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Article — Published Version Education and the Timing of Family Formation: Evidence from Quantile Regression Analysis

Development and Change

Provided in Cooperation with: John Wiley & Sons

Suggested Citation: Batyra, Ewa (2024) : Education and the Timing of Family Formation: Evidence from Quantile Regression Analysis, Development and Change, ISSN 1467-7660, Wiley Periodicals, Inc., Hoboken, NJ, Vol. 55, Iss. 5, pp. 1018-1050, https://doi.org/10.1111/dech.12846

This Version is available at: https://hdl.handle.net/10419/313792

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Education and the Timing of Family Formation: Evidence from Quantile Regression Analysis

Ewa Batyra 匝

ABSTRACT

The relationship between educational level and the age at which women start families has been extensively researched. However, studies have primarily explored how additional schooling shifts the mean or, more broadly, only one point of the age at first union and first birth distributions. This ignores variation in the association between education and the timing of family formation, and the fact that schooling might shape behaviours of vulnerable and more privileged women differently. Using quantile regressions, this article examines heterogeneity in the relationship between education and the age at first union and first birth across the distribution of these events within 50 low- and middle-income countries. It investigates whether additional schooling shifts relatively early union formation and childbearing (that is, lower parts of distributions) similarly or differently than it shifts other parts of the distributions. It finds that association between an additional year at school and the age at first union and birth is weaker in the lower than the upper parts of the distributions. Education has a relatively weak effect on the reduction of early first unions and births and plays an unequalizing role in shaping family formation within countries. These findings are key to understanding persistently high levels of early marriage and pregnancy, despite the expansion of education.

INTRODUCTION

The positive association between the level of women's education and the age at which they start forming families is one of the most studied relationships in demography, sociology and development studies. To date, however, researchers have focused on examining this association at only one point of the distributions of the age at first union and the age at first birth,

The author acknowledges funding from the Marie Sklodowska-Curie Action Individual Fellowship from the European Commission (Grant Number 101060237-DISPARFAM). She thanks the anonymous referees for their helpful comments on earlier drafts of this paper. Open access funding enabled and organized by Projekt DEAL.

Development and Change 55(5): 1018-1050. DOI: 10.1111/dech.12846

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limiting our understanding of variation in how education is related to the timing of family formation within populations. Existing studies ascertain how additional education (for example, years of schooling completed or a given educational level) shifts the mean and the median of the age at first union and the age at first birth distributions, or the proportion of women who experience these events by a given age (for example, first union before age 18), while disregarding the rest of the distributions of these events (e.g., Bongaarts et al., 2017: Esteve and Florez-Paredes, 2018: Grant, 2015: Gupta and Mahy, 2003; Ikamari, 2005). These approaches, while providing valuable information. do not account for the fact that estimates of the effect of education on, for example, the mean of the age at first union might not necessarily be indicative of the size and nature of these effects in other parts of the age at first union distribution (for example, in the lower or upper parts of the distribution). Thus, existing studies have paid little attention to the possibility that the association between education and the timing of family formation might not be the same for all individuals within countries.

In places where women start families early, such as in many low- and middle-income countries (LMICs), evidence of a positive association between schooling and the timing of family formation has led researchers and policy makers to emphasize the importance of expanding access to education and generating incentives for parents to send girls to school, not only for the sake of human capital improvements, but also because of education's capacity to delay the age at first union and birth. Early family formation has long been a concern thanks to numerous studies documenting its negative effects on women's well-being and that of their children (Sunder, 2019; Urdinola and Ospino, 2015; Yount et al., 2018). This, in essence, means that the lower parts of the distributions of the age at first union and birth (that is, early unions and births) are associated with higher risks for well-being. These findings, together with growing international attention to policies concerning the minimum age at marriage, motivated researchers to scrutinize the determinants of union formation and transition to parenthood by the age of 18. However, from the theory and policy perspective, it is relevant to ask not only how education affects one point of the age at first union or birth distributions, but also whether the association is similar, weaker or stronger in the lower than the upper parts of the distributions of these events. So far, there have been no attempts to answer such questions and to measure links between education and the age at first union or birth across the distributions of these events. This is an important issue in the context of LMICs given the high levels of early marriage and pregnancy, despite the expansion of education.

Relatedly, there is a growing body of research documenting stark polarization and wide disparities in the timing of family formation within countries, which raises questions about the factors contributing to this heterogeneity. Based on this evidence — discussed in depth in the following section — I hypothesize that education may be not uniformly associated with age at first union and birth within countries. Namely, additional schooling may shift lower parts of the distributions of these events (that is, relatively early union formation and childbearing) less than it shifts the upper parts of these distributions (that is, relatively late union formation and childbearing). To examine this proposition, I take a novel perspective by using quantile regression analysis. This methodological approach allows me to take account of the fact that schooling might have a heterogeneous association with the timing of family formation within populations, so that the strength of this association may differ across the quantiles of the distributions of the age at first union and birth.

To offer a comprehensive analysis, this study takes a cross-country perspective and focuses on 50 nations in Africa, the Americas, Asia and the former Soviet Republic. In examining whether education affects relatively early union formation and childbearing similarly or differently than it affects other parts of the distribution of these events, I conduct robustness checks, including using women's height as a proxy for socio-economic background that may be associated with both educational level and women's family formation trajectories. Overall, by showing that the strength of the association with education is weaker in the lower parts of the age-at-first union and birth distributions, this study highlights that schooling might provide insufficient impetus for the reduction of early unions and childbearing, and delivers important insights regarding policy formation.

BACKGROUND

Numerous studies, both descriptive and focused on estimating causal effects, provide evidence that schooling influences family formation behaviour, ranging from fertility desires (Behrman, 2015) and levels (Jejeebhoy, 1995), to the timing of events such as the age at first union and first birth. The latter outcomes have received particular attention as they mark the beginning of the family-formation process. Although the relationship between the level of education and the age at which women start forming families differs in magnitude between countries, there is overwhelming evidence that women with higher levels of schooling are less likely to have their first union and first child early, compared to women with a lower level of education.¹

Existing studies have employed a number of methodological approaches; nonetheless, these can be classified into two broad groups. The first, descriptive approach, relies on the level of education to disaggregate the

^{1.} There is an extensive literature on this, but see, for example, Bongaarts et al. (2017); De Paoli (2011); Ferré (2009); Frye and Lopus (2018); Gupta and Mahy (2003); Gyimah (2018); Ikamari (2005); Jejeebhoy (1995); Li and Cheng (2019); Lloyd (2005); Melesse et al. (2021); Stoebenau et al. (2021).

mean or median age at women's first union or birth (e.g., Bongaarts et al., 2017; Stoebenau et al., 2021) or the proportion of women who have their first union or child by a specific age, for example, by the age of 18 (e.g., Esteve and Florez-Paredes, 2018; Melesse et al., 2021). The second common approach, which controls for additional factors associated with the main outcome and predictor, involves using regressions. Some studies have used ordinary least square regression (OLS) or related two-stage least square regression to model shifts in the mean of the distribution, where the dependent variable is the age at first union or birth (e.g., Grant, 2015; Ikamari, 2005; Li and Cheng, 2019). Others have used probit and logit regressions to model shifts in one particular outcome, sometimes within the instrumental variables framework whereby the dependent variable is an indicator of the timing of first union or birth (for example, describing whether a woman experienced these events before age of 18) (e.g., De Paoli, 2011; Ferré, 2009; Gupta and Mahy, 2003).

