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Education and Polygamy: Evidence from Cameroon*

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Abstract. — Has secular education contributed to the decline of polygamy in Africa? To answer this question, we study a wave of public school construction in late-colonial Cameroon. Our difference-in-differences and event-study specifications show that school openings have simultaneously increased education and the chances to be in a polygamous union for men and, more surprisingly, for women. We estimate a structural model of marriage to explain why education made women more likely to be in a polygamous union. The main estimated channel is marriage to educated men who are more often polygamists than uneducated men, not direct preferences for polygamy.

JEL classification: J12, I20, O12.

Keywords: polygamy, education, marriage, matching models.

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1. Introduction

Polygamy is a widespread marital custom in many parts of the world, especially in the African “polygyny belt” between Senegal and Tanzania, where between a third and half of married women are in a polygamous union (Jacoby, 1995; Fenske, 2015).¹ Over the past century however, polygamy has been decreasing substantially, a marriage market transformation “as dramatic as the rise of divorce in the US” (Fenske, 2015). This paper contributes to testing whether the concurrent expansion of education – Sub-Saharan African gross primary enrollment rates increased from 29.2% in 1950 to 89.3% in 2010 (Barro and Lee, 2013) – has contributed to the decline of polygamy. To do so, we estimate whether and how education causally affects polygamy, studying the marriage practices of cohorts who were young during a late-colonial period of massive primary school expansion in Cameroon, a Central-African country. Polygamy is still important in Cameroon, but it has been declining in recent decades.²

Education can affect polygamy directly, through the *transformation of cultural norms*. In Africa during the colonial period, Christian missions and colonial governments provided formal education. In this paper, we focus on public, nonreligious education. Anthropologists and historians have stressed the role of Christian missions and schools in promoting monogamous marriage and contributing to the decline of polygamy on the continent (Egboh, 1972; Ekechi, 1976; Oduyoye, 1993; Falen, 2008; Walker-Said, 2018). This has been confirmed by more recent statistical analyses (Kudo, 2017; Becker, 2022). Colonial governments also tended to view monogamy as a superior marital custom, but, in part to ensure the cooperation of local elites, were less hostile to polygamy than Christian missions (Walker-Said, 2018, for Cameroon). Secular education provided in colonial government schools might nonetheless have contributed to a transformation of marital customs in line with European norms.

Education can also affect polygamy through the channel of *female empowerment*. Edu-

¹Throughout this paper, we use the word “polygamy” as a synonym for “polygyny” (the practice of a man having several wives). Technically, “polygamy” also includes the rare practice of a woman having several husbands (“polyandry”).

²In the 2005 census, 25% of married women were in a polygamous union. This figure was 43% in 1976.

cation provides opportunities for women on the labor market, making them less dependent on marriage for their material well-being and improving their bargaining power. Higher bargaining power with respect to their parents could increase the weight of their own preferences in the choice of a husband. In the model from Tertilt (2006), transferring the right of choosing a husband from fathers to daughters reduces polygamy. Higher bargaining power with respect to their husband could prevent him from taking a second wife.

Finally, education could affect polygamy through the channel of *marriage-market returns to education* (Chiappori et al., 2009). If education is valued by potential partners, more-educated individuals will attract suitors with more desirable traits (education, but not only). Setting aside any direct cultural effect of education on the acceptance of polygamy, education could make men more desirable, allowing them to marry more than one wife. For women, this channel is more complicated. If educated women are more desirable on the marriage market, it will be easier for them to marry monogamously if they value monogamy. However researchers disagree on the extent to which African women valued monogamy (Ekechi, 1976; Falen, 2008). Furthermore women (and their families, who are also involved in marriage decisions) might value both monogamy *and* some husband traits correlated with polygamy, like wealth or social prestige.

To investigate the causal effect of education on the likelihood that men and women are in a polygamous union, we exploit exogenous variation in schooling caused by a late-colonial-era wave of public school construction: selection into polygamy on observed and unobserved characteristics is an important problem, as shown by Fenske (2015) and Arthi and Fenske (2018). We observe the individuals who benefited from this education boom in Cameroonian census microdata from 1976, matched with administrative data on the universe of schools. We use variation in the timing of public, nonreligious school openings in different villages to identify the causal effect of education on marriage-market outcomes. Our event study shows that the opening of a public school in a village does not predict the education or polygamy of individuals too old to go to that school, lending credence to the common trend assumption. In a quasi difference-in-differences setting, we

instrument an individual's education by the stock of public, nonreligious schools in their village when they were of school age. In the sample of married people,³ education increases the likelihood of being in a polygamous union for men (by 10.1 percentage points) and for women (by 6.6 percentage points). The effect on women is entirely explained by women entering a polygamous union as a first wife – education does not change the probability of entering a polygamous union as a second wife.

The effect on men is easily explained: the positive causal effect of education on polygamy means that the channel of marriage-market returns trumps any direct cultural effect. For women, a positive causal effect of education on polygamy is harder to explain. Female education seems desirable on the marriage market: educated women marry more-educated husbands with better labor-market outcomes. Did Cameroonian women in later colonial Cameroon really prefer polygamy, at least as first wives? That African women prefer monogamy has appeared obvious to many, including anthropologists such as Goody (1989). However, others including Ekechi (1976) and Falen (2008) have stressed this view as Euro-centric and have underlined that marrying a polygamist can be desirable and socially prestigious. In fact, what these authors imply is not a preference for polygamy itself, but for some other husband traits correlated with polygamy. Falen (2008) writes that, in Benin, “many women [...] strongly prefer monogamy, but some accept becoming a cowife for financial security.” Of West African women, Ekechi (1976) writes that they “derive social prestige by being married, especially to a wealthy polygamist.” In the language of economists, we would say that the choice of a husband involves trade-offs between different husband characteristics that are correlated. The positive reduced-form effect of women education on polygamy could be explained by a preference of more-educated women (and their family, as the choice is rarely made by the woman alone) for more-educated, richer men who are also more likely to be polygamists.

To investigate this, we estimate a structural model of marriage with utility transfers (Choo and Siow, 2006) that we adapt to polygamy, allowing men to marry several women. Given data on marriages (who is married with whom), this model allows us to estimate

³We also find that education does not significantly affect the likelihood of marriage for men and women alike.

a matrix of “affinities” between different characteristics of husbands and wives (Dupuy and Galichon, 2014). These affinities tell us the extent to which different characteristics of men and women attract or repulse one another on the marriage market. We are particularly interested in the affinity between the number of wives the husband has and the education of each wife. More precisely, we want to know if this parameter depends on whether we allow a direct affinity between education of the husband and education of the wife. Our reduced-form results might suggest that educated women and polygamous men attract one another. We want to know whether it is still the case when we consider the fact that educated men and women attract one another (assortative matching on education). The structural estimation of an affinity matrix also allows us to disentangle the matching on education from the matching on age (age and educational attainment are strongly negatively correlated in a context of schooling expansion).

We show that the affinity parameters we are interested in can be recovered by estimating a logit on pairs of couples within the same village. We first estimate the affinity parameter between a husband’s number of wives and each wife’s education. It is positive, as expected from the reduced-form result: when we do not consider assortative matching on education, polygamy and education appear to attract. Then, we allow a direct affinity between education of the husband and education of the wife: the affinity between the husband’s number of wives and each wife’s education is divided by three, and it loses statistical significance. Looking separately at first wives and wives of higher rank, we find that the positive affinity between education and polygamy is entirely explained by first wives. Considering wives of higher-order rank, the affinity between education and polygamy is positive but small and not significant, even without allowing for assortative matching on education. When allowing for assortative matching on education, the affinity between polygamy as a wife of higher-order rank and woman’s education turns negative (but it is not statistically different from zero).

In sum, education increased the likelihood that women would become the first wives of polygamists and did not decrease the likelihood that they would become second wives of polygamists, but this is not because women prefer polygamy per se. Rather, it’s because

polygamy is correlated with other valuable attributes of a husband, such as education.

Contributions Our paper contributes to the literature on the economics of polygamy, studying its economic consequences, its origins, and the reasons for its decline. Though we do not contribute directly to the literature on the consequences of polygamy, it is useful to summarize it to frame the subject's importance. Tertilt (2005) builds a model where banning polygamy decreases fertility and increases savings and output per capita. Arthi and Fenske (2018) find a positive relationship between polygamy and child mortality but conclude it is likely largely explained by selection into polygamy. Barr et al. (2019) and Rossi (2019) study intrahousehold cooperation in polygamous households. Bertocchi and Dimico (2019) find that HIV infection rates are higher where polygamy is more widespread and use the demographic shock of the slave trade as an instrument to show a causal relationship.

Our paper contributes more specifically to the literature trying to understand the existence of polygamy, and its decline. A first group of works attempts to explain the existence of polygamous and monogamous societies. Becker (1973) makes the point that the existence of polygamous unions is the equilibrium outcome when there is inequality among males. Boserup (1970) proposes that polygamy is explained by female productivity in agriculture, an idea tested by Jacoby (1995) in Cote d'Ivoire. Dalton and Leung (2014) argue that the greater prevalence of polygamy in West Africa is explained by the effect of the Atlantic slave trade on sex ratios.

A second group of works tries to explain the dynamics of polygamy. Anthropologists and historians have underlined the role of cultural factors, notably the hostility of Christian missions to polygamy. This is discussed, for example, by Egboh (1972) for Iboland, by Ekechi (1976) for West Africa more generally, and by Falen (2008) for Benin. In her work on Christianity and social change in Cameroon during the interwar period, Walker-Said (2018) documents how Christian missionaries attempted to transform African marital customs and entered into conflicts with the French colonial state, which in theory was also hostile to bridewealth and polygamy but did not want to alienate local chiefs (who

were needed to administer taxation and forced labor). Recently, economists have looked at this question more quantitatively. Fenske (2015) finds that areas that received a lot of missionary schools during the colonial period have lower rates of polygamy today. Kudo (2017) finds that in Malawi, rates of polygamy decrease as one moves away from the Livingstonia mission station. Becker (2022) studies the consequences of the Church's hostility to polygamy for educational expansion in Africa. He argues that the missions' insistence on monogamy lowered demand for education in places where polygamy was common, and he finds that the positive correlation between distance to a Christian mission and education is much smaller in ethnic groups traditionally practicing polygamy.

In discussing the dynamics of polygamy, economists typically put forward less-cultural factors. In Lagerlöf (2005), societies become monogamous as inequality among men falls. In Lagerlöf (2010), monogamy is modeled as a pacifying institution established by the elite to prevent the rebellion of poorer men (who do not get to marry in a polygamous society where rich men marry several women). In such a framework, the relationship between inequality and polygamy is hump-shaped. De La Croix and Mariani (2015) propose a unified model of the transition from polygamy to monogamy, and then to serial monogamy (divorce and remarriage), where income inequality among men and women plays a key role. Explanations relating the dynamics of polygamy to the dynamics of inequality are maybe more suited to Europe than to the recent decline of polygamy in Africa. Indeed, all available evidence points toward inequality remaining stable (at a high level) in Africa in the last 40 years (Ravaillon, 2014; Alvaredo et al., 2018). Gould et al. (2008) build a model where the dynamics of economic development, rather than the dynamics of inequality, explain the decrease in polygamy, as men substitute fewer educated children and wives for a large number of children and wives.

