, while Angrist and Evans (1998) use parental preferences for a mixed sibling-sex composition.

Less effort has been made to assess the endogeneity of marital status in the participation decision. Assaad and Zouari (2003) estimate a structural model that takes into account the endogeneity of fertility and of the timing of marriage in the participation decision. They model participation in the labor market as a polychotomous variable that distinguishes between different forms of participation. First, they estimate a duration model for age at first marriage and a probit model on the probability of being married, using the total number of married sisters as an exclusion restriction. Then, they estimate a negative binomial model on the number of children using the predicted age at first marriage as a regressor and the average contraceptive use by age at the provincial level as an instrument. As a third stage, they estimate a nested logit model on labor market participation, including the predicted probability of marriage and the predicted number of children among the regressors.

The works by Angeles, Guilkey, and Mroz (2005); Brien and Lillard (1994); Upchurch, Lillard, and Panis (2002); and Glick, Handy, and Sahn (2011) provide the most direct guidance for the approach we adopt in this paper. Angeles, Guilkey, and Mroz (2005) study the effect of education and family planning on fertility in Indonesia. They jointly estimate education decision, age at marriage, and fertility, using maximum likelihood procedures that assume that the heterogeneity terms of the three equations are correlated. The joint distribution of the unobservables is incorporated using a semi-parametric discrete factor method, as suggested by Heckman and Singer (1984) and extended by Mroz and Guilkey (1995) and by Mroz (1999). They report that, controlling for the endogeneity of education and marriage, the impact of the increase of education in reducing fertility is quite low. Conversely, they find that family planning services are very effective in fertility reduction in Indonesia.

Brien and Lillard (1994) study the interrelations between education, timing of marriage, and timing of first conception in Malaysia. They build a sequential probit model to estimate the schooling decision and model the timing of marriage and first conception through hazard models, allowing for correlation among the heterogeneity components of the three equations.[6] Identification is possible thanks to the hypothesis that, for each equation, there is a common heterogeneity component for sisters and that, conditional on this component, sisters' behaviors are otherwise independent. They find that education significantly delays the age at marriage, and that the increase in the age at first conception is due to the delayed marriage.

Brien, Lillard, and Waite (1999) estimate entry into marriage, cohabitation, and non-marital conception using a similar framework. They model each outcome with a continuous hazard model, and they account for the simultaneity of the three related processes. In this case, they are able to observe multiple episodes of each outcome for a subsample of women, and this allows for the identification of the degree of variation in individual specific components for each outcome and the correlation among those components. The same framework is used by Upchurch, Lillard, and Panis (2002), who estimate a model where education, marriage, and fertility decisions influence one another and where each outcome is affected by a woman's characteristics. They find that the risk of conceiving depends on

[6]This model is a combination of the simultaneous equations for hazard models by Lillard (1993) and the sequential choice model of education by Lillard and Willis (1994).

education for white and Hispanic women, but not for black women, while all women make simultaneous choices regarding childbearing and marriage.

More recently, Glick, Handy, and Sahn (2011) have examined the relationships among education, age at marriage, and age at first birth in Madagascar, with a multiple equation model, showing that schooling is very effective in delaying marriage and childbearing. Our paper is in many respects similar, with the major difference being that we not only jointly estimate the determinants of education, age at marriage, and age at first birth, but also age at entry in the formal labor market for young women in Senegal. We assume that there are common factors that influence the four behaviours we are analyzing. Some of these factors are observable, like parents' education and working experience, household wealth, characteristics of the place of residence, and so on. Some other factors are impossible to observe, like, for example, women's preferences.

Moreover, some of the outcomes of interest have a direct impact on some others. The higher the education level of the woman, for example, the more likely she is to postpone her marriage and thus her first childbearing. Education is also expected to facilitate entry into the labor market. At the same time, marital status is likely to influence the decision to enter the labor market. And, of course, marital status is expected to be a key determinant of childbearing.

In order to take into account for all these interrelations, we use a multiple equation framework. Differences in the characteristics of the dependent variable informed the choice of the models that are used to estimate each equation: an ordered probit model is used to analyze the number of completed years of schooling; hazard models are used to analyze age at marriage, age at first childbearing, and age at entry in the labor market. Our estimation approach is fully consistent with the theoretical description of the determinants of these key decisions, and it extends those adopted in Brien and Lillard (1994); Upchurch, Lillard, and Panis (2002); Angeles, Guilkey, and Mroz (2005); and Glick, Handy, and Sahn (2011).

As in Brien and Lillard (1994), we identify the covariance matrix of unobserved heterogeneity components, assuming that sisters share identical heterogeneity components for each equation. Moreover, identification of each equation is possible through the reliance on exclusion restrictions. We identify the completed years of education using detailed retrospective information on local schools and the age at marriage through the use of information on the marriage market. We use the information on the timing of the introduction of family planning programs in order to identify the age at first childbearing. Finally, we use the information on labor market shocks and changes in labor market characteristics to identify the decision to enter the labor market. In all these cases, we rely on retrospectively collected data, especially since many of the sources of identification, such as labor market entry, are dependent on events that transpire after schooling has been completed. While we are comfortable with these exclusion restrictions, even in the absence of instruments, we are able to deal with the unobserved heterogeneity through our estimation technique; our exclusion restrictions can thus be considered largely a bonus, as emphasized by Brien and Lillard (1994); Upchurch, Lillard, and Panis (2002); and Angeles, Guilkey, and Mroz (2005), and discussed further below.

In the remainder of this paper, we begin with a discussion of the empirical strategy in Section 2, followed by a presentation of the data we use in Section 3. Section 4 presents the main results, and Section 5 concludes.

2. Empirical Strategy

We simultaneously model four key decisions in a woman's life course: (i) the level of education, (ii) the age at marriage, (iii) the age at first child, and (iv) the age of the first entry in the formal labor market. We then estimate the same model replacing the age of the first entry in the informal labor market instead of the formal labor market.

Differences in the characteristics of the dependent variables inform the choice of the four models that are used to estimate each equation where we build upon the work of Brien and Lillard (1994); Upchurch, Lillard, and Panis (2002); Angeles, Guilkey, and Mroz (2005); and Glick, Handy, and Sahn (2011) to account for the endogeneity that arises from the simultaneity of the decisions that we model.

We model the number of grades completed for woman i living in community j, with an ordered probit model. An individual will have k grade of schooling, $G_{ij} = k$, if $\mu_k < G_{ij}^* < \mu_{k+1}$, where G_{ij}^* is the latent continuous variable that generates the observed G_{ij}, and μ_k and μ_{k+1} are the cut-off points to be estimated. Equation (1) describes the determinants of the latent variable G_{ij}^*

$$G_{ij}^* = \beta_0^G + \boldsymbol{\beta}_1^{G\prime}\boldsymbol{X}_{ij} + \boldsymbol{\beta}_2^{G\prime}\boldsymbol{C}_j + \boldsymbol{\beta}_3^{G\prime}\boldsymbol{E}_j^G + \varepsilon_{ij}^G \qquad (1)$$

$\boldsymbol{X}_{ij}^G$ is a vector of individual and household characteristics, including age information on the father's and mother's education, ethnicity, religion, region of birth, whether resident is in a rural or urban area, and whether the mother and father are dead. To avoid reverse causality, we also include information on the housing assets of a woman when she was 10 years old and information on the availability of health infrastructures, also measured when the woman was 10 years old. $\boldsymbol{C}_j^G$ is a vector of community-level factors, including the widespread access to electricity, the availability of piped water, and presence of three types of credit institutions—community-based small scale formal insurance institutions, insurance institutions that offer credit, and small individual lenders. $\boldsymbol{E}_j^G$ is a vector of school characteristics, including the share of local teachers in the closest primary school with at least five years of experience and the number of years of education of the school director. We also include information on access to primary and secondary schools when the woman was 10 years old, again since the choices about schooling investments were likely made on the basis of information on school access when children were young, not as adults when schooling is complete. We use the notation $\boldsymbol{E}_j^G$ to indicate a vector of variables that are included in the schooling model but excluded from the other equations: $\varepsilon_{ij}^G = u_{ij}^G + \eta_{ij}^G$, where u_{ij}^S represents unobserved characteristics which influence the grades completed by a woman i living in community j, and η_{ij}^S follows an identically and independently distributed normal distribution, i.e., $\eta_{ij}^S \overset{iid}{\sim} N(0, \sigma_s^2)$.