These studies have delivered important insights by focusing on policyrelevant age thresholds, and some are also able to account for potential endogeneity between education and the timing of family formation by using causal inference techniques. At the same time, these methodological approaches and ensuing results provide only a partial view of the link between the two processes because they focus on the association at just one point of the distributions of the age at first union and birth. While providing valuable information, existing studies have not paid attention to the fact that the relationship might vary at different points in the distributions of these outcomes and, thus, that the association between education and the timing of family formation might not be uniform within countries. For example, schooling might not be contributing to the reduction of early first unions and births to the same extent that it contributes to shifting the average age upwards, or to shifting the age at which women with relatively late pattern of family formation experience these events.

There are reasons to expect that the association with education might differ across the distributions of the age at first union and birth. Overall, education may have heterogeneous effects on different outcomes across the population, ranging from economic outcomes (Brand and Xie, 2010) to demographic outcomes like realized fertility (Brand and Davis, 2011), planned fertility (Miranda, 2008) or female autonomy (Urbina, 2022). For example, Brand and Davis (2011) found variation by socio-economic background in the effect of college education on the timing of first birth and completed fertility in the USA. Miranda (2008) showed that, among teenagers in Mexico, more education was associated with a decrease in planned fertility, but this link was not uniform across the population and the association was most pronounced among women with strong preferences towards children. Urbina (2022) found that in Colombia, Peru and Bolivia, in the context of high levels of gender inequality and prevailing hypergamy norms, the association between additional schooling and female autonomy varied across the population, and that among more vulnerable women, schooling in fact curtailed autonomy. Overall, as highlighted by Brand and Xie (2010), individuals are likely to differ in the way they respond to common treatments and thus it is important to study variability, particularly with respect to the role of schooling.

Based on this knowledge, and evidence related to the patterns of age at first union and birth in LMICs, one might expect that not all individuals will respond in the same way in terms of the timing of family formation when receiving more years of education. First, although women who spend more vears in school form families later than those with fewer years of education. the educational expansion in many LMICs has been accompanied by a much less pronounced delay in the age at first union and birth than expected (or, at times, no delay at all) (Lloyd, 2005; Mensch et al., 2005). For example, in spite of school attendance already being high during the 1990s, during the 2000s and 2010s Latin American countries continued to evidence a low mean age at first birth and exceptionally high levels of adolescent fertility. especially vis-à-vis the total fertility rate (Rodríguez Vignoli, 2013, 2014). In West Africa, educational expansion occurred along with an increase in the mean age at first union, but the region continued to have one of the highest levels of child marriage in the world (Koski et al., 2017; Pesando et al., 2021).

One of the hypothesized explanations is that the expansion of schooling did not result in the provision of good-quality education for all students (Bongaarts et al., 2017; Esteve and Florez-Paredes, 2018; Grant, 2015; Lloyd, 2005; Mensch et al., 2005). Despite increasing educational participation, the percentage of students completing primary school with basic literacy has remained low in many countries (Psaki et al., 2018): 9 out of 10 children in low-income countries were reportedly unable to read by the age of 10 (World Bank, 2019) and overall global progress in learning has been limited (Angrist et al., 2021; Pritchett, 2013; World Bank, 2018). One of the contributing factors has been increasing enrolment putting a strain on educational systems and not being accompanied by sufficient increases in resources, including qualified teachers. In some countries, unsustained investments in education, especially in the context of economic downturns and structural adjustment programmes during the 1990s, played a role (Lloyd et al., 2000, 2005), and had lasting impacts on social spending. Moreover, in the context of high levels of economic inequality, as seen in many LMICs but particularly in Latin America, the opportunities to access high-quality education have been unevenly distributed (Daude, 2013), and have resulted in vast differentials in the skills and knowledge that students gain.

This evidence suggests that not all students have been able to benefit from spending more time in school in terms of what they have been learning and, subsequently, in terms of future labour market prospects — one of the channels through which education influences family formation. It can thus be

hypothesized that in contexts where some girls and young women attend school but gain little knowledge, these women might not alter their family formation behaviours as much as their peers who benefit more from schooling. In other words, additional schooling may shape the behaviours differently across the population. This could result in the persistently high levels of early unions and births among more vulnerable women (that is, unions and births in the bottom of the age at first union and birth distributions), even if changes in family formation associated with receiving more years of schooling are taking place in other parts of the distributions (for example, in the median or upper parts).

There are other, related channels through which education may have a heterogeneous association with the timing of family formation. Education may shape union formation and transition to parenthood not only through knowledge, skills and future opportunities but also by imparting new ideas which challenge traditional gender norms (Caldwell, 1997). However, there is evidence that information received at school may be insufficient to induce a meaningful change in perceptions about women's status in society. In some schools, the curriculum and teachers' practices may themselves be gendered and thus reinforce traditional social norms (Aikman et al., 2005; Stromquist, 2001). Similarly, additional schooling may have unexpected implications and result in decreased female autonomy due to prevailing gender role imbalances in society (Urbina, 2022).

Relatedly, in settings with high levels of gender inequality, the effects of education may be limited among some women, as school attendance is seen, including by girls' families, as a means to prepare them to be better wives and mothers (Kabeer, 2005). Strong norms in relation to women's role in society, in terms of reproduction and marriage, may mean that the behaviour of some women is hardly affected or remains unchanged despite school attendance. Overall, as Frye and Lopus (2018) argue, the association between education and, in the case of their analysis, union formation will depend on the social meaning attached to schooling. Finally, in settings with high levels of socio-economic inequality, the degree of interaction between different social groups, by which norms and behaviours spread across the population, may be limited. For example, although individuals across all social strata may be attending school, even the same school, teachers may treat students differently depending on their background, thus reinforcing existing social divisions (Kabeer, 2005).

In short, not all students may benefit from the transformative potential of schools. High levels of early union and childbearing, issues around disparities in education quality and labour market prospects, and variation in the meanings attached to schooling and its power to transform norms, are some of the explanations for education's heterogeneous association with the timing of union formation and childbearing within LMICs; this association might be stronger for some women and weaker for others. In fact, there is a growing body of research documenting a stark polarization in the timing of family formation in many LMICs. Studies have highlighted the coexistence of very early and very late family-formation patterns within populations, a fact exacerbated in contexts of high economic inequality such as Latin America and Africa (Batyra and Kohler, 2022; Batyra et al., 2021; Castro Torres et al., 2022; Lima et al., 2018; Stoebenau et al., 2021). This evidence of the diversity of family-formation trajectories merits further investigation: to what extent are the commonly studied determinants of age at first union and birth contributing to this diversity? While the above studies have documented disparities in the timing of events within countries, they have not explored the possibility that additional schooling may have a heterogeneous association with the timing of family formation.

This article aims to provide a comprehensive view on the topic by focusing on 50 countries. I use quantile regression analysis to answer the central question: does an additional year at school shift the 25th percentile of the age at first union and birth distributions (representing relatively early union formation and childbearing) equally, more, or less than it shifts the 75th percentile (capturing relatively late union formation and childbearing)? The article also explores the extent to which the patterns differ between Africa, the Americas, Asia and the former Soviet Republics.

By employing a cross-country perspective, this article provides the broadest possible view of the variation in the studied associations. Because of its large scale, it cannot offer an analysis of causal effects or account for potential endogeneity in the link between education and the timing of family formation. Rather, it focuses on the associations themselves, accounting for potential factors that may be linked with both education and the timing of family formation. Despite this, the study contributes to a better understanding of one of the most researched associations in demography, sociology and development studies by highlighting the importance of considering distributional heterogeneity when studying links between education and family processes. It also opens important avenues for future investigation on the extent to which these associations may be causal.