To the best of our knowledge, the only attempt besides ours to estimate the causal relationship between education and polygamy is Fenske (2015), who finds that a variety of recent education supply shocks that increased women's education did not decrease their likelihood of being in a polygamous union. At the same time, he finds that the regions that received a lot of government schools during the colonial period have lower

rates of polygamy today. The discrepancy between the two results asks the question of whether colonial-era education might have different effects on marriage markets and polygamy than education received post-independence in the era of mass education. One reason might be that the content of education changed; another possibility is that colonial education was of higher quality because it was restricted to a limited number of students. Contrary to Fenske (2015), we also consider the effect of education on polygamy for men, we distinguish between first wives and higher-order wives, and we use a structural model to take into account the role of assortative matching on education.

We also contribute to the literature on the interplay of labor and marriage-market returns to education for women. Goldin (1993) studies the changing meaning of college in the lives of American women over the 20th century. Chiappori et al. (2009) develop a model where education has both marriage-market and labor-market returns. Fewer papers focus on developing countries, but Zha (2020) studies the role of education in the marriage market in Indonesia, and Ashraf et al. (2020) study the role of the bride-price custom in explaining whether parents send their daughters to school in response to a school-expansion program. In Cameroon (where most ethnic groups have some form of bride price), we find that women’s education increased in response to school construction and that education had labor-market as well as marriage-market returns.⁴

Our paper contributes to the literature on female empowerment and economic development (Duflo, 2012), specifically the literature estimating the effect of education on women’s lives (see Heath and Jayachandran, 2018, for a recent review). Osili and Long (2008), Keats (2018), Chicoine (2012) and Chicoine (2016) use schooling-expansion reforms in Nigeria, Uganda, Kenya, and Ethiopia to show that education decreases fertility. In Keats (2018), Chicoine (2012) and Chicoine (2016), this decrease is partly explained by increased use of contraception. Dupas (2011) finds that Kenyan girls informed about the relative risks of HIV infection by age substitute away from older partners. Friedman et al. (2016) find, also in Kenya, that girls who prolonged their education as a result of a

⁴Of the 50 ethnic groups in the Ethnographic Atlas (Gray, 1999) we can locate in Cameroon using Murdock’s (1959) map, 43 traditionally practice bride price, corresponding to 96% of the population. Three groups practice bride service or token bride price (2.6% of the population). Four groups traditionally practice female relative exchange (1.5% of population).

merit scholarship program are less likely to accept domestic violence. Like Fenske (2015), we do not find that women who received education following a wave of public school construction are less likely to be in a polygamous union. In fact, in our context, we find they are *more likely* to be in a polygamous union, and we show the role of assortative matching on education in the marriage market in explaining this result. More generally, we show that taking matching into account is important when studying the returns to education on the marriage market.

Finally, our paper contributes to the literature on matching models of the marriage market, particularly the branch of this literature concerned with empirical estimation. Chiappori et al. (2012) estimate a model where individuals match on a single index aggregating all the characteristics of a mate. Reynoso (2019) extends this approach to the polygamous case. She assumes complementarity between co-wives skills. In her model, high-skilled and low-skilled women are co-spouses in polygamous households with high-skilled men, men and women of intermediary skills marry monogamously, and low-skilled men remain single. In this approach, women and men match on a single characteristic (skills), while our approach builds on models where men and women match on several characteristics: Choo and Siow (2006) estimate a model of matching on several discrete attributes of men and women, while Dupuy and Galichon (2014) extend the model to continuous attributes. We adapt Choo and Siow (2006) to a setting where men are allowed to marry several women and we show that, as in the monogamous case, the equilibrium is unique and maximizes total surplus. We characterize the equilibrium's matching function and we show that, as in Dupuy and Galichon (2014), we can identify the second derivatives of the joint utility function of a match with respect to the characteristics of men and women.

In the rest of the paper, we present the data (Section 2), the quasi difference-in-differences strategy and the estimated returns to education (Section 3), and the model of marriage and the results of the structural estimation (Section 4).

2. Data

To identify the effect of education on the marriage market in Cameroon, we use two main sources of data: (1) the full-count 1976 population census of Cameroon, and (2) administrative data from 2016 giving, for every school in Cameroon, its date of opening, status (public or private), and the name of the locality.⁵ Because this is not historical data, it gives us information about historical school supply insofar as there was no attrition — schools, once opened, did not close. During Cameroon’s rapid population growth (it increased sevenfold between 1900 and 2014) and ever-expanding school supply, this assumption seems reasonable, at least for public schools. In the Data Appendix, we cross-validate our source with historical data for the colonial period. We compare the yearly flows and stocks of public and private schools in Cameroon during the colonial period in historical data with the flows and stocks implied by our retrospective data. We find that the two series align well for public schools, but not for private missionary schools.

We were able to precisely geolocate villages and schools in the two datasets (see the Data Appendix).⁶ This allowed us to compute the stock and flows of schools in a radius of 10 km around each village for every year of the 20th century. Knowing the birth year and village of residence of individuals in 1976, we could then compute the stock and flow of schools in their village at each age. The census gives no information on the village of birth, but does give information on the district of birth.⁷ For individuals still residing in their district of birth, we assume that they were schooled in the village in which they were living in 1976. Out-of-district migrants, representing roughly a third of the sample, are excluded from our main specification. This procedure creates two concerns: (1) individuals who stayed in their district might have moved within the district between their primary school years and 1976, and (2) education, our independent variable of

⁵The Data Appendix describes this source in greater detail and discusses the procedure used for constructing variables such as spouse characteristics, which are not given directly.

⁶Village-level geolocation was impossible in the 16 districts of the Bamiléké region; we therefore excluded from our estimations individuals born in these districts (14% of the population in 1976). We also excluded from our estimations individuals born in Yaoundé (the administrative capital) and Douala (the economic capital) because of the difficulty of precisely geolocating enumeration areas within these agglomerations. Yaoundé and Douala represented roughly 5% of the population in 1976.

⁷There were 12,125 villages and 138 districts in Cameroon in 1976; see Appendix Figure A.2.

interest, is likely to affect the decision to migrate, so that the sample of nonmigrants might be selected. To address these two concerns, we also present results using district-level variation in the number of available schools.⁸ These results are estimated both on the sample of nonmigrants, and on the full sample, including out-of-district migrants (see Section 3.5 and Appendix Tables C.4 and C.5).

Table 1 presents some descriptive statistics for the sample of men and women older than 15 in 1976: men have 3.0 years of education on average, versus 1.4 for women; 17% of men are wage earners, versus only 1% of women.⁹ Because of polygyny, the percentage of married men (56%) is lower than the percentage of married women (67%). Married women are on average 10 years younger than their husband. This is important for our identification strategy, because it means that the opening of a school in a village does not necessarily affect both groups of potential mates in the same way (see below, Section 3.4). This large average age difference explains why the percentage of widows is much higher among women than men (14% vs 2%). Forty-four percent of married women are in a polygamous union, versus 24% of men; 71% of polygamists have two wives, 19% have three wives and 10% have four wives or more. People born before 1940 had on average 0.26 schools in a radius of 10 km around their village when they were 7 (0.12 public and 0.14 private). People born after 1940 had on average 1.5 public schools and 1.25 private schools.¹⁰

Appendix Figure B.1 depicts the geographical distribution of polygamy. A mild north-south gradient is evident, but polygamy is prevalent in every district. The share of married women in a polygamous union is below 20% only in a handful of districts around the economic capital (Douala) and the administrative capital (Yaoundé).

In addition to population census and administrative school data, we use a variety of other geographical and historical data sources to build village-level controls. These

⁸More precisely, we use the district average number of available schools for nonmigrants.

⁹The census contains no observation on wages. In the paper, we use information on whether an individual is working for a wage as a coarse proxy for socioeconomic status. In the context of a developing country in the 1970s, earning a wage rather than working as an independent should be interpreted as a sign of a higher socioeconomic status. See the Data Appendix.

¹⁰The slight discrepancy between men and women is due to differences in gender composition across villages explained by different migration patterns. These figures are computed for nonmigrant only, and men tend to migrate more than women.

Table 1: Descriptive Statistics

	Women Older Than 15		Men Older Than 15	
	Mean	Observations	Mean	Observations
<i>Full sample</i>				
Age	35.13	1,974,625	35.86	1,754,334
Years of schooling	1.37	1,964,294	2.97	1,747,632
Wage earner	0.01	1,941,572	0.17	1,711,810
Out-of-district migrant	0.30	1,975,117	0.33	1,754,807
Married	0.67	1,966,349	0.56	1,740,155
Single	0.16	1,966,349	0.39	1,740,155
Widow	0.14	1,966,349	0.02	1,740,155
Divorced	0.03	1,966,349	0.03	1,740,155
<i>Married sample</i>				
Age	33.12	1,310,614	42.98	975,344
In a polygamous union	0.44	1,184,845	0.24	975,468
No. of wives			1.36	970,940
<i>Sample in a polygamous union</i>				
No. of wives			2.46	236,263
2 wives			0.71	236,263
3 wives			0.19	236,263
4 wives or more			0.10	236,263
<i>Nonmigrant sample</i> (excluding Yaoundé, Douala, and the Bamiléké)				
Schools in 10-km radius at age 7				
born before 1940				
public	0.12	453,701	0.11	464,053
private	0.14	453,701	0.11	464,053
born after 1940				
public	1.48	689,018	1.57	551,207
private	1.25	689,018	1.34	551,207

Sample: all men and women older than 15 in 1976.

sources are described in greater detail in the Data Appendix.

3. Estimation of Returns to Education

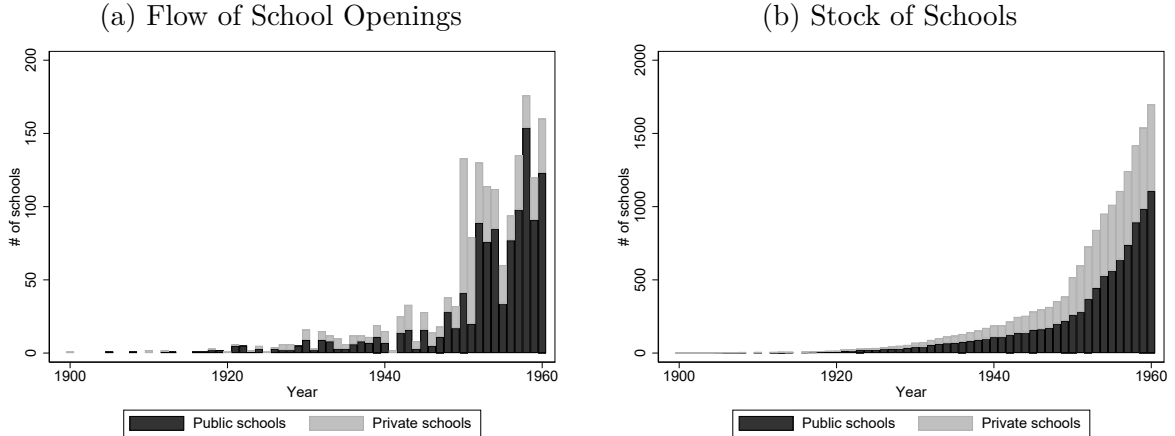
This section measures the marriage-market and labor-market returns to education. The endogeneity concerns in estimating labor-market returns to education are well known. As for marriage-market returns, it is extremely likely that a number of cultural characteristics and personal unobservables determine both educational choices and marriage-market outcomes, especially the probability of being in a polygamous union (Fenske, 2015). Like a number of papers following Duflo (2001), we take advantage of an expansion in school supply — the wave of primary school construction in Cameroon that started after World War II.