We model the age at marriage with a proportional hazard model, as it is depicted in Equation (2):

$$\ln h_{ij}^M(t) = \beta_0^M + \beta_1^M Age_{ij}(t) + \boldsymbol{\beta}_2^M{}' \boldsymbol{X}_{ij} + \boldsymbol{\beta}_3^M{}' \boldsymbol{C}_j + \beta_4^M G_{ij} + \beta_5^M{}' E_{ij}^M + \varepsilon_{ij}^M \qquad (2)$$

where $h_{ij}^M(t)$ represents the ratio between the probability of getting married at time *t* over the cumulative probability of not having married up to time *t*. The baseline hazard is represented by a generalized Gompertz model, which allows the baseline hazard rate to be a non-monotonic function of time; $Age_{ij}^M(t)$ is the piecewise linear duration dependency spline. The risk of marriage begins at age 11. Thus:

$$\beta_1^M(t) = \begin{cases} \beta_{11}^M t & if\ t \leq t_1 \\ \beta_{11}^M t_1 + \beta_{12}^M (t - t_1) & if\ t > t_1 \end{cases}$$

We select the time t_1 as the modal age at marriage in our sample, which stands at 17 years, so that we expect $\beta_{11}^M > 0$ and $\beta_{12}^M < 0$. The spline in age determines the woman- and time-specific impact of time *t* on the log hazard, which we denote as γ_{ijt}. The survival function, $S_{ij}^M(t)$, which denotes the probability of not having married up to time t, is given by:

$$S_{ij}^M(t) = e^{-\int_0^t h_{ij}^M(u)du} = e^{-\lambda(\gamma_{ijt})^{-1}\left(e^{\gamma_{ijt}t}-1\right)}$$

where $\lambda = e^{\boldsymbol{\beta}_2^M{}'\mathbf{X}_{ij}+\boldsymbol{\beta}_3^M{}'\mathbf{C}_j+\beta_4^M S_{ij}+\beta_5^M E_{ij}^M+u_{ij}^M}$. Other regressors include G_{ijt}, the number of completed grades of schooling, and E_{ij}^M, which represents the ratio of men to women in the same age cohort as i, the other covariates are as described above.[7] $\varepsilon_{ij}^M = u_{ij}^M + \eta_{ij}^M$, where η_{ij}^M,is an identically and independently distributed error term.

Similarly, age at first birth is modeled with the hazard model described in Equation (3):

$$\ln h_{ij}^P(t) \ = \beta_0^P + \beta_1^P Age_{ij}(t) + \beta_2^P Mar_{ij}(t) + \boldsymbol{\beta}_3^P{}'\mathbf{X}_{ij} + \boldsymbol{\beta}_4^{P'\mathbf{C}_j} + \beta_5^P G_{ij} + \boldsymbol{\beta}_6^P{}'\mathbf{E}_{ij}^P + \varepsilon_{ij}^P \qquad (3)$$

where $\ln h_{ij}^P(t)$ is the log-hazard of parenthood at time *t.* Risk of parenthood begins at age nine; $Age_{ij}^P(t)$ contains a node at age 21, that is the modal age at first child in our sample. We include multiple sources of duration dependence in the model, including $Age_{ij}(t)$ and $Mar_{ij}(t)$, which is the duration dependency spline indicating the marriage duration, with its coefficient allowed to change after three years since marriage.[8] $\mathbf{E}_{ij}^P$ is a vector of exclusion restrictions (including the availability of condoms in the community and the year when they were first available). $\varepsilon_{ij}^P = u_{ij}^P + \eta_{ij}^P$, where η_{ij}^P is an independently and identically distributed error term.

[7]Becker (1981) argues that, controlling for labor market opportunities, a larger relative supply of potential partners for women will raise their likelihood of marriage.

[8] We have tried including in the model a duration dependency spline indicating the time elapsed since labor market entry, but it turns out not to be significant.

Finally, Equation (4) presents the hazard model for age at entry in the formal labor market:

$$\ln h_{ij}^{L}(t) \;= \beta_0^L + \beta_1^L Age(t) + \beta_2^L Mar_{ij}(t) + \beta_3^L Par_{ij}(t) + \boldsymbol{\beta}_4^L \boldsymbol{X}_{ij} + \boldsymbol{\beta}_5^L \boldsymbol{C}_j + \beta_6^L G_{ij} + \boldsymbol{\beta}_7^L \boldsymbol{E}_{ij}^L + +\varepsilon_{ij}^L \qquad (4)$$

where $\ln h_{ij}^{L}(t)$ is the log-hazard of entry in the labor market at time *t.* Risk of labor starts at age five, and $Age_{ij}^{P}(t)$ contains a node at the modal age at entry in the formal labor market, that is 15 years. $\boldsymbol{E}_{ij}^{L}$ is a vector of dummies, indicating if a positive or a negative shock in the labor market occurred after leaving school; $Par_{ij}(t)$ is the duration dependency spline indicating time since first child. $\varepsilon_{ij}^{L} = u_{ij}^{L} + \eta_{ij}^{L}$, where η_{ij}^{L} is an independently and identically distributed error term.

We estimate a similar model for age at entry in the informal labor market. Its specification is the same as Equation (4), with the exception of $Age_{ij}\,(t)$, that contains a node at age 10, that is the modal age at entry in the informal labor market in our sample.

Individual unobserved heterogeneity poses two main challenges in our efforts to estimate these four interrelated outcomes. Observations with the same values for all covariates are not identical in terms of their hazards: some are more likely to experience failures than others because there are unobservables in the error term, ε_{ij}, that influence the decision processes that we are going to analyze. Ideally, we would identify the two different terms contained in ε_{ij}: the random error term (η_{ij}) and the unobserved heterogeneity component term (u_{ij}).

The second main challenge is represented by the fact that the unobserved factors that appear in Equations (1)–(4) are likely to be correlated, i.e., the same unobserved individual-specific characteristics simultaneously influence the four decisions we want to model. If this is the case, these influences give rise to an endogeneity problem. Consider, for instance, the completed grades of schooling S_{ij} which appear on the right side of Equations (2)–(4): G_{ij} is determined by ε_{ij}^{S}, thus whenever $corr\left(\varepsilon_{ij}^{S}, \varepsilon_{ij}^{k}\right) \neq 0$, with $k = M, P, L$, then G_{ij} will be correlated with the unobserved component of the error terms in Equations (2)–(4), and G_{ij} will thus be an endogenous regressor.

In order to deal with these challenges, we opt for an estimation strategy that is consistent with the fact that the decisions about schooling, marriage, childbearing, and labor market participation are interrelated, and we jointly estimate the four models. We consider these decisions as interrelated—in the sense that they are all influenced by individual characteristics, and that some of the endogenous outcomes of interest have a direct impact on other outcomes. Some of these characteristics are observed, while others are unobserved. We assume that, after conditioning for all observed variables, the heterogeneity term captures all sources of correlations among the four decision processes. The likelihood functions of each of the four models are independent if we are able to condition for the relevant observed and unobservable characteristics. If this is the case, the joint conditional likelihood of the set of observed outcomes for the four decision processes is the product of the conditional probabilities of the four models.

Identification of our four-equation system requires adding some structure on unobservable factors. Ideally, if we were able to repeatedly observe the choices made by a single woman under different observable conditions, we could control for the invariant unobserved component of the error term. The multiple outcomes per woman would allow separating the observation's specific heterogeneity component, (η_{ij}), from the random error term, (u_{ij}). Our data do not allow for such an ideal setting, so we need to introduce assumptions which allow us to identify the covariance matrix of unobserved heterogeneity components. Since we do not have repeated outcomes for the same individual, we assume that all the sisters living in the same household share identical heterogeneity components for each equation, as in Brien and Lillard (1994). This is a reasonable hypothesis since sisters are exposed to the same family circumstances and come from the same background, i.e., the same social context and the same value system. This assumption allows for the estimation of the degree of variation in the sisters-specific component for each process and the correlation among these components.