DATA

This study is based on data from the Demographic and Health Surveys (DHSs), which collect data from more than 90 countries worldwide, and are nationally representative of women aged 15–49.² It focuses on 50 countries grouped into five broad world regions (see Table 1). In the process of selecting the surveys for this study, those that did not have the requisite information were excluded, that is, those without information about the age at first union or birth, and for which only ever-married women

^{2.} See: https://www.dhsprogram.com/

Americas Bolivia 2008 1958 1973 5529 Colombia 2015 1965 1981 13243 Dominican Rep. 2013 1963 1978 3208 Guatemala 2014–15 1965 1980 7724 Guyana 2009 1959 1974 1886 Guyana 2001 1951 1966 3841 Peru 2012 1962 1977 8941 Asia Cambodia 2014 1964 1977 6672 India 2015–16 1965 1981 239729 Indonesia 2017 1967 1982 2086 Myammar 2015–16 1966 1981 4316 Former Soviet Republics Former Soviet Republics Armenia 2015–16 1966 1981 2370 Azerbaijan 2006 1955 1970 3010 Tajikstan 2017 1967 1982 <t< th=""><th>Region</th><th>Country</th><th>Survey</th><th>Oldest cohort</th><th>Youngest cohort</th><th>Sample size</th></t<>	Region	Country	Survey	Oldest cohort	Youngest cohort	Sample size
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Ethiopia 2016 1966 1981 4196 Gabon 2012 1962 1977 2619 Kenya 2014 1964 1979 9205 Malawi 2015–16 1965 1981 6571 Mozambique 2011 1961 1976 3897 Rwanda 2014–15 1965 1980 3770 Tanzania 2015–16 1965 1981 4011 Uganda 2016 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa Z017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast		DRC	2013-14	1963	1979	5109
Gabon 2012 1962 1977 2619 Kenya 2014 1964 1979 9205 Malawi 2015–16 1965 1981 6571 Mozambique 2011 1961 1976 3897 Rwanda 2014–15 1965 1980 3770 Tanzania 2015–16 1965 1981 4011 Uganda 2016 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia		Ethiopia	2016	1966	1981	4196
Kenya 2014 1964 1979 9205 Malawi 2015–16 1965 1981 6571 Mozambique 2011 1961 1976 3897 Rwanda 2014–15 1965 1980 3770 Tanzania 2015–16 1965 1981 4011 Uganda 2016 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali		Gabon	2012	1962	1977	2619
Malawi 2015–16 1965 1981 6571 Mozambique 2011 1961 1976 3897 Rwanda 2014–15 1965 1980 3770 Tanzania 2015–16 1965 1981 4011 Uganda 2013–16 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa Zon 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Kenya	2014	1964	1979	9205
Mozambique 2011 1961 1976 3897 Rwanda 2014–15 1965 1980 3770 Tanzania 2015–16 1965 1980 3770 Uganda 2015–16 1966 1981 4011 Uganda 2013–14 1963 1979 4461 Zambia 2015 1965 1980 2770 West Africa Benin 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Malawi	2015-16	1965	1981	6571
Rwanda 2014–15 1965 1980 3770 Tanzania 2015–16 1965 1981 4011 Uganda 2016 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Mozambique	2011	1961	1976	3897
Tanzania 2015–16 1965 1981 4011 Uganda 2016 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa Benin 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Rwanda	2014–15	1965	1980	3770
Uganda 2016 1966 1981 4834 Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa Benin 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Tanzania	2015-16	1965	1981	4011
Zambia 2013–14 1963 1979 4461 Zimbabwe 2015 1965 1980 2770 West Africa Benin 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Uganda	2016	1966	1981	4834
Zimbabwe 2015 1965 1980 2770 West Africa Benin 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Zambia	2013-14	1963	1979	4461
Benin 2017–18 1967 1983 4511 Burkina Faso 2010 1960 1975 4970 Gambia 2013 1963 1978 2434 Ghana 2014 1965 1979 3147 Guinea 2018 1968 1983 3253 Ivory Coast 2011–12 1962 1977 2736 Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841	***	Zimbabwe	2015	1965	1980	2770
Benin2017-18196719834511Burkina Faso2010196019754970Gambia2013196319782434Ghana2014196519793147Guinea2018196819833253Ivory Coast2011-12196219772736Liberia2018196319782911Mali2018196819832841	West Africa	D	2017 10	10/7	1002	4511
Burkina Faso2010196019754970Gambia2013196319782434Ghana2014196519793147Guinea2018196819833253Ivory Coast2011–12196219772736Liberia2013196319782911Mali2018196819832841		Benin	2017–18	1967	1983	4511
Gambia2013196319782434Ghana2014196519793147Guinea2018196819833253Ivory Coast2011–12196219772736Liberia2013196319782911Mali2018196819832841		Burkina Faso	2010	1960	1975	4970
Ghana2014196519793147Guinea2018196819833253Ivory Coast2011–12196219772736Liberia2013196319782911Mali2018196819832841		Gambia	2013	1963	1978	2434
Guinea2018196819833253Ivory Coast2011–12196219772736Liberia2013196319782911Mali2018196819832841		Ghana	2014	1965	1979	3147
Ivory Coast2011–12196219772736Liberia2013196319782911Mali2018196819832841		Guinea	2018	1968	1983	3253
Liberia 2013 1963 1978 2911 Mali 2018 1968 1983 2841		Ivory Coast	2011-12	1962	1977	2/36
Mali 2018 1968 1983 2841		Liberia	2013	1963	1978	2911
		Malı	2018	1968	1983	2841
Niger 2012 1962 1977 3165		Niger	2012	1962	1977	3165
Nigeria 2018 1968 1983 13354		Nigeria	2018	1968	1983	13354
Senegal 2017 1967 1982 4478		Senegal	2017	1967	1982	4478

Table 1. Countries, DHS Surveys, Cohorts and Sample Sizes

(Continued)

Region	Country	Survey	Oldest cohort	Youngest cohort	Sample size
	Sierra Leone	2013	1963	1978	4952
	Togo	2013-14	1963	1979	3065

Table 1. Continued

Note: samples encompass women who were at least 35 years old at the time of each survey. *Source*: author's calculations from DHS.

were interviewed, since this makes comparable analysis for the age at first birth impossible. For each country, I use the latest available survey and focus on countries for which the latest survey was conducted after the year 2000 so that the period of data collected is similar (2001–18, with the majority covering the period 2011–18).

I use two retrospective questions about the timing of family formation: the age at first union and the age at first birth. The age at first union corresponds to the age at first marriage or cohabitation. The information about these events from the DHS is the most comprehensive data available for LMICs when it comes to geographical coverage, making it particularly useful for a cross-country analysis. It should, however, be noted that these data are not without limitations. For example, they can be influenced by recall errors. Older respondents, in particular, might have difficulty recalling events that happened early in life (Pullum, 2006). Questions about union status can be subject to biases related to interviewer effects, particularly in contexts where survey concepts might be different from local concepts (Randall et al., 2013). For some African countries, the main limitation of questions about union status is that union formation might be a long process and it might be difficult for an individual to state the exact age at which the union was initiated (Bledsoe and Pison, 1994; Meekers, 1992). Although it is important to acknowledge these limitations, there are currently no other measures that are free from such shortcomings and that could be used for conducting a large-scale study of the timing of family formation.