3.1. Historical Background

After World War I, German Cameroon was divided between the British and the French and administered under mandates from the League of Nations. In both parts, before World War II, public expenditure on education was low and Christian missionaries (Protestant and Catholic) were the main providers of education (Dupraz, 2019). After World War II, education expenditure increased massively in both parts. In British Cameroon, real expenditure per school-age child quadrupled between 1937 and 1955. In French Cameroon, it grew by a factor of 30 (Dupraz, 2019). As the stated goal of education policy shifted from educating a small elite to providing universal primary education, the colonial governments of British and French Cameroon increased subsidies to missionary schools and started building more public schools. School construction continued after both parts of Cameroon gained independence in 1960 and were reunited in 1961, but this construction does not contribute to our estimation strategy.

The first panel of Figure 1 displays the annual number of school openings in Cameroon for each year between 1900 and independence in 1960 according to our administrative data. The annual number of school openings started increasing around 1950, from an

Figure 1: Stock and Flow of Public and Private Schools in Cameroon, 1900–1960



Note: The bars representing private schools are stacked on the bars representing public schools so that the graph shows the total number of schools.

average of 16 between 1930 and 1949 to an average of 119 in the 1950s. As a result, the total number of schools in Cameroon increased sixfold from 1945 (283) to 1960 (1,699) — see the second panel of Figure 1. The yearly number of *public* school openings rose from an average of nine between 1930 and 1949 to an average of 81 in the 1950s, and the total number of public schools in Cameroon was seven times higher in 1960 (1,107) than in 1945 (158).

Our analysis always distinguishes between public and private schools, for several reasons. First, as discussed above, it is likely that our data are incomplete for private schools, giving us only the schools that were of sufficient quality to survive until 2016, and second, we have every reason to expect heterogeneous effects of public and private education on the marriage market, especially on polygamy. Private schools opening in Cameroon before 1960 were almost all Christian, missionary schools.¹¹ There is ample evidence that Christian missions were targeting polygamy specifically, putting a lot of effort into promoting monogamy (Egboh, 1972; Ekechi, 1976; Oduyoye, 1993; Falen, 2008). In Cameroon, the African clergy was particularly active in criticizing elements of marital customs such as bridewealth and polygamy (Walker-Said, 2015, 2018). Catholic missionaries established “sixas” or “bride schools” to prepare young girls for a Christian

¹¹In our data, out of 591 private schools built before 1960, 580 (98%) are Christian missionary schools, four are private Islamic schools, and seven are secular private schools.

wedding, and to welcome converted women who wanted to divorce a polygamous husband (Tsoata, 1999; Walker-Said, 2015, 2018). What was the attitude of the colonial government, and of nonreligious colonial schools? The French administration viewed polygamy negatively. In the 1920s, it passed marital laws granting Christian women the right to divorce their polygamous husband (Walker-Said, 2018).¹² However, the French colonial government never criminalized polygamy, and the marital laws were imperfectly enforced because the government did not want to alienate local chiefs, who played an important role in administering taxation and forced labor. Apparently, in contrast to Christian schools, public schools didn't actively fight polygamy. We investigated the portrayal of families in *Mamadou et Bineta lisent et écrivent couramment* (Davesne, 1951), a reading textbook used throughout French Africa from the 1930s onward, and we found several nonjudgmental mentions of polygamous households (though the protagonists Mamadou and Bineta live in a monogamous family).¹³ However, we cannot completely exclude the fact that public school teachers, especially European ones, imparted a negative view of polygamy to their students.

3.2. Event Study and Common Trend Assumption

To obtain plausibly exogenous variation in the education of men and women, we use local, village-level variation in the number of schools over time. Our main estimation strategy is a quasi difference-in-differences (which we present in Section 3.3). Difference-in-differences estimates are sensitive to the common trend assumption, which we test with an event study. This event study estimates the effect of school openings at different ages. This checks that school construction increased enrollment only for the children of school age after the school was built. Put another way, this event-study tests whether school construction took place in areas with different trends in education. One concern is that the government was responding to local demand for education and that schools were built in villages that were already seeing an increasing educational trend. If that were

¹²These laws targeted other aspects of Cameroonian marriage customs, giving widows the right to marry who they wanted, and setting limits to the bride price.

¹³p. 43-45: Karibi, the protagonist of a tale, has several wives; p. 73: describes the huts of a compound and mentions “the room of wives” (*la chambre des femmes*).

the case, one would estimate a positive correlation between education and the opening of a school after school entry age (say, at 25). We estimate the following equation:

$$E_{ivc} = \alpha_v + \delta_c + \delta_c \times BR + \sum_{a=-10}^{a=30} \left(\beta_{public,a} n_{vc}^{public,a} + \beta_{private,a} n_{vc}^{private,a} \right) + x'_{vc} \theta + e_{ivc} \quad (1)$$

E_{ivc} is the education (in years) of individual i , born in village v in year c , and α_v and δ_c are village and cohort fixed effects. To allow the trend in education to be different in British Cameroon, we interact the cohort fixed effects with a dummy BR equal to one in the two Western provinces, which were part of British Cameroon until independence. $n_{vc}^{public,a}$ is the number of public schools that opened in village v when an individual born in year c was age a (negative numbers are years before birth). $n_{vc}^{public,-10}$ is the stock of public schools 10 years before birth.¹⁴ $n_{vc}^{private,a}$ is the same for private schools. For example, for individual i born in 1940 in village v where the only school, public, opened in 1947, $n_{v,1940}^{public,7} = 1$, and all other school-opening variables are equal to zero. We do not consider $a > 30$; individuals who had a school opening after 30 serve as the reference.¹⁵ x_{vc} is a vector of time-invariant village controls interacted with a quartic cohort trend (see Section 3.3 below for a detailed discussion). We estimate equation (1) separately for men and women. Standard errors are clustered at the village level.¹⁶

Equation (1) is similar to an event-study specification, with a few differences: the date of the event (the opening of a school) is not the same in each village, there can be several events per village, and the effect of the event is estimated for different age cohorts rather than at different time periods. In theory, Cameroonian children were supposed to attend primary school from 7 to 12, but the colonial authorities were flexible regarding school entry age: according to a 1950 decree, children could start primary school as late as

¹⁴To make the graph more readable, we do not consider schools built more than 10 years before birth, which is why $n_{vc}^{public,-10}$ is the stock, rather than the flow, of schools in the village 10 years before birth: it takes into account all schools opening more than 10 years before birth.

¹⁵If we were to consider all $a > 30$, the sum of every $n_{vc}^{public,a}$ would be the number of public schools, a constant at the village level, and would be captured by the village fixed effects.

¹⁶When showing the first-stage and reduced-form results (Tables 4 and 5), we also show Conley standard errors allowing for serial and spatial autocorrelation, but the procedure is computationally very demanding with many fixed effects and a large sample size, so we do not adopt it systematically throughout the paper.

10 and enter the final grade as late as 16.¹⁷ Furthermore, parents might have bypassed official regulations to allow their children to benefit from the opening of a school in their village. There are no detailed statistics on late school entry during the colonial period, but Appendix Figure B.2 shows school attendance by age in the 1976 census for boys and for girls. For both girls and boys, attending the first year of primary school was quite common until the age of 12 (but became negligible after). A large percentage of boys and girls older than 12 were still attending primary school. There is also some error in age in the census, notably in the form of age heaping (see Appendix Figure B.3). We therefore expect school openings to have a positive effect on education before age 7, but also to a certain extent between 8 and 16 (because of late school entry, and because the opening of an additional school can alleviate capacity constraints in an existing school). However, school openings after age 17 should not have a positive effect on education.¹⁸ Because individuals who had a school opening after age 30 serve as the reference group, this event study is similar to a test of parallel pretrends in a more classical difference in differences setting.

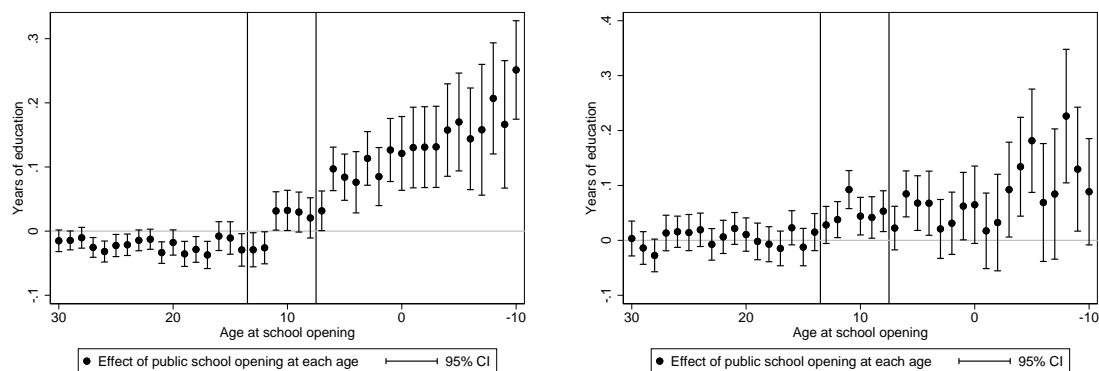
In estimating equation (1), and in the rest of the paper, we consider only the schools that opened during the colonial period, before 1960. These are the schools that mattered for the individuals we consider in the rest of the paper (older than 25 in 1976, born before 1951).

Figure 2 displays the β associated with public education, for men and women separately. Schools opening before age 7 increase education of both men and women. The effect is larger and more precisely estimated for women than for men. If, when a public school opened in a village, some boys were already attending private school while girls were not, we would expect the opening of a public school to matter more for girls' education. Schools opening between ages 8 and 12 also increase education, but the magnitude is

¹⁷*Arrêté du 27 juillet 1950*, reproduced in France, Ministère des Colonies (1950). *Rapport annuel du Gouvernement Français à l'Assemblée Générale des Nations Unies sur l'administration du Cameroun placé sous la tutelle de la France*.

¹⁸We pick age 17 because children could enter the final grade of primary school as late as 16, according to the 1950 decree. Because of late school entry and age heaping, it is hard to determine an exact threshold. Picking an older threshold confirms the conclusion that school openings matters only for people young enough to go to school.