As described above, each equation contains exclusion restrictions, i.e., covariates that are included only in one equation and are excluded from the others. But the inclusion of these exclusion restrictions would not be essential in order to identify the model (Lillard, 1993; Brien, Lillard, and Waite, 1999).

2.1. Marginal effects

The computation of the marginal effects of any regressor needs to account for both the direct effect on each of the four models, as well as for the indirect effects that go through the outcomes of earlier models, as our four-equation system is recursive.

Let $\boldsymbol{Z}_{ijt}$ denote all the possibly time-varying regressors which are included in at least one of the four models; without loss of generality, assume that the first element in $\boldsymbol{Z}_{ijt}$ is a continuous variable.[9] The partial derivative of the predicted number of completed grades $E\left(G\middle|\boldsymbol{Z}_{ij}\right)$, with respect to Z_{1ij}, is given by:

$$\frac{\partial E\left(G\middle|\boldsymbol{Z}_{ij}\right)}{\partial X_{1ij}} = \sum_{k=0}^{N} \frac{\partial Prob\left(G = k\middle|\boldsymbol{Z}_{ij}\right)}{\partial X_{1ij}} k \; = \sum_{k=0}^{N} \beta_1^G \phi\left(\mu_k - \boldsymbol{\beta} \mathrm{Z}_{\mathrm{ij}}^{\mathrm{G}}\right)$$

[9]If one element of this vector is excluded from model h with $h = G, M, P, L$, then its coefficient is constrained to zero.

Then, we compute the impact of the marginal variation in $\boldsymbol{Z}_{ijt}$ upon the median age at marriage[10] as predicted by the duration model in Equation (2). This, in turn, depends on (i) the direct impact captured by the estimated coefficient in Equation (2); and (ii) the impact which goes through the influence of the variation in $E(G|\boldsymbol{Z}_{ij})$ and the related impact on the out-of-school spline.

Similarly, the influence of a marginal variation in $\boldsymbol{Z}_{ij}$ upon the parenthood model in Equation (3) has to account for its influence upon the timing of the marriage. To give an idea of the richness of these indirect effects, we can observe that the influence of, say, the death of the father upon the age of entry in the labor market of a woman in the sample depends on 17 coefficients estimated in the four models.

3. Data Sources and Descriptive Statistics

The data we use in this paper is the 2003 Household Survey on Education and Welfare in Senegal (EMBS), conducted in 33 rural and 30 urban communities.[11] The sample consists of 2,668 females between the ages of 15 and 30 years of age. Although, as discussed by Glick and Sahn (2009, 2010), the sample is not truly nationally representative; it is part of a cohort study of young children, and efforts were made to randomly select into the sample new households to ensure that it is as close as possible to a random sample. Indications from comparison with other national surveys indicate that this effort was quite successful, and that the sample of 1,820 households is representative of the population in terms of religion, ethnic groups, and demographic characteristics, as well as other characteristics such as education.[12]

[10]This is defined as the predicted age at marriage at which $S_{ij}^{M}(t)$ is equal to 0.5; no closed form expression exists for the mean time to failure predicted by a Gompertz model.

[11] See Glick and Sahn (2009) and (2010) for details about the survey design.

[12]For example, net primary enrollment in our sample (primary enrollments of children 7–12) is 66 percent compared with 63 percent for the country as whole in 2000 (World Bank, 2006).

In our analysis, we rely extensively on the education, labor market, and demographic modules of the EMBS, as well as the module which contains information on the current residence, and on retrospective questions for adults above age 21 about where they lived, as well as the household and community characteristics, when they were 10 years old. This data is a key component of our methodology, because it allows us to observe the childhood characteristics that we use to explain the marriage, fertility, and labor market decisions. We also use the community and school modules that collect detailed information on the local infrastructure in general, as well as the characteristics of schools in the community. This includes the experience and credentials of the principal and management of the school, as well as the number of teachers, their qualifications and pedagogical practices, and building and classroom conditions. Up to three schools were interviewed in the 60 clusters (i.e., communities) where the survey was conducted. These are used as control variables in the analysis that follows. In addition, we have information about the availability of family planning services and the availability of contraceptive devices, all of which were collected as part of the community survey. In each community, it was determined whether each service type was available and, if so, when it first became available.

Descriptive statistics for variables used in this paper are reported in Table 1 and indicate that the average age of women in our sample is 20.52, that 32 percent of the sample is married, and 24 percent are parents. Birth before marriage is not very common, occurring for only 6 percent of our sample.

Women in the sample have an average of 4.28 years of schooling, although, 25 percent of the women are still in school. Thus, mean completed school among this cohort will be higher. Only 36 percent of their fathers and 23 percent of their mothers attended some school.

In terms of labor market activities, we distinguish between those women working in the formal wage and those working in the non-wage labor market. Around 18 percent of our sample report working in the formal sector wage labor market, and the average age of entry is 17. Those working in the informal non-wage sector comprise 27 percent of the women in our sample.

About half of our sample is rural, and it is 96 percent Muslim. The largest ethnic group is Wolof, comprising 38 percent of the sample, followed by the Poular and Serere, each representing approximately one-fifth of the sample. Ninety-percent of the communities have a primary school, but only 61 percent have a lower secondary school. Among other community infrastructure, around four-fifths of the communities have piped water, electricity, and readily available condoms at health clinics, with the average year that condoms became available being 1994. Approximately 60 percent of the communities report having access to each of the following types of credit sources: credit communautaire, insurance, and individual lenders. However, among our communities, 25 percent have none of these sources, 24 percent have only one of these sources, and 51 percent have all three of these sources of credit.

4. Results

We present the results of the education, marriage, fertility, and labor market models in Table 2. Since the parameters themselves are often difficult to interpret, we focus on the marginal effects and highlight some of those results in Tables 3 and 4.

4.1 Education

We find that the education of the mother and father have a powerful impact on schooling attainment of their daughter. Interestingly, the magnitude of the coefficients of father's education is substantially higher than that of the mother. When we compute the marginal effects, we find that if the mother has some primary schooling, the impact on her daughter's education is modest, just 0.03 years. However, if the mother has completed primary school, the daughter is likely to have completed 0.67 additional years of school. If the mother has at least completed lower secondary school, the predicted effect on schooling of the daughter just about doubles to 1.4 years of schooling. The impact of a father's education is much greater; a daughter of a father who has completed primary school will have 1.2 additional years of schooling, with the comparable number of lower secondary schooling being 1.9 more years for his daughter, as compared to a father with no schooling. The death of a father before the child turns 10 reduces the expected years of schooling by around 0.45 years; no such effect is observed for the death of a mother, as indicated by the insignificant parameter estimate. One plausible explanation for this is that the father's death has a greater impact on household resources and thus contributes to earlier school withdrawal. However, we would have expected that girls substitute in terms of home production when their mother dies.

The household asset index when the girl was 10 years of age also has the expected positive impact on schooling outcomes. As mentioned above, relying on this lagged asset variable avoids the possibility of any reverse causality and more accurately reflects how wealth at or around the time that a child is just enrolling in school affects long-term schooling outcomes. In terms of magnitudes, we find that an increase in the asset index of one standard deviation contributes to 0.24 year increase in the number of years of schooling.

The presence of a primary school within 5 km, at age 10, has a large impact. A child living in such a community is expected to have completed 2.8 more years of school than a child living in a community without such a school. The presence of a lower secondary school within 5 km, at 10 years of age, has about one-third the impact on schooling as the presence of a primary school.