I focus on women born between the 1950s and 1980s (Table 1), and on women who, at the time of the survey, were at least 35 years old because, in the countries covered by the analysis, women have their first union and child relatively early. Thus, by the age of 35, the vast majority of women have experienced these events (see Appendix Table A1). Even though first family formation among women after the age of 35 is relatively rare, I also conduct sensitivity analysis that focuses on women aged at least 40. This is done to check whether this higher age cut-off leads to similar conclusions (Appendix Figures A1 and A2). The upper age limit is 49 years, as family histories of women up to that age are collected as part of the DHS. The sample sizes range from 1,886 in Guyana to 239,729 in India (Table 1).

Finally, I use data on the number of years of schooling completed, which is the primary source of information about women's educational levels. I also utilize data about place of residence, which describes whether women lived in an urban or rural area at the time of the survey, and women's height at the time of the survey. The rationale for inclusion of these two variables is described in the methods section, below.

METHODS

I use unconditional quantile regression (Firpo et al., 2009; Porter, 2015) to explore the relationship between the number of years of schooling completed and the age at first union and birth across the distributions of these two outcomes. Unconditional quantile regression allows for the assessment of how this association varies across the distribution of the age at first union and birth, adjusting for the confounding factors as measured by the control variables. In unconditional quantile regression (unlike in conditional quantile regression), quantiles are defined with reference to the unconditional distribution of an outcome, and the inclusion of the covariates does not redefine quantiles, that is, it has no effect on which observations are defined to be at the median or other quantiles of the distribution (Killewald and Bearak, 2014). It thus facilitates answering the main question in this study, namely: does the association with education vary for women in the lower and the upper parts of the distributions of the age at first union and the age at first birth, adjusting for the confounding factors as measures with the control variables? Conditional quantile regression models are less useful for answering such research questions, as they are concerned with the dispersion of the outcome variable within groups, as defined by the control variables (Firpo et al., 2009; Killewald and Bearak, 2014; Porter, 2015).

Quantile regression differs from the more commonly used regressions, such as, for example, OLS or logit and probit models, which only summarize the association of a given predictor with the mean of an outcome or a particular point in the outcome's distribution. Thus, these approaches provide only a partial view of the associations of interest because they neglect the relationships at different points in the outcome's distribution. Quantile regression sheds light on the relationship between regressors and an outcome across various points of the distribution of that outcome, providing a more thorough picture of the relationship of interest. Despite this advantage, the application of quantile regression models has been largely limited to research on labour market outcomes or student achievements (Chamberlain, 1994; Eide and Showalter, 1998; Glauber, 2018; Levin, 2001), or health research, primarily outcomes such as malnutrition or body weight (Abrevaya, 2001; Amugsi et al., 2020; Bann et al., 2020; Rahman and Hossain, 2020). To the best of my knowledge, this study is the first to use quantile regression models to explore the association between women's years of schooling and the timing of family formation.

More specifically, quantile regression allows for the examination of whether this relationship differs between the lower and upper parts of the distributions of the age at first union and birth. I explore if an additional year at school is associated with postponement of family formation uniformly in the lower quantiles (that is, where the age at family formation is lower, thus early) and the upper quantiles (that is, where the age at family formation is higher, thus late). Consequently, I focus not only on the 50th percentile (median), but also on the 25th and 75th percentiles. The unconditional quantile regression used in this study is based on the transformation of the dependent variable into the recentred influence function (RIF), defined as follows:

RIF
$$(Y; q_{\tau}) = q_{\tau} + (\tau - 1 \{Y \le q_{\tau}\}) / f_Y(q_{\tau})$$
 (1)

In Eq. (1), τ is a given quantile, q_{τ} is the value of the outcome variable, Y, at the τ^{th} quantile. $f_Y(q_{\tau})$ is the density of Y at q_{τ} , and $1\{Y \le q_{\tau}\}$ indicates whether the outcome variable falls at or below the specified quantile (dummy variable equals 1, otherwise 0). As described in the previous section, I focus only on women aged above 35 and conduct sensitivity analysis using a sample of women aged at least 40. This approach allows me to avoid the issue of right censoring, that is, the fact that some women have not yet experienced a given event. As a result, I do not need to resort to survival analysis, the application of which has not been well developed in unconditional quantile regression models.

I run two sets of models: in the first, the dependent variable is the age at first union, in the second, the age at first birth. The main explanatory variable is a continuous variable describing women's completed number of years of schooling. First, I estimate the 50th percentile quantile regression and compare the results with those from the OLS. This allows me to explore whether these two approaches, which focus on the modelling of the central tendency, provide similar results. I subsequently estimate the quantile regressions at the 25th and 75th percentiles and compare the point estimates, and their 95 per cent confidence intervals (CI), to examine whether the difference between the coefficients is statistically significant. This difference represents heterogeneity in the way additional schooling is associated with a change in the age at first union and birth at the lower and the upper parts of the distributions of these outcomes.

Apart from the completed years of schooling, I include other independent variables: women's birth cohort and place of residence. To define birth cohort, I use information about women's year of birth, grouping them into five-year cohorts. Since the surveys were conducted in different years, the cohort range differs between countries, but, in most cases, it covers individuals born between around 1960 and 1980 (see Table 1). To ensure that the sample size of each cohort is sufficient to conduct the analysis, I focus on cohorts that include at least 50 women. I add a control for women's place of residence in the models because it is known that there are

large differences in the timing of family formation and educational levels between women living in urban and rural areas (Montgomery et al., 2003). It is crucial to bear in mind that the information about the place of residence refers to that at the time of the survey and not necessarily at the time a given woman had her first union or child. Given the lack of alternatives, I use this current status information while calling attention to this limitation.

Other measures in DHS, particularly regarding socio-economic status, which are potentially associated with education and the timing of family formation (for example, the wealth index), are even more likely to suffer from this limitation. Because family formation is likely to result in moving households, the wealth index is likely to change over the life course. Thus, its value at the time of the survey may fail to adequately reflect the wealth of the household a woman lived in at the time of first union or birth. Moreover, the DHS wealth index is known to be difficult to compare within and across countries due to its conceptual complexity (Brockington et al., 2021; Chakraborty et al., 2016; Vandemoortele, 2014). The DHS provides no other information on socio-economic status, such as parental income or parental education. This makes it challenging to capture early-life conditions and women's economic background, which could be associated with both their educational level and the timing of family formation.

To address this issue, I use information about women's adult height, which is a known proxy for early childhood conditions, such as nutritional intake, parental resources devoted to children's health, and the overall economic conditions of a family and household in early life (Alderman et al., 2006; Peck and Lundberg, 1995; Silventoinen, 2003). Based on the literature highlighting this link, I run additional models that include a continuous variable describing women's height for countries for which this information is available. It should be emphasized that, despite accounting for various factors that might be associated with women's education and family formation, this study is not able to account for all potential factors. Another potential source of endogeneity that this study does not account for is reverse causality, as schooling and timing of family formation may be linked through complex pathways with both processes influencing one another (for example, schooling may affect when women form their first union, but union formation may also shape schooling trajectories). Thus, this study documents cross-sectional associations and does not isolate the causal effects of education on the timing of family formation.