Figure 2: Event-Study Graphs: Effect of Public School Openings on Education
(a) Women (b) Men



Note: Both figures display the β_{public} coefficients of equation (1), estimated separately for men and women. The $\beta_{private}$ are estimated jointly, but for readability, they are presented on a different graph (Appendix Figure C.1). Standard errors are clustered at the village level.

lower for women. We find no evidence that schools responded to local demand and were built in villages that were already on an upward educational trend. If it were the case, we would estimate positive β coefficients for schools opening after age 17 (remember that the reference group are individuals who had a school opening after age 30). However, the β coefficients for schools opening after age 17 are very close to zero (and if anything slightly negative for women). The picture is similar for private schools (Appendix Figure C.1).

Appendix Figure C.2 displays event study graphs for men and women when we also consider postcolonial schools (opened after 1960). These graphs are very similar to Figure 2, but there is a significant (though small) correlation between years of education and the number of public schools built in the village when an individual was 20 to 30 years old. This is likely because, in the postcolonial period, school supply reacted to local demand and educated parents of school-age children were able to lobby for more schools. For this reason, we never consider schools built after 1960 in the rest of the paper. This hardly matters for our results because we consider only individuals older than 25, who were born before 1951.

Next, we present event-study results in a more compact way and we test more formally the parallel trend assumption that schools opening after age 17 are not positively cor-

related with education or labor-market or marriage-market outcomes. We estimate the following equation:

$$\begin{aligned}
E_{ivc} &= \beta_1 N_{vc}^{public,7} + \beta_2 n_{vc}^{public,8-12} + \beta_3 n_{vc}^{public,13-16} + \beta_4 n_{vc}^{public,17-22} \\
&+ \pi_1 N_{vc}^{private,7} + \pi_2 n_{vc}^{private,8-12} + \pi_3 n_{vc}^{private,13-16} + \pi_4 n_{vc}^{private,17-22} \\
&+ \alpha_v + \delta_c + \delta_c \times BR + x'_{vc} \theta + e_{ivc}
\end{aligned}$$

where $N_{vc}^{public,7}$ is the stock of public schools in village v when individual i was 7, and $n_{vc}^{public,a-b}$ is the number of public school openings between ages a and b .¹⁹ β_4 and π_4 test the parallel-trend assumption for, respectively, public and private schools. A positive β_4 (π_4) would indicate that individuals aged 17 to 22 at the opening of a public (private) school were already on a positive educational trend. β_3 and π_3 should be small, but they might be positive if there is late school entry and measurement error in age. We estimate equation (2) not only for education, but also for a labor-market outcome (the probability of being a wage earner) and for the number of spouses (for married men) and co-spouses (for married women).²⁰

Table 2 displays the coefficients of equation (2) for women. An additional public school in the village at age 7 increases schooling by 0.11 years and the probability of being a wage earner by 0.16 percentage points. It also increases the number of co-spouses of married women by 0.02 (we obtain very similar results in Appendix Table C.2 when we also consider single, divorced, and widowed women with a number of co-spouses equal to zero). This is the central result of the paper: a positive education shock makes women *more*, not *less* likely to enter a polygamous union. Public schools did not open in villages that were already on a positive educational trend: consistent with Figure 2, the coefficient

¹⁹ $N_{vc}^{private,7}$ and $n_{vc}^{private,a-b}$ are the same for private schools.

²⁰ For marriage-market outcomes, we might estimate β_4 and π_4 different from zero even if the parallel trend assumption is valid because of general equilibrium effects on local marriage markets. If the building of a school makes women who were young enough to attend school relatively more attractive on the local marriage market, it also makes women who were too old to attend school relatively less attractive. The same general equilibrium argument can be applied to labor-market outcomes, but labor markets are typically larger than marriage markets, which makes the problem less important.

Table 2: Effect of School Construction After Age 17, Women

	(1)	(2)	(3)
	Years of Schooling	Wage Earner	No. of Co-spouses
No. public schools at 7	0.1110*** (0.0167)	0.0016*** (0.0006)	0.0197** (0.0082)
No. public sch. openings 8-12	0.0331*** (0.0122)	0.0009** (0.0004)	-0.0010 (0.0064)
No. public sch. openings 13-16	-0.0087 (0.0094)	-0.0001 (0.0003)	0.0030 (0.0065)
No. public sch. openings 17-22	-0.0166*** (0.0057)	-0.0001 (0.0002)	-0.0057 (0.0057)
No. private schools at 7	0.1019*** (0.0171)	0.0023*** (0.0006)	-0.0519*** (0.0135)
No. private sch. openings 8-12	0.0166 (0.0119)	0.0007 (0.0005)	-0.0419*** (0.0132)
No. private sch. openings 13-16	-0.0183** (0.0091)	0.0006* (0.0003)	-0.0339*** (0.0122)
No. private sch. openings 17-22	-0.0167** (0.0071)	0.0007* (0.0003)	-0.0328*** (0.0110)
Village fixed effects	✓	✓	✓
Cohort fixed effects	✓	✓	✓
Cohort fixed effects × Br. Cameroon	✓	✓	✓
Village controls × cohort quartic trend	✓	✓	✓
Observations	700,986	693,045	490,045

Notes: Sample: In columns (1) and (2), nonmigrant women aged 25–60 in 1976 (the slight discrepancy in number of observations between columns (1) and (2) is explained by missing employment status); in column (3), all nonmigrant married women 25–60. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

on the number of public schools opening between ages 17 and 22 is actually negative. The effect is statistically significant but very small: -0.016 years of schooling, about a seventh of the effect of public schools²¹ at age 7. Public schools did not open in villages that were on a differential trend in the percentage of wage earners or polygamy: in columns (2) and (3), the coefficient on the number of public schools opening between ages 17 and 22 is statistically indistinguishable from zero.

The coefficients on private schools are comparable to the coefficients on public schools for education and the probability of being a wage earner, but they are very different for the number of co-spouses of married women. An additional private school in the village

²¹Table C.10 shows the robustness of our results to small amounts of endogeneity; see the discussion in Section 3.5.

at age 7 *decreases* the number of co-spouses by 0.05. Given that Christian schools in Cameroon were actively fighting polygamy (Tsoata, 1999; Walker-Said, 2015, 2018), one natural interpretation of this result is that Christian and secular education have very different effects on the marriage market. However, we find evidence that the parallel-trend assumption does not hold in the case of private, Christian schools. The opening of a Christian school in the village between ages 17 and 22 is associated with 0.03 fewer co-wives, which means that Christian schools were opening in villages where polygamy was already declining, probably because these were already Christianized villages.

Table 3: Effect of School Construction After Age 17, Men

	(1)	(2)	(3)
	Years of Schooling	Wage Earner	No. of Spouses
No. public schools at 7	0.0572*** (0.0158)	0.0087*** (0.0023)	0.0111** (0.0047)
No. public sch. openings 8-12	0.0524*** (0.0112)	0.0017 (0.0013)	0.0098*** (0.0036)
No. public sch. openings 13-16	0.0143 (0.0103)	0.0028* (0.0014)	0.0065* (0.0034)
No. public sch. openings 17-22	0.0012 (0.0085)	0.0012 (0.0011)	0.0017 (0.0032)
No. private schools at 7	0.0847*** (0.0200)	0.0117*** (0.0019)	0.0176** (0.0072)
No. private sch. openings 8-12	0.1005*** (0.0202)	0.0113*** (0.0019)	0.0070 (0.0059)
No. private sch. openings 13-16	0.0330** (0.0167)	0.0066*** (0.0019)	0.0074 (0.0063)
No. private sch. openings 17-22	0.0233 (0.0146)	0.0081*** (0.0017)	0.0080 (0.0062)
Village fixed effects	✓	✓	✓
Cohort fixed effects	✓	✓	✓
Cohort fixed effects × Br. Cameroon	✓	✓	✓
Village controls × cohort quartic trend	✓	✓	✓
Observations	608,454	602,210	470,247

Notes: Sample: In columns (1) and (2), nonmigrant men aged 25–60 in 1976 (the slight discrepancy in number of observations between columns (1) and (2) is explained by missing employment status); in column (3), all nonmigrant married men 25–60. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 3 displays the coefficients of equation (2) for men. The results are broadly in line with Table 2 for women, but a couple of important differences emerge: first, an additional

public school in the village at age 7 increases education by 0.06 years (versus 0.11 for women), maybe because boys were more likely to be already enrolled in a private school; and second, the number of schools opening between ages 13 and 16 increases education, the probability of being a wage earner, and the number of co-spouses, maybe because boys were more likely than girls to start school late.²² Men who had access to a private school at age 7 have more spouses, not less, but the parallel-trend assumption seems to be violated: private schools opened in villages with a positive trend in wage labor (column 2).

3.3. First Stage and Reduced Form

To estimate the labor-market and marriage-market returns to education, we use the stock of public schools in the village when an individual was of school age as an instrument for education. In a first stage, we estimate the following equation:

$$E_{ivc} = \alpha_v + \delta_c + \delta_c \times BR + \beta N_{vc}^{public,a} + x'_{vc} \theta + \varepsilon_{ivc} \quad (2)$$

where the excluded instrument $N_{vc}^{public,a}$ is the stock of public schools in village v when an individual born in year c was age a . We choose the age a that maximizes the first stage F-test of the excluded instrument. This age is 7 for women, and 13 for men (see Appendix Figure C.3). As discussed in Section 3.2 above, the reason why schools opening after age 7 seems to matter more for men than for women is that late school entry is a more common phenomenon for men.²³ As a robustness exercise, we also present results where we use the same excluded instruments for men and women: the stock of schools at age 7, and the number of schools opening between ages 8 and 13 (see Appendix Tables C.11). As in equation (1), α_v and δ_c are village and cohort fixed effects, and BR is a

²²The magnitude of the effect on the number of spouses is important (0.0065 wives). This might be explained by the fact that, in the cohort of boys theoretically too old to enter school when a new school opened, the few who attended school enjoyed a particularly big advantage in the marriage market later in life, while those who obtained education along with the rest of their cohort had to compete with other educated men.

²³In the 1976 census, 46% of boys in the first year of primary school are older than 10 versus 38% for women.

binary equal to one in the two Western provinces, which were part of British Cameroon until independence. x_{vc} is a vector of time-invariant village controls interacted with a quartic cohort trend. Standard errors are clustered by village, and we also provide Conley standard errors allowing for serial and spatial autocorrelation.²⁴

The village fixed effects capture any village characteristic correlated with both education and school construction. One might be worried, for example, that more schools are built in urban areas, where the returns to education are larger. Controlling for village fixed effects will perhaps be even more important in the second stage, when we will put marriage-market outcomes on the left-hand side. Fenske (2015) shows that, in a number of African surveys, the negative correlation between polygamy and education is largely explained by geographical controls correlated negatively with education and positively with polygamy. The cohort fixed effects ensure that we will not interpret a spurious correlation between an increasing trend in education (or a declining trend in polygamy) and an increasing trend in school supply. The cohort fixed effects are interacted with the British Cameroon binary to allow for a different trend in regions subjected to a different colonial rule. Because of the village fixed effects, our identification comes from within-village differences between age groups. However, villages sharing similar observable characteristics may have similar histories, and this may be related to trends in school construction. This is why our specification includes the vector x_{vc} of time-invariant village controls interacted with cohort trends.²⁵ The village-level controls are precipitation, temperature, elevation, ruggedness, a malaria stability index and agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, mission station in Roome (1925), 1913 German mission school, and 1913 German government school (see Data Appendix).²⁶ In Appendix Table C.1, we show that the coefficients of the first stage and reduced form

²⁴See Conley (1999, 2008). We use a distance cutoff of 100 km, a cutoff for serial correlation of 20 years, and a Bartlett kernel. To implement Conley standard errors, we make use of the very useful code provided by Hsiang (2010) and Fetzer (2014).