Among schooling characteristics, we examine the impact of the education of the director and the experience of the teachers. In the case of the former, our results indicate that each year of additional education of the director increases the expected level of education by 0.04 years. Similarly, a 10-percentage point increase in teachers with at least five years of experience raises the expected years of schooling by 0.65 years. Some caution is warranted in interpreting these school characteristics variables since they may in part be capturing the effect of the wealth and other general characteristics of village; thus it is possible that community heterogeneity is driving the results. While we cannot rule this out, we include a range of other community covariates to at least deal in part with this problem. Included are indicator variables that capture whether most of the households have access to electricity, whether there is piped water available, the distance to land-line telephones, and the presence of various credit institutions, as well as access to health facilities when the women were 10 years old. These are largely intended as controls, and thus caution is necessary in interpreting the individual parameters in a causal fashion.

Among other marginal effects of note is that Muslim girls are expected to complete 0.9 less years of schooling that other religious groups, and among the Diola ethnic group, 0.7 years more than the predominant Wolof ethnic group. Being born in certain regions also results in far lower schooling achievement. For example, those from Diourbel realize 0.68 fewer years of schooling, and conversely, those born in Louga complete more schooling than the region of Dakar.

4.2 Marriage

In considering the determinants of marriage, we again concentrate our discussion on the presentation of the total marginal effect, which includes both the direct effect as well as the indirect effect of the parameters, operating through the impact on schooling. More specifically, the indirect schooling effect is reflected in both the grade attained and the number of years out of school spline. Additionally, the normal aging process that is captured separately by the age splines affect the timing of marriage. One of the most notable results is that a women's own education reduces substantially the hazard of marriage. We find that each additional year of schooling contributes to a decline in in the median marriage survival time by approximately half a year. Other individual characteristics that seem to be important in terms of affecting the marriage hazard include that of being a Muslim, which decreases the survival time to marriage by a great deal, 3.76 years; being a member of the Pular ethnic group also increases the hazard of marriage in a substantial way.

In examining the impact of the characteristics of the household in which the woman lived when she was a child, we find that the women's mother and father having some education increases the survival probabilities; but the deaths of a mother and father have the opposite effect, raising the hazard of entering into marriage. More specifically, the time to marriage among women in our sample is increased by 4.01 years when their mothers have some education, nearly four times the magnitude of the effect of their fathers having some education. The effect of the death of a mother reduces the survival function by nearly two years. Interestingly, we find little effect of a father's death on hazard of marriage.

Another interesting finding is that the higher the ratio of men to women, the higher the hazard of getting married. This is consistent with our expectations insofar as the more men relative to women in the local marriage market, the shorter the survival time to marriage.

4.3 First birth

Figure 1 presents the survival functions for marriage and age at first birth for a rural women with the following individual and household characteristics: Wolof, Muslim, 22 years of age, did not complete first grade, mother and father had no education, the mother died, and the asset index value was 11.48 (equivalent to 0.5 standard deviation above the mean). The survival probability function for marriage crosses the 50 percent probability line around 21 years of age, and by their late twenties, the vast majority of women with these simulated characteristics are expected to have married. It is about that same age of 20 that a woman's hazard of first birth begins to accelerate rapidly, and by around age 23—that is, during the first three years of marriage—nearly all the women with the simulated characteristics are predicted to have their first birth.

Like with marriage, the women's own education and that of her parents is of critical importance in terms of the timing of first births. An additional grade of education increases time to first birth by 0.41 years. The mother having some primary education in turn delays the age of first birth by 1.8 years, which is over two times the magnitude of the effect on the father having some education.

In modeling the hazard of first birth, we find that the coefficients on the time since marriage, as entered as a spline for the first three years, and on more than three years of marriage, have the expected positive sign and are statistically significant at standard levels. This can be interpreted as suggesting that the hazard of having a first birth increases with time during the first three years of marriage, and thereafter, the hazard remains the same, although, the survival function continues its downward trend. In terms of the magnitude of the impact, we calculated the marginal effect and find that delaying marriage by one year (i.e., decreasing marriage duration by one year) reduces the median parenthood age by 0.61 years. This is portrayed in the survival function shown in Figure 2.

The death of the women's a mother also has a large impact in terms of increasing the hazard of first birth, comparable in magnitude to the effect on age at marriage. The death of a father has a smaller impact than the death of a mother, as it did with the hazard of marriage, but the delay of first birth by 0.66 years is much higher than the 0.08 years that the hazard of marriage is reduced.

We also find that the availability of condoms in the community reduces the hazard of first birth. The marginal effect implies that first birth is delayed by 0.88 years when they are available. However, in those communities where condoms have been available longest, condom availability has the least impact on the hazard of motherhood. This might be explained by the fact that there is a great influence with the more recent introduction of condoms, both because of their novelty, as well as the possibility that more recent efforts at condom diffusion are more effective in terms of broad-based behavioral change.

4.4 Labor market entry

The magnitude of the marginal effect of own education is relatively small, with each additional grade of schooling delaying entry by 0.18 years into the formal wage sector, and 0.09 years in the case of the informal sector. This presumably reflects that more education is associated with higher expectations for a good wage job, and thus, longer queuing times for entry into the formal sector, in particular. Such a finding has been observed in other countries in Africa where education has been shown to be associated with greater queuing for higher paid formal sector employment opportunities (Glick and Sahn, 1997).

One interesting aspect of the story about the role of education in terms of delaying entry into employment is observed in Figures 3 and 4 where we draw the survival function for entry into the labor market. In the case of the formal wage labor market (Figure 3), by age 25, approximately 15 percent of the women with the simulated characteristics would be expected to have entered into the formal sector labor market. For those having an additional year of schooling, this figure would be closer to 13 percent. What is particularly interesting in this figure is that we distinguish between the direct and indirect influence of grade attainment. In this case, the direct effect, also shown in the figure, is very small. Instead, most of the impact of grade attainment on the risk of labor market entry passes through the influence of schooling on the timing of marriage and parenthood. This in turn explains why the two survival functions are nearly identical up until the age when a woman is making decisions regarding marriage and fertility. In the case of the non-wage labor market, we see that the direct and indirect effects of education are working in opposite directions: the direct effect contribution to an increase in the hazard of work, while the indirect effect is doing just the opposite (Figure 4).

We also find that the hazard of entering the formal labor market increases as a result of a positive economic shock having occurred in the previous six years, conditional upon having completed schooling. Presumably, this reflects the greater employment opportunity in the wage sector that follows from such as shock once a woman has exited school. There is no such effect observed in terms of entering the non-wage sector, which again is not surprising since it is in part a reservoir for those unable to find better jobs in the wage sector. Interestingly, just the opposite applies to negative shocks that occur after leaving school: such a shock is not associated with any change in the hazard of entering wage employment, but it is associated with a increase in the risk of entering the non-wage sector. We have no strong expectation as to the impact of such events since, on the one hand, a negative shock would be expected to increase the need for a woman to find a job to cope with the stress of the shock, but at the same time, a negative shock may reduce the employment possibilities in the formal wage sector due to increased economic circumstances. The marginal effects also show that a woman whose father or mother has some primary schooling is associated with an increased hazard of entering the formal sector. The marginal effects of a father having some primary education are particularly high: the duration to entry into the formal labor market is 3.33 years shorter if the father has some education, more than twice the magnitude of the effect of the mother's education. Just the opposite is the case for the role of father's education in the informal sector. Perhaps this reflects that these men are less likely to be working in the informal sector, and that this in turn reduces the likelihood that their daughters would be engaged in informal sector work. The death of a mother and a father also delays entry into the formal labor market, just like it does for marriage and age of first child.

We also consider how the time since marriage affects the timing of entry into the labor market. The negative and significant effect of being married for 0–3 years indicates that the hazard of entering the formal and informal sector (the change in the instantaneous probability) decreases with time during the first three years of marriage, and thereafter does not change (as seen by the insignificant parameter estimates for marriage duration of greater than three years). In terms of magnitude, delaying marriage by one year reduces the expected duration until entering the formal labor market by 0.31 years.

The negative coefficients on the motherhood duration splines also suggest that the hazard of entering the labor market decreases as a consequence of motherhood, although, it is only the parameter estimate for motherhood of two or more years that is significant in the informal sector model.