RESULTS

Table 2 shows the descriptives summarizing the key variables: the 25th, 50th and 75th percentiles of the distributions of the age at first union and birth, and the mean years of education. The countries exhibit a vast degree of variation in the profiles of women regarding the timing of key events and

			Age at first union			Age at first birth		
Region	Country	25th	50th	75th	25th	50th	75th	Mean Years
Americas								
-	Bolivia	17.8	20.3	24.1	18.4	20.8	24.2	6.9
	Colombia	17.9	20.8	24.9	18.7	21.3	25.3	9.2
	Dominican Rep.	16.3	19.1	22.8	18.1	20.7	24.3	9.3
	Guatemala	16.7	18.8	21.8	17.9	19.9	22.8	4.5
	Guvana	17.6	20.2	24.3	18.3	20.3	23.4	8.3
	Honduras	16.6	18.9	22.2	17.9	20.0	23.0	6.2
	Nicaragua	15.7	17.9	20.9	17.3	19.4	22.3	5.5
	Peru	18.0	20.8	24.9	18.8	21.3	25.3	8.9
Asia								
	Cambodia	17.4	19.8	22.8	19.2	21.5	24.3	3.9
	India	15.9	18.3	21.1	18.2	20.5	23.3	4.6
	Indonesia	17.4	20.3	23.9	19.1	21.7	25.2	8.6
	Myanmar	18.1	20.7	24.8	19.9	22.7	26.6	5.0
	Nepal	15.8	17.5	19.9	18.3	20.1	22.3	2.5
	Philippines	19.1	22.2	26.0	20.3	23.1	26.8	10.7
	Timor-Leste	18.2	21.2	24.7	19.8	22.6	26.0	5.4
Former So	oviet Republics							
	Armenia	18.4	20.1	23.0	19.6	21.4	24.3	11.9
	Azerbaijan	19.4	21.7	24.5	20.7	23.0	25.9	10.9
	Kyrgyzstan	18.7	20.2	22.5	20.0	21.6	23.9	11.9
	Rep. of Moldova	19.1	20.5	22.3	20.3	21.8	23.7	11.6
	Tajikistan	18.3	19.7	21.8	19.8	21.3	23.6	10.4
	Ukraine	18.9	20.4	22.5	20.1	21.9	24.3	14.0
East & Ce	entral Africa							
	Angola	17.0	19.9	24.6	17.4	19.7	22.8	4.0
	Burundi	18.0	20.3	23.3	19.3	21.6	24.4	2.5
	Cameroon	15.5	17.8	21.3	16.9	19.2	22.1	5.0
	Chad	14.8	16.2	18.9	16.1	18.6	21.8	1.1
	Congo	16.7	19.3	23.0	17.2	19.3	22.2	7.4
	DRC	15.9	18.3	21.7	17.3	19.7	22.8	5.3
	Ethiopia	14.6	16.3	19.6	16.5	18.7	21.5	1.7
	Gabon	17.3	20.8	27.1	17.0	18.8	21.4	7.6
	Kenya	17.3	19.8	22.8	17.9	20.1	22.8	7.8
	Malawi	16.2	18.2	20.6	17.3	19.1	21.3	4.3
	Mozambique	16.2	18.9	23.0	17.0	19.7	22.8	2.4
	Rwanda	18.7	20.9	23.4	19.8	22.0	24.7	4.5
	Tanzania	16.8	18.8	21.7	17.8	19.5	21.8	5.6
	Uganda	16.0	18.2	21.3	16.8	18.9	21.3	4.8
	Zambia	16.1	18.0	20.6	17.2	18.8	21.0	5.9
	Zimbabwe	17.6	19.8	22.4	18.3	20.2	22.6	9.1
West Afri	ca							
	Benin	16.5	19.4	23.0	17.8	20.3	23.5	1.4
	Burkina Faso	16.3	17.8	19.8	17.7	19.5	21.8	0.9
	Gambia	15.1	17.6	20.8	16.5	19.2	22.3	2.3
	Ghana	17.0	19.8	23.8	18.0	20.6	24.0	5.8
	Guinea	15.8	18.6	22.9	17.1	20.1	24.2	1.0
	Ivory Coast	16.1	18.8	22.8	16.7	19.0	22.1	2.6
	Liberia	15.8	18.2	21.8	16.3	18.3	21.3	3.1

Table 2. Descriptive Statistics for the Outcomes (Percentiles of the Age atFirst Union and Birth Distributions) and the Key Predictor (Mean Years ofEducation) in the Samples

(Continued)

		Age	Age at first union		Age at first birth			Education	
Region	Country	25th	50th	75th	25th	50th	75th	Mean Years	
	Mali	15.6	17.8	21.0	17.3	19.4	22.8	1.0	
	Niger	14.6	15.6	17.6	16.7	19.0	22.2	0.8	
	Nigeria	15.4	18.8	23.9	17.1	20.0	24.3	5.8	
	Senegal	17.0	19.9	23.8	18.5	21.3	24.8	2.4	
	Sierra Leone	15.6	17.8	20.9	16.8	19.6	22.8	1.7	
	Togo	17.0	19.8	23.3	18.0	20.7	23.9	3.0	

Table 2. Continued

Source: author's calculations from DHS.

education. The timing of family formation is the earliest in West Africa and the latest in former Soviet Republics. It is also in these two regions that women have the lowest and the highest levels of schooling, respectively.

The Relationship between Years of Schooling and the Timing of Family Formation: Differences by Quantiles

The results of the OLS and the median quantile regressions describing the relationship between schooling and the age at first union and birth are presented in Table 3. The point estimates of the OLS and the median quantile regression are similar in magnitude, which means that the two approaches focusing on the central tendency provide similar evidence. Although there is a noticeable variation in the magnitude of the associations between countries, with coefficients ranging from 0.1 to 0.6, they are positive and statistically significant at the 5 per cent level in most nations, with the exception of Congo, Ivory Coast and Liberia in the median quantile regression, a one-year increase in schooling completed is associated with 0.3 years (around four months) and 0.45 years (around six months) increase in the (average/median) age at first union in Bolivia and Peru, respectively. This is in line with literature documenting a positive relationship between years of education and the timing of family formation.

A noteworthy pattern is that, while in the Americas, Asia and the former Soviet Republics the magnitude of the association is roughly similar for the two events, in African countries, particularly East & Central Africa, the link between education and the age at first union is stronger than it is with the age at first birth. This finding is consistent with a study on Madagascar, according to which an additional year of education delayed women's marriage more than it delayed their first birth (Glick et al., 2015). One explanation could be high levels of premarital childbearing in Africa. Bongaarts et al. (2017) have shown that, while the mean age at first birth is higher than the mean age at first union across most countries covered by their analysis,