²⁵We use a quartic cohort polynomial, which is less computationally demanding than interacting the village controls with the vector of cohort fixed effects (but we show in Tables 4 and 5 below that the two approaches give very similar results).

²⁶Jedwab et al. (2018) have shown that historical mission atlases such as Roome (1925) tend to report only selected mission locations, hence the importance of also using the data on mission schools in Schlunk (1914), digitized by Dupraz (2019).

are invariant to also including ethnic-group-level variables from the *Ethnographic Atlas* (precolonial centralization and whether polygyny is general or limited); these variables are unfortunately available for only two-thirds of the villages in our data. The vector x_{vc} also contains the number of private Christian schools in the village at a . The stock of private schools is not used as an excluded instrument, because the parallel-trend assumption is less credible for private schools, but we include it as control to account for potential complementarity or substitution between private and public schools.²⁷

Equation (2) is a quasi difference-in-differences specification. It would be similar to a staggered difference-in-differences if $N_{vc}^{public,a}$ was a binary variable. In this case, we would be comparing the education gap of young and old cohorts between villages where a school opened and villages where no school opened. Because $N_{vc}^{public,a}$ is a count variable, we also use the difference in the education gap of young and old cohorts between villages where N schools opened and villages where only $N - 1$ schools opened.²⁸ A recent literature shows that when the time of treatment varies, this type of two-way fixed-effects model can produce biased estimates in the presence of heterogeneous effects (de Chaisemartin and D’Haultfoeuille, 2020; Goodman-Bacon, 2021). We show the robustness of our results to using the estimator of de Chaisemartin and D’Haultfoeuille (2020), which solves the issue (see Section 3.5 below).

Table 4 displays the results of both the first stage and the reduced form on the number of co-wives for married women. (In Section 3.4, we obtain very similar results when we also consider single, divorced and widowed women with a number of co-spouses equal to zero.) Column (1) is a simplified specification, where the village controls are interacted with a linear, rather than quartic, cohort trend. One additional school in the village at age 7 increases education by 0.096 years of education. Column (2) is our preferred specification, where we interact the village controls with a quartic cohort trend: the

²⁷If public and private schools are substitutes, the government might build fewer public schools in villages that already have private schools. The government might also build more public schools in villages that already have private schools and where the demand for education is high.

²⁸In our sample of women, 72% had no school in the village at age 7, 20% had one school, 5% had two schools, and 3% had three schools or more, with a maximum of 11 schools. In our sample of men, 62% had no school in the village at age 13, 22% had one school, 8% had two schools, and 8% had three schools or more, with a maximum of 12 schools.

Table 4: First Stage and Reduced Form for Married Women

	Dep. var.: years of education				Dep. var.: No. of (co-)wives			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
No. public schools at 7	0.0957*** (0.0108) [0.0151]	0.0852*** (0.0096) [0.0098]	0.0825*** (0.0096) [0.0099]	0.0857*** (0.0094) [0.0099]	0.0176*** (0.0063) [0.0053]	0.0188*** (0.0067) [0.0056]	0.0180*** (0.0069) [0.0057]	0.0205*** (0.0068) [0.0056]
No. private schools at 7	0.1178*** (0.0149) [0.0153]	0.0943*** (0.0169) [0.0164]	0.0942*** (0.0168) [0.0162]	0.0828*** (0.0151) [0.0151]	-0.0289*** (0.0083) [0.0068]	-0.0293*** (0.0087) [0.0071]	-0.0301*** (0.0087) [0.0071]	-0.0257*** (0.0086) [0.0072]
Village fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
Cohort fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
Cohort fixed effects × Br. Cameroon	✓	✓	✓		✓	✓	✓	
Village controls × cohort trend	✓				✓			
Village controls × cohort quartic trend		✓		✓		✓		✓
Village controls × cohort fixed effects			✓				✓	
Cohort fixed effects × province fixed effects				✓				✓
Observations	490,045	490,045	490,045	490,045	490,045	490,045	490,045	490,045

Notes: Sample: All nonmigrant married women ages 25–60 in 1976. Standard errors clustered at the village level are in parentheses. Conley standard errors allowing for spatial and serial autocorrelation are in brackets (distance cutoff of 100 km, cutoff for serial correlation of 20 years, Bartlett kernel). * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

coefficient drops slightly, to 0.085 years. It then hardly changes when we interact the village controls with the full vector of cohort fixed effects in column (3), which shows that this computationally demanding estimation is not required. In column (4), we interact the vector of cohort fixed effects with a vector of province fixed effects (there were seven provinces in Cameroon in 1976; see Appendix Figure A.2), and the coefficient on the stock of public schools at age 7 is barely affected. This reassures us that our first stage is capturing the causal effect of school supply on education rather than a spurious correlation between trends in education and province-specific trends in school construction. Columns (5) to (8) display the results of the reduced form: an additional public school in the village at age 7 increases the number of co-wives by about 0.02 (while an additional private school decreases the number of co-wives by about 0.03). Table 5 displays the results of the first stage and of the reduced form on the number of wives for married men. An additional public school at age 13 increases education by about 0.05 years and the number of wives by about 0.01.

3.4. Labor-Market and Marriage-Market Returns to Education

We use our first stage to instrument for education in the following equation:

$$y_{ivc} = \alpha_v + \delta_c + \delta_c \times BR + \tau E_{ivc} + x'_{vc} \phi + \nu_{ivc} \quad (3)$$

Table 5: First Stage and Reduced Form for Married Men

	Dep. var.: years of education				Dep. var.: No. of wives			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
No. public schools at 13	0.0431*** (0.0100) [0.0100]	0.0495*** (0.0098) [0.0089]	0.0483*** (0.0098) [0.0088]	0.0571*** (0.0099) [0.0087]	0.0082*** (0.0027) [0.0024]	0.0089*** (0.0029) [0.0024]	0.0091*** (0.0029) [0.0024]	0.0120*** (0.0030) [0.0024]
No. private schools at 13	0.0795*** (0.0181) [0.0166]	0.0756*** (0.0178) [0.0162]	0.0749*** (0.0180) [0.0161]	0.0629*** (0.0175) [0.0159]	0.0100** (0.0047) [0.0038]	0.0104** (0.0048) [0.0039]	0.0103** (0.0048) [0.0039]	0.0118*** (0.0044) [0.0035]
Village fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
Cohort fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
Cohort fixed effects × Br. Cameroon	✓	✓	✓		✓	✓	✓	
Village controls × cohort trend	✓				✓			
Village controls × cohort quartic trend		✓		✓		✓		✓
Village controls × cohort fixed effects			✓				✓	
Cohort fixed effects × province fixed effect				✓				✓
Observations	470,247	470,247	470,247	470,247	470,247	470,247	470,247	470,247

Notes: Sample: All nonmigrant married men aged 25–60 in 1976. Standard errors clustered at the village level are in parentheses. Conley standard errors allowing for spatial and serial autocorrelation in brackets (distance cutoff of 100 km, cutoff for serial correlation of 20 years, Bartlett kernel). * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

where y_{ivc} can be an individual’s own labor-market outcome (for example, the probability of being a wage earner) or marriage-market outcome (for example, the number of wives/co-wives, or education of the spouse).

For the exclusion restriction to be met, it is not enough that the parallel-trend assumption holds. It must also be the case that school construction affects outcome y_{ivc} only through the own education of individual i . However, because of matching on age, the construction of a school in one’s village also affects the education of one’s potential mates, who tend to be of similar ages. This is particularly important when we estimate whether one’s own education causes that person to choose a more-educated spouse (i.e., matching on education). If husbands and wives always married within the same age cohort, it would be impossible to disentangle the matching on education from the matching on age and the direct effect of school construction on spouse education. In Cameroon around independence, individuals did not typically marry within the same age cohort — husbands were on average 10 years older than their wives (see Table 1). However, the stock of schools when an individual was of school age and the stock of schools when their potential mates were of school age are highly correlated. The necessity of disentangling matching on age and marriage-market returns to education is one additional motivation for using the structural model of Section 4, which allows us to estimate matching on education (and the affinity between polygamy and education) taking into account

matching on age. Before we get to that model, though, as a first step, we will present the instrumental variable estimates. In Section 3.5 below, we probe the sensitivity of these IV estimates to controlling for the stock of schools when potential husbands were of school age, and we find the effect of women’s education on their likelihood of being in a polygamous union nearly unchanged.

Selective migration is a second threat to the validity of the exclusion restriction. In equations (2) and (3), $N_{vc}^{public,a}$ is the stock of public schools at age a in the village where the person *lived in 1976*. Because we know the district of birth of individuals, we can exclude out-of-district migrants from the sample (about 30% of men and women). If education affects the decision to migrate, then our sample might be selected. However, we show that we obtain qualitatively similar results when we define the educational shock at the district level and use the full sample, including migrants (see Appendix Tables C.4 and C.5 and discussion in Section 3.5).

Table 6 presents estimations of labor-market and marriage-market returns to education for women (Panel A) and men (Panel B). We first present the results of the reduced form, then returns to education estimated by OLS (conditioning on the same set of controls as in the reduced-form equation, including cohort and village fixed effects), followed by returns to education estimated by 2SLS, using as the excluded instrument the number of schools in the village at age 7 (for women) or age 13 (for men).

We first consider labor-market returns. We use information on employment status, because the census gives no information on income.²⁹ In the context of a developing country in the 1970s, earning a wage rather than working as an independent should be interpreted as a sign of higher socioeconomic status (see discussion in the Data Appendix). One additional year of education increases the likelihood of earning a wage by 1.4 percentage points for women and 8.7 percentage points for men (Table 6, column 2). The larger effect for men is likely due to many formal wage-earning positions being closed to women in 1976 Cameroon. For both men and women, the positive effect comes from a higher likelihood of earning a wage conditional on being in the labor force, and not from higher

²⁹We also find asset indices unattractive here: these indices proxy the joint wealth of the household and are not specific to the contribution of either household member.

labor-force participation (Appendix Table C.2, column 1): for women, an additional year of education actually decreases the likelihood of being in the labor force, though the effect is not statistically different from zero. The much lower labor-market returns to education for women than for men are important for understanding our results on marriage and polygamy. In the near absence of labor-market returns to education, marriage-market returns to education should become a particularly important dimension, for both women and their parents.³⁰

We next turn to the effect of education on the extensive margin of marriage: an additional year of schooling slightly decreases the likelihood of having ever been married for women (-0.02) and increases it for men (0.04), but these effects are not statistically different from zero in the 2SLS specification (Table 6, column 2).