Among the other covariates, Muslim women have the expected lower risk of entering the wage labor market. In fact, the duration until formal labor market entry is expected to be 3.46 years longer for Muslim than non-Muslim women. This is illustrated additionally in Figure 5 which shows the survival function for labor market entry distinguishing between Muslim and non-Muslim women. Muslim women enter the labor market earlier, but by age 19 the survival functions cross and thereafter, the risk of labor market entry for non-Muslim women accelerates, while this is not the case for Muslims. So, by ago 30, there is a large difference in the accumulated risk of having entered the labor market, with Muslim women far less likely to do so.

As with the other models, we include a range of community covariates, largely as control variables. We do take particular note of the impact of accessibility of credit institutions in terms of increasing the hazard of job entry. Access to these credit institutions, as expected, showed no direct effect on marriage or fertility choices, although, access to a

"Mutuelle" did contribute to increased schooling, which would be consistent with the notion that credit constraints may be one explanation for early dropout.

4.5 Heterogeneity correlations

Finally in Table 2, we present the heterogeneity standard deviations and correlations. All of the standard deviations are significant, and the correlation between marriage and parenthood is positive. This later relationship presumably reflects the fact that those women, who are inclined to putting off marriage, do the same for first birth.

5. Conclusions

In this paper we have estimated the complex relationship between schooling, marriage, age at first birth, and entry into the labor market, employing a method that is designed explicitly to control for the endogeneities that result from the fact that these outcomes are jointly determined. Our goal is to gain some insight into the links between these simultaneously determined outcomes and to better understand the importance of individual attributes, family background, and communities' characteristics.

In terms of the determinants of schooling, we highlight the importance of factors when the child was 10 years of age in determining eventual schooling attainment. This includes the parent's education and the wealth of the parent's household. Additionally, we find that the accessibilities of primary and secondary schools at age 10, and their characteristics, are important in determining schooling outcomes. So, too, is the availability of community infrastructure, such as health services, when the cohort of women we study were young children.

While understanding the factors that determine school attainment is important, much of our paper focuses on the effect of schooling on the other outcomes we model. We find that grade completed is an important in terms of delaying marriage, with the marginal effect on age at first birth being approximately half of marriage. This effect on fertility is operating through delayed marriage, as well as directly on the decision to have a first birth. Education also eases entry into the wage labor market, again with the effect operating indirectly through marriage and fertility choices, as well as directly on the labor market entry decision. Beyond own education, mother's education is also important for determining age of marriage and first birth. Father's education has a lesser effect on the survival function. Death of mothers and fathers is also important in affecting all the outcomes of interest, and like education, the impact of the death of a mother is far greater than the father. Women who grew up in household with more assets also are less likely to marry at younger ages, but increase the hazard of labor market entry as well as age at first birth.

Several other important results emerge from our analysis, including the role of an individual's religious background and the gender ratios in the community in affecting the time to marriage; condom availability in reducing the hazard of first birth; and the importance of positive economic shocks and availability of credit institutions in terms of promoting labor market entry.

Finally, the significant correlations and standard deviations across equations indicate that there are endogeneities that need to be dealt with in modeling these outcomes, and thus the importance of our efforts to deal with them through jointly modeling these outcomes. Our findings therefore adds to the limited set of studies that take into account that the outcomes of schooling, marriage, fertility, and labor market participation are likely to be correlated and which address the endogeneities and unobserved heterogeneity that have implications for the modeling of these outcomes.

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Table 1. Mean, minimum and maximum values of variables

Variable	Mean	Min	Max
Grade completed	4.28	0	14
Age	20.52	15	30
Muslim	0.96	0	1
Ethnicity, Wolof	0.38	0	1
Ethnicity, Poular	0.20	0	1
Ethnicity, Serere	0.18	0	1
Ethnicity, Dioola	0.06	0	1
Ethnicity, Mandingue	0.13	0	1
Father dead	0.22	0	1
Mother dead	0.07	0	1
Father has no education	0.62	0	1
Father has primary education	0.06	0	1
Father has completed primary	0.16	0	1
Father has completed college	0.15	0	1
Mother has no education	0.76	0	1
Mother has primary education	0.07	0	1
Mother has completed primary	0.12	0	1
Mother has completed college	0.05	0	1
Father has no education	0.62	0	1
Household Asset index at age 10	46.63	0	100
Rural area	0.48	0	1
Distance to phone	1.10	0	15
>75% of Household Use Electricity	0.81	0	1
Pipeline network	0.82	0	1
Health Service within 5km at 10 years	0.81	0	1
Credit communautair	0.61	0	1
Credit, insurance	0.60	0	1
Credit, individual lender	0.57	0	1
Primary school within 5km at 10 yrs	0.90	0	1
Lower secondary within 5km at 10 yrs	0.61	0	1
Number years school of primary school Director	13.33	10	17
% teachers' in primary school with at least 5 years experience	0.69	0	1
Men to women ratio in cohort	0.81	0.53	0.97
Condoms available	0.80	0	1
Year condoms first available	1994	1972	2003
No positive shock after leaving school	0.43	0	1
Positive shock within 0-3 years of leaving school	0.09	0	1
Positive shock, 3-6 years after leaving school	0.10	0	1
Positive shock, more than 6 years after leaving school	0.38	0	1

Source: EMBS 2003.

Table 2. Education, marriage, first birth and labor market entry joint estimation results

	Formal labor market		Informal labor market	
Completed grades				
Age	-0.012 *	(1.72)	-0.012 *	(1.70)
Muslim	-0.418 ***	(2.77)	-0.411 ***	(2.72)
Father dead	-0.212 ***	(2.97)	-0.210 ***	(2.95)
Mother dead	-0.005	(0.04)	-0.011	(0.09)
Father education, some primary	0.230 *	(1.72)	0.224 *	(1.70)
Father education, primary completed	0.560 ***	(6.11)	0.563 ***	(6.11)
Father education, college completed	0.886 ***	(8.06)	0.882 ***	(8.01)
Mother education, some primary	0.015	(0.12)	0.014	(0.11)
Mother education, primary completed	0.308 ***	(2.99)	0.307 ***	(2.94)
Mother education, college completed	0.628 ***	(4.38)	0.609 ***	(4.25)
Asset index	0.009 ***	(5.48)	0.010 ***	(5.58)
Rural	0.426	(1.64)	0.429	(1.64)
Distance to telephone from community	-0.017	(1.09)	-0.017	(1.08)
At least 3/4 hhs use electricity in community	0.109	(1.19)	0.106	(1.16)
Pipeline network in the community	0.354 ***	(2.99)	0.354 ***	(2.97)
Health service at 5 Km (dummy)	0.002	(0.02)	-0.002	(0.03)
Credit communautaire <5km	-0.038	(0.28)	-0.027	(0.20)
Mutuelle de credit <5km	0.224	(1.64)	0.221	(1.61)
Preteur individuel <5km	0.426 **	(2.38)	0.426 **	(2.37)
Primary school within 5km at age 10	1.517 ***	(11.38)	1.510 ***	(11.34)
College within 5 km at age 10	0.435 ***	(5.30)	0.435 ***	(5.33)
Years of education of the school director	0.02	(0.93)	0.02	(0.94)
Percentage of teachers with at least 5 years of exp.	0.307 *	(1.67)	0.311 *	(1.69)

continued

Table 2 *cont.* Education, marriage, first birth and labor market entry joint estimation results