		Age	at first union	Age at first birth		
Region	Country	OLS	QR, median	OLS	QR, median	
Americas						
	Bolivia	0.29	0.28	0.29	0.25	
	Colombia	0.39	0.42	0.42	0.40	
	Dominican Rep.	0.42	0.40	0.40	0.36	
	Guatemala	0.35	0.33	0.33	0.29	
	Guyana	0.28	0.31	0.14	0.14	
	Honduras	0.38	0.35	0.36	0.31	
	Nicaragua	0.37	0.34	0.36	0.34	
	Peru	0.45	0.46	0.45	0.43	
Asia						
	Cambodia	0.20	0.22	0.20	0.21	
	India	0.31	0.30	0.26	0.24	
	Indonesia	0.54	0.57	0.49	0.52	
	Myanmar	0.49	0.48	0.45	0.46	
	Nepal	0.29	0.26	0.21	0.19	
	Philippines	0.53	0.60	0.48	0.51	
	Timor-Leste	0.17	0.21	0.17	0.20	
Former Sov	viet Republics					
	Armenia	0.54	0.55	0.52	0.56	
	Azerbaijan	0.21	0.30	0.23	0.31	
	Kyrgyzstan	0.35	0.34	0.37	0.38	
	Rep. of Moldova	0.24	0.20	0.27	0.24	
	Tajikistan	0.29	0.23	0.27	0.24	
	Ukraine	0.16	0.16	0.19	0.17	
East & Cen	tral Africa					
	Angola	0.11	0.15	0.06	0.05	
	Burundi	0.34	0.28	0.3	0.27	
	Cameroon	0.50	0.48	0.28	0.23	
	Chad	0.23	0.19	0.11	0.10	
	Congo	0.14	0.11	0.03 (ns)	0.02 (ns)	
	DRČ	0.20	0.18	0.12	0.12	
	Ethiopia	0.28	0.19	0.21	0.14	
	Gabon	0.33	0.33	0.11	0.06	
	Kenya	0.31	0.31	0.21	0.20	
	Malawi	0.23	0.20	0.16	0.15	
	Mozambique	0.21	0.22	0.14	0.10	
	Rwanda	0.28	0.26	0.23	0.24	
	Tanzania	0.31	0.23	0.24	0.18	
	Uganda	0.28	0.25	0.24	0.19	
	Zambia	0.31	0.26	0.22	0.17	
	Zimbabwe	0.41	0.42	0.43	0.40	
West Africa	1					
	Benin	0.38	0.36	0.37	0.32	
	Burkina Faso	0.32	0.18	0.28	0.17	
	Gambia	0.29	0.28	0.22	0.24	
	Ghana	0.26	0.20	0.23	0.18	
	Guinea	0.19	0.20	0.14	0.11	
	Ivory Coast	0.34	0.24	0.17	0.04 (ns)	
	Liberia	0.22	0.18	0.02 (ns)	0.01 (ns)	
				× ,		

Table 3. Coefficients from the Ordinary Least Square Regression (OLS) and the Quantile Regression (QR) for the 50th Percentile (Median), Describing Association between Schooling and the Age at First Union and Age at First Birth

(Continued)

Region	Country	Age	at first union	Age at first birth		
		OLS	QR, median	OLS	QR, median	
	Mali	0.36	0.22	0.11	0.11	
	Niger	0.43	0.17	0.31	0.34	
	Nigeria	0.48	0.56	0.39	0.39	
	Senegal	0.48	0.38	0.35	0.28	
	Sierra Leone	0.22	0.17	0.09	0.10	
	Togo	0.35	0.32	0.32	0.26	

Table 3. Continued

Notes: Models include a control for women's birth cohort and place of residence.

Ns means that a given coefficient is not statistically significant at 5% level. All remaining coefficients are statistically significant from zero at 5% level.

Source: author's calculations from DHS.

the nations that were an exception to this pattern were mainly situated in sub-Saharan Africa. Moreover, evidence suggests that more educated women in sub-Saharan Africa might have a greater likelihood of giving birth premaritally (Meekers, 1994). Relatedly, in their study of premarital childbearing in sub-Saharan Africa, Smith-Greenaway and Clark (2018) suggest that, since getting married is generally incompatible with continuing school, women who are unmarried at the time of their first birth may decide to remain so in order to complete their education before marrying their future partner. This is in line with a study on Malawi, which found that premarital childbearing is more common among more educated women and often happens in the context of educational advantage (Smith-Greenaway, 2016). The weaker relationship between education and the timing of first birth, compared to first union, could reflect a pattern of more educated women transitioning to motherhood but postponing entering unions.

Figures 1 and 2 present results of the quantile regressions for the 25th and 75th percentiles for the age at first union and first birth, respectively. For comparability, in both figures, the countries are sorted according to the magnitude of the coefficients of the median quantile regression for the first union (from the smallest to the largest). For each country, two-point estimates are shown, one for the 75th percentile (black) and one for the 25th percentile (grey), with 95 per cent confidence intervals. In countries where the confidence intervals for the two estimates do not overlap, the Figures depict a difference between these two-point estimates (marked as Diff.).

The results corroborate the findings from the median quantile regression in that completing more years of education is associated with an increase in the age at first union and first birth (see positive coefficients of the quantile regressions at the 25th and 75th percentiles). However, in many countries, the magnitude of the relationship varies considerably across the distributions of these outcomes. Namely, the association is stronger in the upper quartile (that is, the upper part of the age at first union and birth distributions) than in the lower quartile. This means that education has a

Figure 1. Coefficients Describing the Association between Schooling and the Age at First Union from the QR (25th and 75th Percentiles). Difference in the Coefficients from the 75th and 25th Percentile QR (Diff.), if Statistically Significant



Notes: Countries sorted according to the magnitude of the coefficient from the median QR. Models include a control for women's birth cohort and place of residence.

Figure 2. Coefficients Describing the Association between Schooling and the Age at First Birth from the QR (25th and 75th Percentiles). Difference in the Coefficients from the 75th and 25th Percentile QR (Diff.), if Statistically Significant



Notes: Countries sorted according to the magnitude of the coefficient from the median QR for the age at first union. Models include a control for women's birth cohort and place of residence.

heterogeneous association with the timing of family formation across the distribution of these events and is significantly weaker in their bottom part.

In terms of the age at first union (Figure 1), these relationships are particularly visible in West Africa and the Americas, where the coefficients from the quantile regression pertaining to the 25th percentile are markedly smaller than those for the 75th percentile. The 95 per cent confidence intervals of the estimates do not overlap in any of the countries in these two subregions, except for Guyana. For example, in Sierra Leone, an additional year of schooling is associated with an increase in the age at first union of 0.10 years at the 0.25 quantile of the distribution of the age at first union; at the 0.75 quantile, the corresponding increase is 0.31 years (difference of 0.21). For some countries, such as Senegal, the disparity is even bigger: the point estimate for the 75th percentile (0.75) more than triples when compared to the 25th percentile (just below 0.25).

The fact that an additional year of schooling raises the 75th percentile by more than it raises the 25th percentile highlights heterogeneity in how education is related to the timing of union formation. While an additional year at school increases the age at first union at all levels of the age at first union distribution, the association is weaker in the lower quantile. Substantively, these results mean that the postponement of family formation associated with more years at school is less pronounced among women who form families relatively early than those who form families relatively late. Thus, schooling has a relatively weaker effect on the reduction of early unions than it has on shifting upwards the age at these events in the upper part of the distributions. Importantly, the existence of marked differences in the coefficients of the quantile regressions and a weaker association at the lower ends of the age at first union distribution also mean that there is a tendency for the dispersion of this distribution to grow with increasing years of schooling. In other words, education is associated with increasing inequality in the timing of union formation because it widens the distribution of the age at this transition.

A similar pattern of heterogeneity for the age at first union can be observed in the majority of nations of East & Central Africa and Asia, although differences between the estimates (Diff.) are smaller than in West Africa and the Americas. In the former Soviet Republics, this association differs between the upper and lower parts of the distribution in half of them. On the one hand, the coefficients from the quantile regressions for the 25th percentiles are smaller than for the 75th percentiles in Tajikistan, Kyrgyzstan and Armenia. However, they are analogous in magnitude in Ukraine, Moldova and Azerbaijan, suggesting that schooling has a similar association with the age at first union in both the lower and upper parts of the distributions.