The last four columns of Table 6 focus on the sample of married individuals. One additional year of schooling increases the education of a woman's husband by 0.56 years and the average education of a husband's wives by 1.34 years (column 3). This large effect is likely partly explained by matching on age and the direct effect of school construction on the education of a husband's potential wives, who are typically younger. This problem motivates the structural estimation in Section 4, which allows us to separate the matching on education from the matching on age.³¹ One additional year of schooling increases the likelihood of being in a polygamous union by 6.6 percentage points for women, and by 10.1 percentage points for men (column 4). Men's number of wives increases by 0.18 while women's number of co-wives increases by 0.22 (column 5). One additional year of education decreases the age of the husband for women (by one and a half years), but it does not decrease the average age of the wives for men (column 6).

It is worth underscoring the difference between OLS and 2SLS results as far as polygamy is concerned. While for labor-market returns to education, OLS and 2SLS estimates have broadly the same sign and magnitude, it is not the case for polygamy. For women,

³⁰It is possible that our results on marriage-market returns, and more specifically on polygamy, would be different in a context with higher labor-market returns to education for women, but this is not something we can explore with the data we have.

³¹In Section 3.5 and Appendix Table C.6, we study the sensitivity of our results to controlling for the stock of schools when potential spouses were of school age.

Table 6: Labor-Market and Marriage-Market Returns to Education

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women						
	Wage Earner	Ever Married	Husband Education	Husband Polygamous	Husband No. of wives	Husband's Age ^(a)
<i>Reduced Form</i>						
No. public schools at age 7	0.0015*** (0.0005)	-0.0020 (0.0016)	0.0476*** (0.0114)	0.0054** (0.0024)	0.0178*** (0.0067)	-0.1430*** (0.0494)
<i>OLS</i>						
Years of schooling	0.014*** (0.001)	-0.019*** (0.001)	0.547*** (0.011)	-0.012*** (0.001)	-0.022*** (0.003)	-0.734*** (0.018)
<i>2SLS</i>						
Years of schooling	0.014*** (0.004)	-0.018 (0.014)	0.558*** (0.112)	0.066** (0.030)	0.221*** (0.083)	-1.613*** (0.581)
K-P F-stat	58.51	59.59	79.29	79.03	79.03	80.05
Mean dep. variable	0.004	0.941	0.950	0.472	1.846	12.234
Observations	693,045	698,736	491,152	490,045	490,045	491,806
No. clusters	9,205	9,214	9,040	9,039	9,039	9,042
Panel B: Men						
	Wage Earner	Ever Married	Wife(s) Education	Polygamous	No. of Wives	Wife(s) Age ^(a)
<i>Reduced Form</i>						
No. public schools at age 13	0.0043*** (0.0014)	0.0021 (0.0019)	0.0672*** (0.0089)	0.0050*** (0.0016)	0.0089*** (0.0029)	0.0175 (0.0264)
<i>OLS</i>						
Years of schooling	0.033*** (0.001)	0.013*** (0.000)	0.239*** (0.005)	0.003*** (0.001)	0.009*** (0.001)	-0.266*** (0.012)
<i>2SLS</i>						
Years of schooling	0.087*** (0.031)	0.040 (0.037)	1.339*** (0.246)	0.101*** (0.039)	0.181** (0.070)	0.344 (0.549)
K-P F-stat	26.56	28.43	26.24	25.71	25.71	25.85
Mean dep. variable	0.073	0.840	0.637	0.258	1.358	-9.740
Observations	602,210	604,670	438,328	470,247	470,247	439,070
No. clusters	9,202	9,201	9,041	9,091	9,091	9,043

Notes: Panel A sample: In columns (1) and (2), all nonmigrant women aged 25–60 in 1976 (the slight discrepancy in number of observations is explained by missing employment status); in columns (3)through(6), nonmigrant married women aged 25–60 in 1976 (the slight discrepancy in number of observations is explained by missing information on husband education and total number of spouses). Panel B sample: In columns (1) and (2), all nonmigrant men aged 25–60 in 1976 (the slight discrepancy in number of observations is explained by missing employment status); in columns (3)through (6), nonmigrant married men aged 25–60 in 1976. The discrepancy in number of observations is explained by missing information: for men, the number of wives is directly observed in the census while the exact characteristics of wives are obtained by matching married men and women living in the same household (see the Section 2). Wife education and wife age are averages for polygamous men. (a) The dependent variable is spouse age minus own age. Because of the birth-year fixed effects, it does not matter for the coefficient of interest, but the mean of the dependent variable is more informative, giving average age difference. All regressions control for the stock of private schools at age 7 (for women) or age 13 (for men), village fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant village controls interacted with a quartic cohort trend: precipitation, temperature, elevation, ruggedness, a malaria stability index, agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, Roome mission station, 1913 German mission school, and 1913 German government school (see Data Appendix). Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

the correlation between education and polygamy is negative. For men, the correlation is positive, but the magnitude much smaller than in the 2SLS (one year of education increases the number of wives by 0.01, Table 6, column 5). These differences are likely explained by the large role of unobservable characteristics in determining who becomes polygamous. One of these unobservable characteristics is religiosity. Christians are in principle not allowed to take several wives. While the census does not indicate what type of schooling each individual received (secular or religious), our instrument estimates the effect of public secular education only.³²

Could the positive effect of education on polygamy be explained by selection into marriage? (One could imagine for example that women who dislike polygamy decide to remain single when they receive education.) We saw above that education has no discernible effect on the likelihood of having ever been married. Appendix Table C.2 column (3) shows that education increases the probability of being *currently* married for women and men alike (though the coefficient is barely significant for women and not statistically significant for men). That is because education decreases the likelihood of being divorced or widowed. In columns (5) and (6) of Appendix Table C.2, we estimate the effect of education on polygamy on the full sample, including single, divorced, and widowed individuals who are not in a polygamous union and have zero co-wives (women) and zero wives (men). The results we obtain are very similar to the results of Table 6. Another concern related to sample selection is that we do not observe the number of co-wives of 8.3% of married women because, though we know they are married, we failed to match them with their husband.³³ It could be, for example, that educated women married to polygamists are more likely to live in a separate residence from the co-wives. Since we would not be able to observe the type of union for these women, this would lead us to underestimate the effect of education on the likelihood of a woman being in a polygamous union. However, in column (6) of Appendix Table C.2, we show that education does not significantly affect the probability of observing a married woman's

³²The compliers of a public school opening obtain secular education.

³³The census asked men about their number of wives, but women were only asked if they were married. To know the number of co-wives of women, we need to match them with their husband.

husband.³⁴

That (secular) education increases polygamy for men is not too surprising: if education increases the attractiveness of men on the marriage market (if only because of very large labor-market returns), then it should give men the opportunity to marry more women in a society that allows polygamy. The fact that education increases the likelihood of a woman being in a polygamous union is more surprising. If education makes women more attractive on the marriage market, and if women prefer to marry monogamous men, then we would expect education to decrease polygamy for women. But what we estimate with our instrumental strategy is a reduced-form result, which is the combined effect of matching on many characteristics. If educated women marry more-educated men and educated men are more likely to be polygamous, then our positive reduced-form effect might mask a negative affinity between female education and polygamy. To explore this possibility, in Section 4 we turn to a structural model of the marriage market. A structural model also has the advantage of accounting for matching on age.

As a first step toward understanding mechanisms, we ask whether more-educated women are more likely to enter polygamous unions as first wives or as wives of lower rank.³⁵ It is important because first wives in polygamous unions typically enjoy greater status and bargaining power (Matz, 2016).³⁶ In Table 7, we decompose the effect of education into different ranks within marriage: column (2) shows the effect of education on the probability of being the first wife in a polygamous union, column (3) on the probability of being the second wife, and column (4) on the probability of being the third wife or lower. An additional year of education increases the probability of being in a polygamous

³⁴Admittedly, it could still be the case that the effect of education on co-residence is different for women married monogamously and women married polygamously. If educated women married monogamously were *more* likely to live with their husband, and educated women married to a polygamist were *less* likely to live with their husband, this would downward-bias the effect of education on polygamy for women.

³⁵The census gives no information on the timing of marriages and therefore on the rank of women within a marriage. We infer the rank from age: a woman is given rank 1 if she is the oldest co-wife, rank 2 if she is the second oldest, etc. Matz (2016) finds that, in Ethiopia, age predicts marital sequence correctly for 83% of women.

³⁶Matz (2016) shows, in Ethiopia, that first wives tend to be from wealthier families and control more resources, and that their children enjoy more education. She also reviews previous works on different African countries, finding that first wives enjoy a higher social status and that their children fare better.

union as a first wife but has no effect on the probability of entering a polygamous union as a second or third wife.³⁷ We ignore whether educated women know that the man they marry as first wives will then take other wives, or whether they think they are entering a monogamous marriage (we do not have this information in our census data).³⁸ There is another concern related to the dynamic nature of polygamy: a woman in a monogamous union now could end up in a polygamous union in the future if her husband takes a second wife. Could it be the case that less-educated women are married with monogamous men who have a stronger likelihood of becoming polygamists in the future? We argue this is unlikely: monogamous men with a strong likelihood of becoming polygamous should be young and of higher socioeconomic status, but less-educated women marry older men of lower socioeconomic status (Table 6).

Our results differ from Fenske (2015), who finds no effect of recent education reforms on women’s polygamy in Sierra Leone, Nigeria, Kenya, or Zimbabwe.³⁹ While Fenske studies later education reforms, well after decolonization, we study a very early school expansion program during the colonial period. As we argue in the second part of the paper, the positive reduced-form effect of women’s education on polygamy comes from (1) matching on education and (2) the returns to education on the labor and marriage markets for men (educated women marry educated men who are richer and more likely to be polygamous). But while we find strong matching on education (that remains strong when we account for matching on age in the structural model), the natural experiments studied by Fenske “have found additional education affected women but have often failed

³⁷Interestingly, when we look at the reduced form, we see that the negative effect of the opening of a private, Christian school on the likelihood of being in a polygamous union comes entirely from the negative effect on the probability of entering as a second wife or lower. This result is consistent with women who received Christian education refusing to enter into a polygamous marriage, but still monogamously marrying men who then take other wives and become polygamous.

³⁸Since 1981, the marriage certificate must mention whether a marriage is monogamous or polygamous (“Ordonnance no 81-02 du 29 juin 1981”, article 49), but this disposition did not exist in late-colonial Cameroon. Our structural model is not dynamic, and therefore it assumes that women know whether the marriage they are entering into will be monogamous or polygamous. In practice, there are strong correlates of whether a man will become a polygamist, observable by women. For example, Boltz and Chort (2019) show that, in Senegal, norms regarding polygamy are transmitted from father to son.

³⁹Fenske’s results sometimes have large confidence intervals, but a meta-analysis (not reported) of the most precise estimate in each country finds a weighted average effect size of 0.7 percentage points with a standard error of 1.3 percentage points, compared with our effect size of 6.6 percentage points (standard error of 3.3 percentage points).