	Formal labor market		Informal labor market	
Age at Marriage	Coefficient	p-value	Coefficient	p-value
Age spline, age 11 intercept	-12.715 ***	(5.02)	-12.830 ***	(4.99)
Age 11-17 slope	0.655 ***	(14.62)	0.658 ***	(14.37)
Age 17+ slope	0.183 ***	(6.45)	0.186 ***	(6.58)
Age	0.125 ***	(2.61)	0.125 ***	(2.60)
Education grade completed	-0.138 ***	(3.94)	-0.135 ***	(3.83)
Muslim	0.926 **	(2.00)	0.925 **	(2.02)
Father dead	-0.028	(0.22)	-0.028	(0.21)
Mother dead	0.535 ***	(2.69)	0.540 ***	(2.64)
Father education, some primary	-0.236	(0.73)	-0.246	(0.76)
Father education, primary completed	-0.288	(1.36)	-0.306	(1.44)
Father education, college completed	-0.237	(0.89)	-0.232	(0.86)
Mother education, some primary	-1.012 ***	(3.34)	-1.042 ***	(3.44)
Mother education, primary completed	-0.656 **	(2.34)	-0.646 **	(2.32)
Mother education, college completed	-0.903 **	(2.22)	-0.939 **	(2.32)
Asset index	0.002	(0.47)	0.002	(0.54)
Rural	0.532	(1.08)	0.569	(1.14)
Distance to telephone from community	0.061 **	(2.45)	0.062 **	(2.49)
At least 3/4 hhs use electricity in community	-0.21	(1.20)	-0.2	(1.13)
Pipeline network in the community	-0.442 **	(2.36)	-0.450 **	(2.36)
Health service at 5km (dummy)	-0.214	(1.39)	-0.219	(1.42)
Credit communautaire <5km	-0.144	(0.58)	-0.137	(0.54)
Mutuelle de credit <5km	-0.384	(1.54)	-0.394	(1.54)
Preteur individuel <5km	0	(0.00)	0.008	(0.03)
Ratio of men to women in the cohort	3.804 **	(2.36)	3.838 **	(2.34)

continued

Table 2 *cont.* Education, marriage, first birth and labor market entry joint estimation results

	Formal labor market		Informal labor market	
Age at first child	Coefficient	p-value	Coefficient	p-value
Age spline, age 9 intercept	88.606 **	(2.37)	91.875 **	(2.45)
Age 9-21 slope	0.464 ***	(9.48)	0.463 ***	(9.60)
Age 21+ slope	0.075	(1.52)	0.08	(1.61)
Marriage duration spline, intercept	1.555 ***	(6.40)	1.560 ***	(6.36)
Marriage duration 0-3 years	0.960 ***	(7.88)	0.988 ***	(8.06)
Marriage duration 3+ years	0.078	(1.06)	0.09	(1.19)
Age	0.009	(0.39)	0.014	(0.62)
Education grade completed	-0.033	(0.65)	-0.031	(0.59)
Muslim	-0.126	(0.23)	-0.192	(0.34)
Father dead	0.573 ***	(2.98)	0.594 ***	(3.06)
Mother dead	0.492 *	(1.96)	0.498 *	(1.93)
Father education, some primary	-0.044	(0.09)	-0.033	(0.07)
Father education, primary completed	0.465	(1.52)	0.462	(1.49)
Father education, college completed	0.098	(0.25)	0.075	(0.19)
Mother education, some primary	0.307	(0.71)	0.293	(0.67)
Mother education, primary completed	0.289	(0.68)	0.301	(0.69)
Mother education, college completed	-0.32	(0.44)	-0.25	(0.33)
Asset index	-0.004	(0.87)	-0.005	(0.94)
Rural	0.379	(0.53)	0.32	(0.44)
Distance to telephone from community	0.023	(0.65)	0.024	(0.66)
At least 3/4 hhs use electricity in community	-0.539 **	(2.13)	-0.502 **	(1.99)
Pipeline network in the community	-0.715 **	(2.24)	-0.764 **	(2.32)
Health service at 5km (dummy)	-0.295	(1.32)	-0.317	(1.40)
Credit communautaire <5km	0.517	(1.33)	0.486	(1.22)
Mutuelle de credit <5km	0.004	(0.01)	0.007	(0.02)
Preteur individuel <5km	0.283	(0.60)	0.243	(0.50)
Year preservatives were first available	-0.049 ***	(2.63)	-0.051 ***	(2.71)
Availability of preservatives in the community	-0.699 **	(2.20)	-0.742 **	(2.28)

continued

Table 2 *cont.* Education, marriage, first birth and labor market entry joint estimation results

	Formal labor market		Informal labor market	
Age at entry in the labor market	Coefficient	p-value	Coefficient	p-value
Age, intercept	-9.258 ***	(8.43)	-6.836 ***	(7.09)
Age 5-10 (or 5-15) slope	0.296 ***	(9.31)	0.896 ***	(15.34)
Age 10+ (or 15+) slope	0.191 ***	(7.29)	0.086 ***	(4.27)
Marriage duration spline, intercept	0.547 *	(1.87)	0.832 ***	(2.83)
Marriage duration 0-3 years	-0.373 **	(2.14)	-0.513 ***	(2.71)
Marriage duration 3+ years	-0.035	(0.45)	0.137	(1.08)
Motherhood duration spline, intercept	-0.118	(0.16)	0.07	(0.08)
Motherhood duration, 0-2 years	-0.352	(0.39)	-0.133	(0.12)
Motherhood duration, 2+ years	0.039	(0.54)	-0.248 *	(1.78)
Age	-0.035 *	(1.70)	-0.066 ***	(3.57)
Education grade completed	0.007	(0.18)	-0.019	(0.47)
Muslim	-0.318	(0.88)	-0.029	(0.07)
Father dead	0.183	(1.07)	0.069	(0.39)
Mother dead	0.367	(1.47)	-0.331	(1.26)
Father education, some primary	0.423	(1.28)	-0.228	(0.67)
Father education, primary completed	0.088	(0.37)	-1.123 ***	(4.00)
Father education, college completed	-0.523	(1.53)	-1.527 ***	(3.98)
Mother education, some primary	-0.181	(0.52)	-0.184	(0.47)
Mother education, primary completed	0.142	(0.49)	0.187	(0.62)
Mother education, college completed	-0.815	(1.59)	0.069	(0.12)
Asset index	-0.003	(0.57)	0.001	(0.13)
Rural	1.928 ***	(3.45)	0.871	(1.43)
Distance to telephone from community	0.044	(1.39)	0.246 ***	(8.27)
At least 3/4 hhs use electricity in community	0.265	(1.12)	0.760 ***	(2.64)
Pipeline network in the community	-0.595 **	(2.18)	-1.061 ***	(4.68)
Health service at 5km (dummy)	-0.211	(1.13)	-0.013	(0.07)
Credit communautaire <5km	0.544 *	(1.70)	-1.712 ***	(5.28)
Mutuelle de credit <5km	0.943 ***	(2.78)	0.514 *	(1.74)
Preteur individuel <5km	0.917 **	(2.50)	-0.596 *	(1.71)
Positive Shock 0-3 years after leaving school	0.751 ***	(2.60)	-0.27	(0.93)
Positive Shock 3-6 years after leaving school	0.987 ***	(3.46)	-0.364	(1.44)
Positive Shock 6 years or more after leaving school	0.659 ***	(2.97)	-0.294	(1.35)
Negative Shock 0-3 years after leaving school	0.011	(0.03)	0.857 **	(2.16)
Negative Shock 3-6 years after leaving school	0.333	(1.07)	0.076	(0.25)
Negative Shock 6 years or more after leaving school	-0.156	(0.76)	-0.031	(0.15)

Table 2 *cont.* Education, marriage, first birth and labor market entry joint estimation results

	Formal labor market		Informal labor market	
Heterogeneity Standard Deviation				
Schooling	0.694 ***	(9.61)	0.693 ***	(9.71)
Marriage	1.362 ***	(9.89)	1.385 ***	(9.82)
Parenthood	2.212 ***	(9.42)	2.277 ***	(9.83)
Labor	1.625 ***	(8.43)	2.220 ***	(15.68)
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Correlations				
Schooling and Marriage	0.049	(0.35)	0.034	(0.25)
Schooling and Parenthood	-0.073	(0.61)	-0.08	(0.68)
Schooling and Labor	-0.021	(0.16)	-0.077	(0.91)
Marriage and Parenthood	0.288 ***	(4.09)	0.284 ***	(4.18)
Marriage and Labor	0.039	(0.43)	-0.007	(0.12)
Parenthood and Labor	0.105	(1.34)	-0.01	(0.18)
Log likelihood	-12163.75		-12711.5	
[1] Wolof is the reference category				
Control for Region of birth included in all models				