Similar conclusions are reached when the link between years of schooling and the age at first birth is considered (Figure 2), although there are some notable regional differences. In the Americas, Asia and the former Soviet Republics, the results for the age at first birth largely mirror those for the age at first union. In West Africa and East & Central Africa, the pattern by which an additional year at school raises the 75th percentile of the first birth distribution more than it raises the 25th percentile is visible in 8 out of 13 and 6 out of 16 countries, respectively. Consequently, the heterogeneous association of education across the distribution of the age at first birth is visible in fewer African countries than is the case for the age at first union, particularly in East & Central Africa. This pattern could be related to the point discussed in the previous section that the association with education. as described by the median regression, is weaker in the case of the age at first birth than the age at first union. These results could suggest that education has a more stratifying role when it comes to the timing of union formation than the entrance into motherhood. Nonetheless, it should be noted that, although apparent in fewer countries, the unequalizing role of schooling across the first birth distribution shown in Figure 2 is still present in most countries in West Africa, and around half of all the African countries.

The presented results highlight that, by characterizing the relationship between the years of schooling completed and the age at first union and birth at different points of the distribution of these events, quantile regression provides a more comprehensive picture than the one offered by examining only their centre. Overall, in many of the countries covered by this study, and particularly in the Americas and West Africa, the association with education weakens as one moves down the distributions of the age at first union and birth.

Robustness Checks and Alternative Model Specifications

As explained above, the main analysis focuses on women who were at least 35 years old at the time of the survey. This is motivated by the need to find a balance between an analysis that focuses on women who are unlikely to still be at risk of experiencing these 'first time' events, and ensuring that the sample sizes are sufficiently large. As a robustness check, I replicate the analysis for a smaller subsample of women who were 40 years old at the time of the survey (see Appendix Figures A1 and A2). As expected, the confidence intervals of the estimates are wider than in the analysis covering the sample of women aged 35+. Nonetheless, the pattern described in the previous section is clearly visible in the majority of countries, with the exception of a few (five and six countries in the case of age at first union and birth, respectively), mainly in West Africa and East & Central Africa (for example, see Guinea in West Africa for the analysis of the age at first union where the confidence intervals of the estimates for the 25th and 75th percentiles now overlap). These results suggest that the findings are largely robust to a more conservative model specification.

Second, I conduct an analysis that includes a control variable capturing women's adult height. Note that all women included in this study were above the age of 35 at the time of survey, an age beyond which height no longer increases. The aim is to proxy and capture women's early-life socio-economic status, which is likely to be associated both with their education and the timing of family formation. This analysis cannot be conducted for all countries because not all surveys collect women's anthropometric information: Colombia, Angola, Indonesia, Philippines, Senegal and Ukraine do not record such data. Moreover, for some countries, these data are collected for subsamples of women, ranging from 35 per cent (for example, Malawi) through to 50 per cent (for example, Benin) and 65 per cent (for example, Cambodia).

The results of the models that account for women's height largely corroborate the findings presented in Figures 1 and 2 (see Figures A3 and A4 in the Appendix), although, as expected, due to the much smaller sample sizes for many countries, the confidence interval estimates are also much wider. In the case of a few countries, the evidence of the heterogeneous association of education across the distributions of the age at first union and birth, as shown in Figures 1 and 2, can no longer be observed when height is included (four and five countries for the age at first union and birth, respectively); for example, Liberia or Congo in the analysis of the age at first union. for which the confidence intervals for the 25th and 75th percentiles now overlap. Overall, however, in most countries where the initial models pointed to heterogeneity in the association of education at the upper and the lower parts of the age at first union and birth distributions, this evidence is still supported. These results suggest that differences in women's early-life socio-economic conditions are unlikely to fully explain the heterogeneous relationship between education and the timing of family formation.

DISCUSSION

Many studies have examined the association between schooling and the timing of union formation and transition to motherhood in LMICs, but there has been a lack of research on whether this association varies within countries. Using quantile regression analysis, this article fills this void, providing evidence of heterogeneity in the relationship between education and the timing of family formation within the populations of women. These results have implications for future research on the determinants of family formation and for policy making.

The results draw attention to the importance of considering the whole distributions of the age at first union and birth when studying links between these events and education. Moreover, due to the evidence that family formation early in life might have negative effects on women's well-being and that of their children (Sunder, 2019; Urdinola and Ospino, 2015;

Yount et al., 2018), it is important to examine not only whether education shifts early family formation, but also the relative strength with which it may do so. The results presented here provide evidence that receiving more schooling does not translate into uniform changes in the timing of family formation within populations. This is particularly evident in the Americas and West Africa, where education has a weaker effect on the reduction of early first unions and births than it has on further shifts in late union formation and childbearing. Moreover, these marked differences in the strength of the association with education between the lower and the upper parts of the distributions of the age at first union and birth mean that, with increased years of schooling, the dispersion of the distributions of (and thus the inequality in) the age at first union and first birth tends to increase. This is an indication of the unequalizing role of education in shaping the age at family formation.

These results imply that — especially in the Americas and West Africa, where many women form families early in life and thus might be at risk of adverse outcomes — increasing the number of years that girls/women stay at school might be insufficient to contribute to shifting this family-building behaviour. These findings are important in understanding the persistence of high levels of early marriage and pregnancy in these subregions, despite educational expansion, and the large disparities in the age at first union and birth. The presented results are consistent with evidence that the processes related to family formation in the Americas and West Africa are very unequal, and, while some segments of women's populations substantially postpone unions and motherhood, others continue to experience these events very early (Batyra and Kohler, 2022; Esteve and Florez-Paredes. 2018: Koski et al., 2017: Lima et al., 2018). By showing that schooling is not uniformly associated with the timing of family formation, this study expands the body of knowledge on the links between heterogeneity in the timing of family formation and socio-economic inequalities in LMICs.

These results are important from a policy perspective, as they underline that, despite the generally positive association between additional schooling and the age at family formation, expanding access to education may provide insufficient impetus for the reduction of early unions and childbearing in LMICs. The results support the mounting evidence that receiving more schooling is not necessarily equivalent to learning and does not necessarily enhance labour market prospects. It has been increasingly acknowledged that educational attainment is not a reliable marker of knowledge achievement and skills, as many students in LMICs have been receiving poor-quality education (Angrist et al., 2021; Psaki et al., 2018; World Bank, 2018, 2019). These disparities in education quality could be one of the drivers of the heterogeneous association between schooling and the timing of family formation. If spending more time in school brings little in terms of learning and work opportunities (one of the channels through which education influences family-formation behaviour) among some women, then one can expect that the way education is related to family formation will vary across the population.

The results can also be interpreted through the lens of research highlighting the limited role of schools in inducing a change in the perceptions about women's status and their role in society (Aikman et al., 2005; Kabeer, 2005; Stromquist, 2001; Urbina, 2022), which are key in transforming family-formation behaviours. As highlighted by Kabeer (2005), in settings with high levels of gender inequality, education might have little influence, given prevailing norms that the role of women in society should be limited to domestic spheres. In fact, schools might actually contribute to reproducing existing, not only gender, but also other social status inequalities, as highlighted by studies on different settings, such as Latin America (Stromquist, 2001). The rigidity of gender norms and educational structures that do little to tackle existing social divisions might explain why school attendance among some girls/women does not necessarily translate into as strong as expected delays in family formation. These aspects are important when thinking about the transformative power of education in the context of entrenched social norms. Overall, it is possible that the barriers to accessing quality education, limited labour market opportunities, and persistent norms related to women's status in society, all weaken the association between additional years of schooling and the postponement of family formation among some segments of the population.