Table 7: Decomposing the Effect on Polygamy by Rank

	(1)	(2)	(3)	(4)
	In Polygamous Union (Any Rank)	In Polygamous Union First Wife	In Polygamous Union Second Wife	In Polygamous Union Third Wife or Lower
Reduced Form				
No. public schools at age 7	0.0054** (0.0024)	0.0057*** (0.0021)	-0.0003 (0.0013)	-0.0000 (0.0012)
No. private schools at age 7	-0.0082*** (0.0025)	0.0005 (0.0020)	-0.0030** (0.0015)	-0.0057*** (0.0018)
OLS				
Years of schooling	-0.012*** (0.001)	0.002*** (0.001)	-0.009*** (0.001)	-0.005*** (0.000)
2SLS				
Years of schooling	0.066** (0.030)	0.067*** (0.026)	-0.003 (0.016)	0.001 (0.014)
K-P F-stat	79.03	79.03	79.03	79.03
Mean dep. variable	0.472	0.244	0.163	0.064
Observations	490,045	490,045	490,045	490,045
No. clusters	9,039	9,039	9,039	9,039

Notes: Sample: All nonmigrant married women aged 25–60 in 1976. “In polygamous union” is a binary variable equal to 1 if a woman is married with a polygamous man and 0 if she is married with a monogamous man. “In polygamous union first wife” is a binary variable equal to 1 if a woman is married with a polygamous man as a first wife, and equal to 0 if she is married with a monogamous man or with a polygamous man as a second wife or lower. “In polygamous union second wife” is a binary variable equal to 1 if a woman is married with a polygamous man as a second wife, and equal to 0 if a woman is married with a monogamous man, with a polygamous man as a first wife, or with a polygamous man as a third wife or lower. “In polygamous union third wife or lower” is a binary variable equal to 1 if a woman is married with a polygamous man as a third wife or lower, and equal to 0 if a woman is married with a monogamous man, or with a polygamous man as a first or second wife. All regressions control for the stock of private schools at age 7 (for women) or age 13 (for men), village fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant village controls interacted with a quartic cohort trend: precipitation, temperature, elevation, ruggedness, a malaria stability index, agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, Roome mission station, 1913 German mission school, and 1913 German government school (see the Data Appendix). Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

to find that the characteristics of their husbands changed as a result” (Fenske, 2015, p. 71). Second, we find that in our setting, returns to education on the labor market for men are particularly strong (one year of education increases the likelihood of wage employment by 9 percentage points). The returns to attending one of Cameroon’s earlier public schools were likely particularly large because average education was still low. It is also possible that school quality was higher during the colonial period because resources were concentrated on a limited number of students (Pritchett, 2001).

Given that Christian missions in Cameroon were actively promoting monogamy, we might expect education (even if secular) to have different consequences in the marriage market for different religious groups. To explore this question, we divide the country between Christian-majority regions (about two-thirds of the population) and Muslim-majority regions using the 2005 census.⁴⁰ Interestingly, in 1976, the prevalence of polygamy was only slightly higher in Muslim-majority regions, with 50% of married women in a polygamous union versus 46% in Christian-majority regions. Appendix Table C.3 presents the first stage, the reduced form, and 2SLS for polygamy and number of wives or co-wives, separately for Christian-majority and Muslim-majority regions, and for men and women. For women, the first stage is much weaker in the Muslim North: one additional public school in the village at age 7 increases years of schooling by 0.03 years, versus 0.12 years in Christian-majority regions. This makes it difficult to compare the 2SLS results in the two samples: coefficients are higher in Muslim-majority regions, but they are not statistically different from zero, because of large standard errors. For men, the first stage is similar in the two regions (if anything, the opening of a public school has a larger effect in the North). As for the effect of education on polygamy, it appears to be smaller in the North (where the effects are not statistically different from zero), but large standard errors do not allow us to conclude that there is meaningful heterogeneity.

⁴⁰The 1976 census does not give the religion of respondents. The Muslim-majority *départements* in 2005 were Djerem, Faro et Deo, Mayo Banyo, Mbere, Vina, Diamare, Logone et Chari, Mayo Louti, and Noun. These correspond to the following *départements* in 1976: Adamaoua, Benoue, Diamare, Logone et Chari, and Bamoun. All other *départements* were Christian-majority.

3.5. Robustness

Migration. In our main specification, we instrument education by the stock of public schools at age a in the village where the individual lived in 1976. Because we know the district of birth of individuals, we assume that people who are still living in their district of birth are also still living in their childhood village, and we discard from the sample people who migrated out of their district of birth (about 30% of the sample). This creates two concerns (1) individuals who stayed in their district of birth might have moved within district⁴¹, and (2) if education affects the decision to migrate, then the sample of nonmigrants might be selected. More precisely, we worry that the women who decided to migrate when they received education had specific preferences regarding polygamy. If women with a strong distaste for polygamy migrated when they obtained education, then selective migration could explain why we estimate a positive effect of education on polygamy. In Appendix Tables C.4 (women) and C.5 (men), we present the results of a specification where we use as an instrument the district average number of available schools at a in the district of birth, which allows us to include migrants in the analysis. Our first stage thus becomes

$$E_{idc} = \alpha_d + \delta_c + \delta_c \times BR + \gamma \bar{N}_{dc}^{public,a} + x'_{dc} \phi + e_{idc} \quad (4)$$

α_d are district fixed effects. $\bar{N}_{dc}^{public,a}$ is the average at the district level of $N_{vc}^{public,a}$ on the sample of nonmigrants (a is 7 for women and 13 for men). The vector x_{dc} contains all variables in the vector x_{vc} averaged at the district level (there are 112 districts). Standard errors are clustered at the district level. Because the census gives the district of birth of everyone, migrants and nonmigrants, we can estimate this model on the sample of nonmigrants and on the full sample. The top panel of Appendix Table C.4 shows the first stage of the district-level instrument for women, on the sample of nonmigrants (columns 1 to 3) and on the full sample (columns 4 to 7). Compared to our main specification, we lose some statistical power, with an excluded instrument F-stat in the 20–30 range versus

⁴¹Though, because we are considering all schools in a radius of 10 km around the village, short distance migration should not be an issue.

60–80 in Table 6, Panel A.

In the bottom panels of Appendix Table C.4, columns (1) to (3), we display the results of the reduced form and the 2SLS on the sample of nonmigrants for the probability of being a wage earner, the probability of having a polygamous husband (conditional on being married), and the number of co-wives. We broadly replicate our main results, which reassures us that within-district migration is not a concern: one additional year of education increases the likelihood of being a wage earner by 1.3 percentage points, increases the likelihood of having a polygamous husband by 4 percentage points, and increases the number of co-wives by 0.2. In columns (4) to (7), we display results estimated on the full sample, including out-of-district migrants. The first thing to note is that an additional year of education increases the likelihood of migrating out of the district by 4 percentage points (column 4). This is not large enough for selection to fully explain the positive effect of education on polygamy: indeed, even in the worst-case scenario where no migrant woman was marrying a polygamous man, selective migration would increase rates of polygamy by a maximum of 1.8 percentage points per year of education.⁴² In columns (5) to (7), we estimate returns to education on the full sample, including migrants. If selective migration alone was explaining the positive effect of education on polygamy, we should not estimate any effect of education on polygamy on this sample. However, we find that one additional year of schooling increases the likelihood of being in a polygamous union by 1.6 percentage points (though the effect is not statistically different from zero), and increases the number of co-wives of married women by 0.14 (statistically significant at the 5% level), against 0.2 on the sample of nonmigrants. The positive effect of education on women’s polygamy is not an illusion created by selective migration.

In Appendix Table C.5, we present the district-level instrument results for men. Unfortunately, as far as men are concerned, the first stage is very weak with the district

⁴²Let us assume that the true percentage of women in a polygamous marriage in a village is 44% (it corresponds to the percentage of married women in a polygamous marriage in the population, including migrant women). If we fail to observe 4% of women in the village because they migrated, and if these women are all monogamous, then the share of married women in a polygamous marriage jumps to 45.8% (44 of 96).

instrument, especially when considering the full sample (with excluded instrument F-stats in the 1–4 range). This might be explained by errors in the district of birth for outmigrant men. This is one of the reasons why we do not use the district-level instrument as our main specification in this paper, because the estimation of our structural model of matching requires a strong instrument for the education of women and for the education of men. However, we note that the point estimates of the 2SLS are hardly affected by the inclusion of migrants.

Direct effect of school construction on the spouse. One concern about our identification strategy is that, in a village, potential husbands and wives would be subject to a similar education shock. However, this is not the case in our context: in Cameroon in 1976 wives were on average 10 years younger than their husband. Still, there is a large correlation between the stock of schools in the same village 10 years apart. It is particularly concerning when estimating assortative matching on education. Do individuals who have access to a school marry more-educated individuals because of assortative matching on education, or do they marry more-educated individuals because of assortative matching on age (they marry individuals of a similar age who also had access to a school)? Disentangling the matching on age from the matching on education, and more generally disentangling the different dimensions of matching, is precisely what our structural model (in Section 4) allows us to do. Here, we study the sensitivity of our estimates to controlling for the stock of schools when potential spouses were of school age. In the top panel of Appendix Table C.6, we show returns to education for women (instrumenting education by the stock of schools at age 7) with and without controlling for the stock of schools in the village at age 3, when the cohort of men 10 years older was of school entry age (13). The effect of an additional year of education is barely affected by controlling for the number of schools at age 3, except for husband’s age (column 6).⁴³ In the bottom panel of Appendix Table C.6, we show returns to education for men (instrumenting education by the stock of schools at age 13) with and without controlling for the stock of schools

⁴³The number of public schools at age 3 is positively correlated with husband’s age, which increases the negative effect of female education on husband’s age. This could be the sign of a trade-off between husband’s age and husband education.

in the village at age 17, when the cohort of women ten years younger was of school entry age (7). In this case, the estimated effect of an additional year of education tends to fall for all variables. Note, however, that the stocks of schools at ages 13 and 17 in the same village are highly correlated (with a correlation coefficient of 0.92 in our sample).⁴⁴

Robustness to heterogeneous treatment effects. A recent literature shows that two-way fixed-effects models can produce biased estimates in the presence of heterogeneous effects (de Chaisemartin and D’Haultfoeuille, 2020; Goodman-Bacon, 2021). In the Appendix, we show that our results are robust to using the DID_M estimator of de Chaisemartin and D’Haultfoeuille (2020), which solves the issue. This estimator compares the average change in outcomes between two consecutive cohorts in the same village exposed to a different number of schools (say the number of schools in the village went from N to $N + 1$) to the average change in outcomes between these two consecutive cohorts in villages where the number of schools remained N . We can also compute the effect of the change in school stock many cohorts after the event, using villages where the stock of schools stayed the same. Likewise, we can compute the effect of the change in school stock for the older cohorts — a placebo check of the absence of a pretrend. To implement the DID_M estimator, we group individuals into five-year cohorts centered around ages ending in 0 and 5. There are two reasons for doing this. First, we would otherwise be comparing the outcomes of individuals aged exactly 7 at the time of a school construction to the outcomes of individuals aged exactly 8, who are partially treated. Instead, we compare the outcome of individuals aged 5–9 to the outcome of individuals aged 10–14 (they might still be partially treated, but the problem is tempered). Second, it improves the balancedness of our panel and increases the number of consecutive cohorts contributing to the estimation.⁴⁵ We redefine the time and treatment variables in the panel as follows: each individual is attributed the central birth date of their cohort and the corresponding number of schools at age 7.⁴⁶ We then implement the estimator of de Chaisemartin and

⁴⁴For women, the stocks of school at ages 7 and 3 in the same village are less correlated with a coefficient of 0.84.