Table 3. School attainment model marginal effects		
Variable	Marginal effect	Discrete change
Age	-0.026	\\\
Muslim	\\\	-0.917
Father dead	\\\	-0.448
Mother dead	\\\	-0.011
Father education, some primary	\\\	0.498
Father education, primary completed	\\\	1.235
Father education, college completed	\\\	1.984
Mother education, some primary	\\\	0.032
Mother education, primary completed	\\\	0.671
Mother education, college completed	\\\	1.393
Asset index	0.019	\\\
Rural	\\\	0.881
Distance to telephone from community	-0.036	\\\
At least 3/4 hhs use electricity in community	\\\	0.234
Pipeline network in the community	\\\	0.749
Health service at 5 Km (dummy)	\\\	0.004
Credit communautaire <5km	\\\	-0.081
Mutuelle de credit <5km	\\\	0.481
Preteur individuel <5km	\\\	0.923
Primary school within 5km at age 10	\\\	2.789
College within 5 km at age 10	\\\	0.973
Years of education of the school director	0.043	\\\
Percentage of teachers with at least 5 years of exp.	0.654	\\\

Table 4: Marginal effects of hazard models

Effect in years on:

Variable	**Age at Marriage**	**Age at first child**	**Age at entry in formal labor market**	**Age at entry in informal labor market**
Grade (1 additional grade)	0.5	0.41	-0.18	-0.085
Father dead (from 0 to 1)	-0.08	-0.66	-1.16	-0.199
Mother dead (from 0 to 1)	-1.89	-1.51	-1.88	3.75
Father has primary education (compared to no education)	1.06	0.76	-3.33	2.01
Mother has some education (from 0 to 1)	4.01	1.8	-1.54	-0.5
Age (1 additional year)	-0.43	-0.24	0.3	1.03
Ratio of men to women in the cohort (1 Standard deviation=.132)	-1.71	-0.88	0.39	1.29
Years preservative were first available (1 Standard deviation=7.33 years)	//	0.44	//	//
Availability of preservatives in community	//	0.81	//	//
Delaying Marriage (1 year)		0.61	-0.31	-0.64

Characteristics: 22 years old, Muslim, Wolof, Mother died, Father no educated, Mother no educated, Asset index 11.48, lives in Rural area, Zero education grades completed.