This study has limitations. First, it did not aim to capture the causal effect of education on the timing of family formation. Although robustness checks suggest that accounting for women's height, which is a proxy for early-life socio-economic conditions, does not change the interpretations, other possible factors could not be accounted for due to the limited amount of retrospective information. It should also be highlighted that the use of height has limitations. For example, height may be affected by temporary shocks very early in life such as environmental or economic crises (Adsera et al., 2021; Rosales-Rueda, 2018), in which case it may be a less accurate reflection of women's overall socio-economic status and the economic conditions of their family and household. Unfortunately, DHS does not offer better proxies of socio-economic status that could be used in a comparative study like this one. The extent to which the associations used could represent causal effect could be explored in future research, for example, by conducting in-depth, country-specific studies using techniques that take account of potential endogeneity, including reverse causality.

Moreover, this study was concerned with cross-sectional associations and did not consider dynamics of change in the relationships of interest. Because of data availability, the coverage of surveys by country differed, with most of them being conducted in the late 2010s, but some a decade earlier (Nicaragua in 2001). This means that this study was not able to account for any transitions (or differences therein) happening across countries within that period, such as shifts in fertility rates. Future studies could use in-depth analyses to cast light on whether the highlighted associations change over time, or conduct macro-level analysis of how they may be shaped by different demographic factors. This would be important once newer data become available and analysis encompassing younger cohorts becomes possible. Moreover, the data currently available for a woman's first union do not distinguish between marriage and cohabitation. Finally, although this study focuses on women below the age of 50 at the time of the survey, and thus does not consider the oldest individuals, it should be kept in mind when interpreting the results that there might also be differentials in mortality by educational level.

This article opens important avenues for future investigation. Research exploring the factors behind the heterogeneous association between schooling and the timing of family formation within countries — such as disparities in quality of schooling and labour market opportunities, and variation in social norms — is needed. The results of this study suggest that further comparative research on the determinants of disparities, focusing on various aspects of socio-economic inequality, would be vital to better understand the growing diversity of family formation behaviours.

APPENDIX

Table A1. Percentage of Women Who Have Never Been in Union or Have Never Given Birth (Women Who Were at Least 35 or 40 Years Old at the Time of the Survey)

		% not ex	periencing an	event (by ag	e at survey)
Region	Country	Union (35)	Union (40)	Birth (35)	Birth (40)
Americas					
	Bolivia	5.3	4.4	4.4	3.9
	Colombia	8.0	7.3	5.7	5.1
	Dominican Republic	2.6	2.2	4.6	4.1
	Guatemala	7.3	6.2	6.3	5.3
	Guyana	8.7	8.3	6.7	6.2
	Honduras	5.2	4.8	4.4	4.3
	Nicaragua	2.7	2.8	3.2	3.2
	Peru	7.1	6.2	5.7	5.1
Asia					
	Cambodia	5.0	4.7	7.6	7.1
	India	1.7	1.5	4.3	3.9
	Indonesia	2.8	2.7	6.3	6.2
	Myanmar	13.9	12.9	17.7	16.3
	Nepal	0.9	1.1	2.5	2.5
	Philippines	7.4	6.9	9.3	8.4
	Timor-Leste	6.2	5.9	7.6	7.4
Former USSR					
	Armenia	5.4	4.6	7.8	7.4
	Azerbaijan	8.3	6.6	12.2	10.3
					(Continued)

(Continued)

		% not experiencing an event (by age at survey)				
Region	Country	Union (35)	Union (40)	Birth (35)	Birth (40)	
	Kyrgyzstan	1.1	1.3	3.0	3.1	
	Republic of Moldova	1.8	1.3	4.7	4.3	
	Tajikistan	3.5	2.7	5.6	4.3	
	Ukraine	2.5	2.1	5.7	4.8	
East & Central Africa						
	Angola	9.6	9.4	3.1	3.3	
	Burundi	4.0	3.4	4.1	3.2	
	Cameroon	3.3	2.8	3.6	3.0	
	Chad	0.4	0.4	1.1	1.1	
	Congo	2.6	2.5	2.3	2.2	
	DRC	2.3	2.1	3.0	2.7	
	Ethiopia	3.6	2.8	5.0	4.3	
	Gabon	7.4	6.5	4.5	4.2	
	Kenya	4.7	4.4	2.1	1.9	
	Malawi	1.2	1.1	1.4	1.4	
	Mozambique	3.2	3.0	4.2	4.3	
	Rwanda	5.4	4.1	3.9	3.7	
	Tanzania	2.8	2.1	2.7	2.4	
	Uganda	2.3	2.0	1.7	1.9	
	Zambia	2.2	1.5	2.5	2.3	
	Zimbabwe	3.2	3.5	3.6	4.2	
West Africa						
	Benin	1.3	1.4	1.7	1.5	
	Burkina Faso	0.3	0.2	0.9	0.7	
	Gambia	1.2	0.8	2.9	3.5	
	Ghana	2.3	1.3	3.0	2.1	
	Guinea	1.8	1.3	4.5	4.5	
	Ivory Coast	4.7	3.7	3.4	2.7	
	Liberia	1.6	1.0	1.3	1.2	
	Mali	1.3	1.3	3.2	2.9	
	Niger	0.8	0.6	2.2	1.6	
	Nigeria	2.9	2.2	3.9	3.4	
	Senegal	2.6	2.1	4.6	4.1	
	Sierra Leone	2.3	1.7	1.9	2.1	
	Togo	1.9	1.2	2.4	2.1	

Table A1. (Continued)

Source: author's calculations from DHS.

Figure A1. Coefficients Describing the Association between Schooling and the Age at First Union from the QR (25th and 75th percentiles). Difference in the Coefficients from the 75th and 25th Percentile QR (Diff.), if Statistically Significant. Subsample of Women Aged at Least 40 Years at Survey



Note: The 95% CI for the 75th pct. in Angola, missing from the output due to larger bottom range than in other countries, is (-0.18, 0.13), respectively. For comparability, countries are sorted according to the magnitude of the coefficient from the median QR for models with control for women's birth cohort and place of residence.

Figure A2. Coefficients Describing the Association between Schooling and the Age at First Birth from the QR (25th and 75th percentiles). Difference in the Coefficients from the 75th and 25th Percentile QR (Diff.), if Statistically Significant. Subsample of Women Aged at Least 40 Years at Survey



Note: The 95% CI for the 75th pct. in Liberia, Guinea, Mali and Angola, missing from the output due to larger bottom range than in other countries, is (-0.11,0.06), (-0.23,0.15), (-0.13,33) and (-0.20,0.03), respectively. For comparability, countries are sorted according to the magnitude of the coefficient from the median QR for the age at first union for models with control for women's birth cohort and place of residence.

Figure A3. Coefficients Describing the Association between Schooling and the Age at First Union from the QR (25th and 75th percentiles). Difference in the Coefficients from the 75th and 25th Percentile QR (Diff.), if Statistically Significant. Countries with Information about Women's Height



Note: For comparability, countries are sorted according to the magnitude of the coefficient from the median QR for models without control for women's height. Models control for women's birth cohort, place of residence and height.

Figure A4. Coefficients Describing the Association between Schooling and the Age at First Birth from the QR (25th and 75th percentiles). Difference in the Coefficients from the 75th and 25th Percentile QR (Diff.), if Statistically Significant. Countries with Information about Women's Height



Note: The 95% CI for the 75th pct. in Liberia, missing from the output due to larger bottom range than in other countries, is (-0.14; 0.06). For comparability, countries are sorted according to the magnitude of the coefficient from the median QR for the age at first union for models without control for women's height. Models control for women's birth cohort, place of residence and height.

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