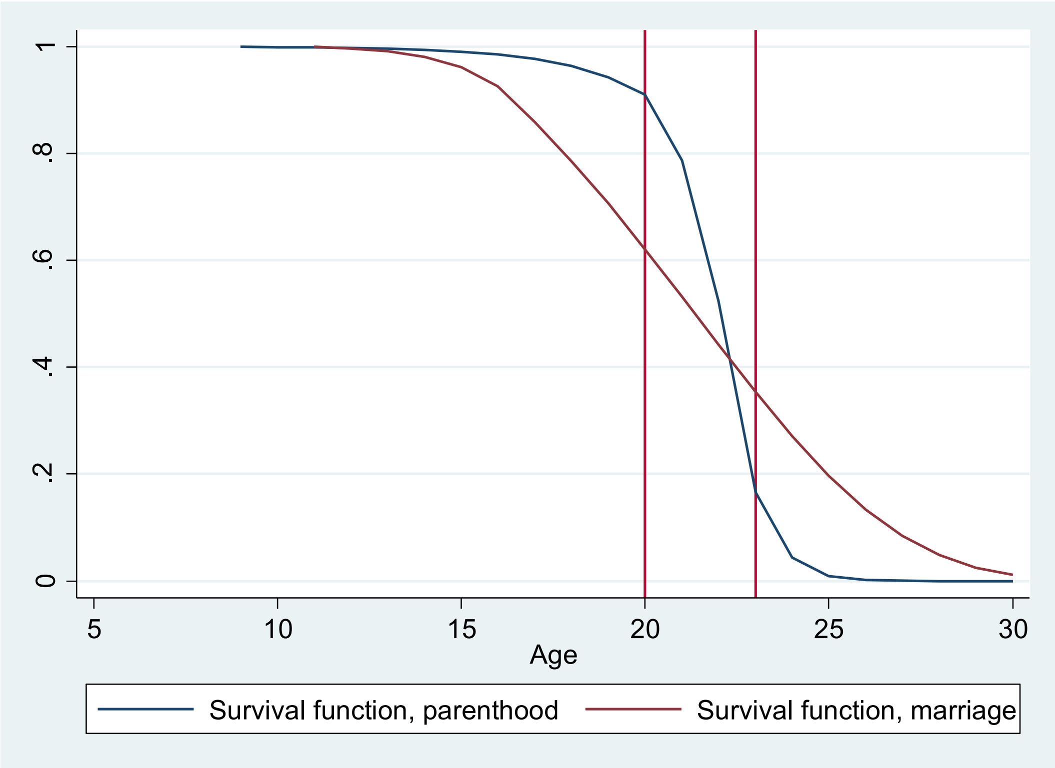

Figure 1. Survival Function of marriage and parenthood

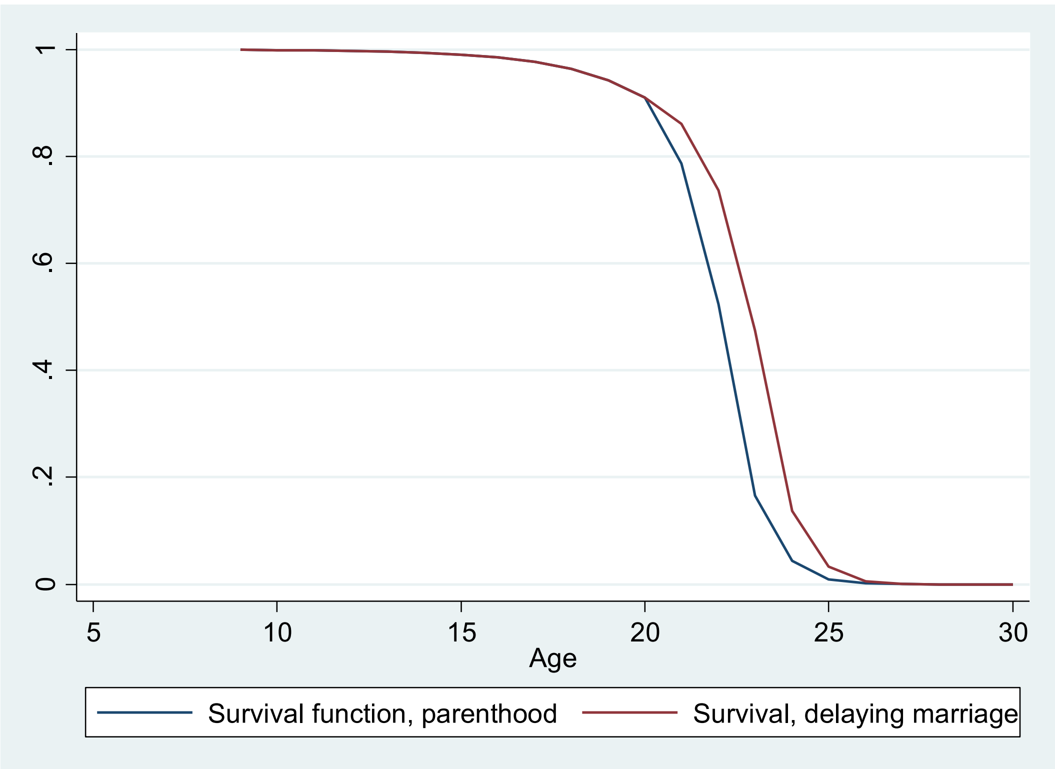

Figure 2. Survival function for parenthood, delaying marriage

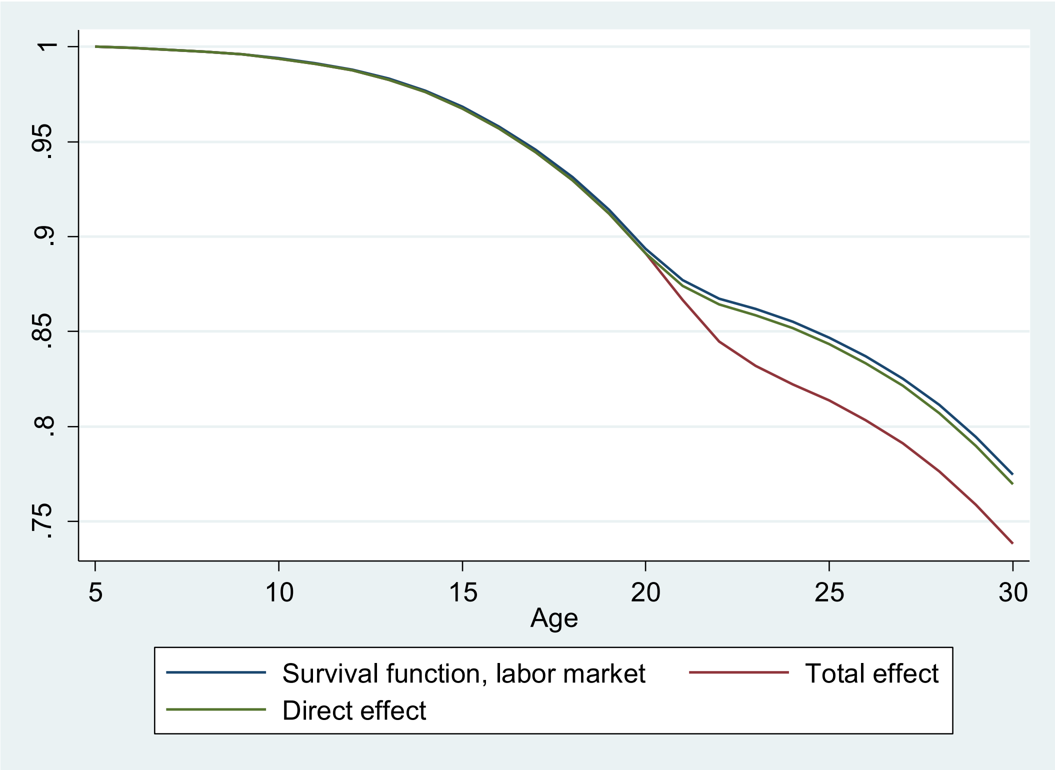

Figure 3. Simulation of the increase of 1 standard deviation of grade (3.5 grades) in formal sector.

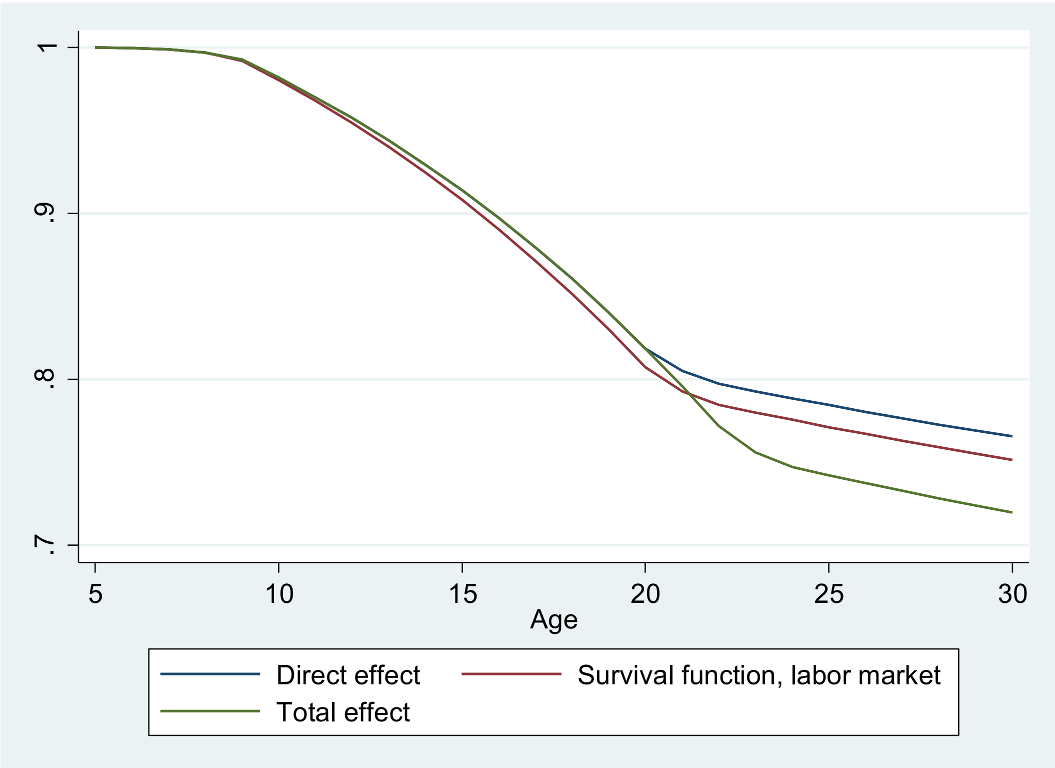

Figure 4. Simulation of the increase of 1 standard deviation of grade (3.5 grades) in informal sector.

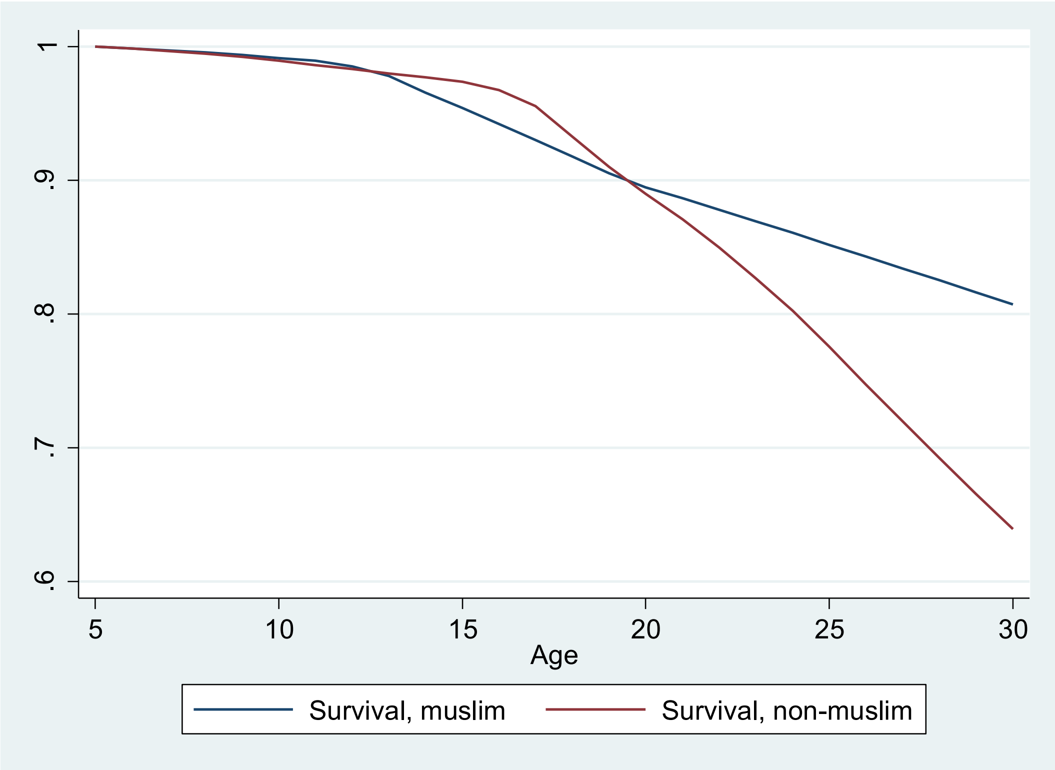

Figure 5. Survival function of formal sector labor market entry by religious group