⁴⁵Not all birth years are systematically represented within each village.

⁴⁶For example, someone born in 1940 in a village where one school opened in 1948 belongs to the 1939–1943 cohort (33 to 37 years old in 1976), whose central date is 1941 and is therefore attributed a

D’Haultfoeuille (2020). Appendix Figures C.4 and C.5 display placebo and dynamic effects for the first stage and reduced form.⁴⁷ Appendix Table C.8 displays the average of the dynamics effects for the five cohorts after the school opening on education (columns 1–2 for women and 5–6 for men) and on the number of (co-)spouses (column 3 for women and column 7 for men). The DID_M estimator cannot easily be used in an instrumental variable estimation, but we display the ratio of the reduced form to the first stage in columns 4 and 8. These ratios are remarkably similar to our main estimates: one year of schooling increases a woman’s number of co-spouses by 0.20 (against 0.22 in our main estimation Table 6) and a man’s number of wives by 0.17 (against 0.18 in our main estimation Table 6). It is therefore unlikely that our two-way fixed-effects results are biased by heterogeneous effects.

Selective mortality. To show that our results are not explained by selective mortality, Table C.9 displays, along with our baseline results, results estimated on the restricted sample of men and women aged 25–45. Results obtained on the young sample are very similar to results obtained on the full sample. One noticeable difference is that the effect of education on the probability of being a wage earner is much lower in the young sample (2.5 percentage points instead of 8.7). This might be because returns to schooling were particularly high in the older cohorts, which were among the first to receive formal education in Cameroon.

Sensitivity to violation of the exclusion restriction. In Appendix Table C.10, we investigate how a violation of the exclusion restriction would affect our estimates of the returns to education on the labor and marriage markets. Adopting the “plausibly exogenous” approach of Conley et al. (2012), we allow the stock of schools at age a ,

number of schools at age 7 equal to 1. Appendix Table C.7 replicates our main two-way fixed-effects results with the redefined cohort and treatment variables: results are qualitatively similar but noisier and somewhat smaller, which is expected since the redefinition likely increases measurement error.

⁴⁷We show four placebo effects and five dynamic effects, which corresponds roughly to the age window we consider in the event-study Figure 2.

$N_{vc}^{public,a}$, to enter directly into equation (3) with a coefficient γ :

$$y_{ivc} = \alpha_v + \delta_c + \delta_c \times BR + \tau E_{ivc} + x'_{vc}\phi + \gamma N_{vc}^{public,a} + \nu_{ivc} \quad (5)$$

This means we allow the exclusion restriction to be violated, either because school construction affect labor and marriage market outcomes through a channel other than education, or because school construction is not perfectly exogenous and is correlated with an unobserved variable that matters for labor and marriage market outcomes. In Conley et al. (2012), priors on the distribution of γ are used to build a confidence interval encompassing the possibility of a violation of the exclusion restriction. We follow common practice and report the largest value of γ that can be assumed while maintaining the effect of education τ statistically different from zero (Appendix Table C.10). γ is reported as a percentage of the effect of the instrument in the reduced form. Appendix Table C.10 shows that large violations of the exclusion restriction are required to make the effect of education indistinguishable from zero. Consider for example the number of co-wives of a woman as the dependent variable (column 2). Even if the value of γ was as high as one-third of the effect of the instrument in the reduced form, the effect of education on the number of co-wives would still be significant at the 5% level.

Alternative instrument. Appendix Table C.11 displays, along with our baseline results, results obtained using the same set of instruments for men and women: the stock of schools in the village at age 7, and the number of schools opening between ages 8 and 13. The results are overall very similar, but the excluded instrument F-stat is lower when using the same instruments for men and women.

4. Structural Estimation of Affinities

We showed that women who achieved more years of education because of an exogenous increase in school supply in their village were more likely to be married to a polygamist as a first wife. However, this reduced-form approach tells us little about preferences on the

marriage market. It is hard to believe that more-educated women prefer polygamous men, all else being equal. However, polygamous men are more likely to come from better-off families, have higher wealth, enjoy better employment prospects, and have achieved more years of education. It is not hard to believe that these characteristics are highly valued by women, and by their families, especially in a context where, as we have seen, there were not many labor-market opportunities for educated women. Assortative matching on education, in particular, is likely to play an important role: if educated women marry more-educated men, but more-educated men are also more likely to be polygamous, then more-educated women might end up more likely to marry polygamists. How can we provide evidence for this mechanism using data on marriages and our education supply shock? The naive and straightforward approach is not warranted here: we cannot control for husband’s education in equation (3), even if we instrument husband’s education using the stock of schools when he was of school age. Because any characteristic of the husband is a choice of the wife and her family, it is almost surely correlated with the error term in equation (3), even if education is exogenously received by the man.

In this section, to show the role played by matching on education in explaining the reduced-form results, we estimate a structural model of marriage with utility transfers. We adapt the model of Choo and Siow (2006) to polygamous unions by allowing men to marry several women. Using data on who marries whom, we cannot recover preferences directly, but Dupuy and Galichon (2014) show that it is possible to estimate the matrix of second derivatives of the joint utility of a match with respect to male and female attributes. These parameters, called “affinities,” describe the extent to which different characteristics of men and women attract or repulse one another on the marriage market. The affinity between two given attributes is net of the affinities between all other attributes included in the model. For example, the education affinity is net of the age affinity — this is important because, in a context of increasing education, people of similar ages tend to have similar levels of education. We are particularly interested in the affinity between the number of wives of a husband and the education of a wife: a positive affinity would indicate that polygamous men and educated women attract one another

on the marriage market (what our reduced-form results appear to imply). A negative affinity would indicate that they repulse one another, all other husband attributes being equal. This could mean that educated women have a dislike for marrying polygamous men, or that polygamous men have a dislike for marrying educated women.⁴⁸ We are more precisely interested in how the affinity between the number of wives of a husband and the education of a wife changes when we allow matching on education, that is, when we allow a direct affinity between the education of husband and wife. If matching on education plays a large role in explaining the reduced-form result, then the affinity between husband’s polygamy and wife’s education should decrease when we allow a direct affinity between education and education.

4.1. A Matching Model of the Marriage Market

In the standard structural model of matching with transferable utility of Choo and Siow (2006) or Dupuy and Galichon (2014), a man with a set of attributes x marries a woman with a set of attributes y .⁴⁹ We adapt the model of Choo and Siow (2006) to the polygamous case, where a man can marry several women.

In our model, women and men choose partners to maximize their utility. For the sake of tractability, we omit from our model some important aspects of polygamous marriages. First, though polygamy is a dynamic process (a man marries a first wife, then a second wife later in his life, and so on), our model is static. To account for the important difference — for women — between marrying a single man and marrying an already-married man, we incorporate in the utility function the rank in the marriage. Second, in the context of Cameroon before 1976, it might be more realistic to envision marriage as a match between two families rather than between two individuals. One can understand the utility functions in our model as applying to the whole family of the man or woman — but we do not explicitly model the strategic interactions within the bride’s or groom’s families.⁵⁰

⁴⁸In the framework of our model, we cannot distinguish between these two interpretations.

⁴⁹In Choo and Siow (2006), the attributes of men and women are discrete. Dupuy and Galichon (2014) show a generalization to continuous attributes in the monogamous case.

⁵⁰This means we apply a unitary model of household decisions to both families. This has well-known

A man m of attributes $x_m = x$ marrying an ordered set of women $W = (w_1, \dots, w_{n_Y})$ of attributes $Y = (y_1, \dots, y_{n_Y})$, where the order gives the rank of wives. He gets the following utility⁵¹

$$\mathcal{U}(m, W) = U(x, Y) + \varepsilon_{mY} = \sum_{r=1}^{n_Y} u(x, y, n_Y, r) + \varepsilon_{mY} \quad (6)$$

$U(x, Y)$ is the systematic part of the utility; it is assumed to be additively separable in the utility given by each match with a woman of attributes y . The utility of a match with a given wife is also allowed to depend on the number of wives n_Y and on her rank r — for example a man might care more about the education of his first wife than subsequent wives (see Matz, 2016, for a framework in which highly productive women have a greater chance of becoming first wives). The utility of a given match does not depend on the characteristics of the other wives. This rules out potential complementarity/substitution between different types of wives. This restrictive assumption keeps the model tractable and simplifies the interpretation of results: we keep our focus on the total number of wives and the characteristics of each wife, disregarding how the interaction between the characteristics of different wives might matter for a man’s choice. We normalize the utility of single men to $\mathcal{U}(m, \emptyset) = 0 + \varepsilon_{m\emptyset}$.

The model has transferable utility: husband and wife can transfer utility to one another to compensate for some attributes. These utility transfers play the role of prices, but they are unobserved by the econometrician. In Cameroon, marriage is often associated with an actual price, the bride price, which is transferred from the groom’s family to the bride’s family.⁵² However, we do not observe the bride price and therefore model it as an unobserved transfer. $U(x, Y)$ is the utility post transfers. ε_{mY} is a randomly drawn “sympathy shock” for each set of spouses Y . Each man m is therefore allowed to have idiosyncratic preferences for his wives’ attributes Y , but not preferences for individual women. ε_{mY} follows a Gumbel distribution, and is independent between Y .⁵³

limitations, but we have little choice here as our data does not provide information on bargaining within families.

⁵¹ n_Y is the total number of wives; in a monogamous marriage, $n_Y = 1$. We note $n_Y = 0$ for single men.

⁵²See footnote 4.

⁵³This assumption is restrictive but very common for discrete choice modeling (e.g. multinomial logit).

A woman w with attributes $y_w = y$ becomes the wife of rank r of a man m of attributes $x_m = x$ and $n_Y - 1$ other wives (in monogamous marriages, $n_Y - 1 = 0$ and $r = 1$). She gets the following utility:

$$\mathcal{V}(m, w, n_Y, r) = v(x, y, n_Y, r) + \eta_{wxn_Yr} \quad (7)$$

$v(x, n_Y, y, r)$ is the systematic part of the utility; it depends on the attributes of the man x , of the woman y , on the number co-wives n_Y , and on the rank r . η_{wxn_Yr} is a randomly drawn “sympathy shock” of woman w for becoming the wife of rank r of a man of type x with n_Y spouses. We normalize the utility of single women to $\mathcal{V}(\emptyset, w) = 0 + \eta_{w\emptyset}$. $v(x, n_Y, y, r)$ depends on the number of wives and on the rank, but not on the other characteristics of co-wives. Recent work has studied empirically the strategic interactions between co-wives in polygamous marriages (Barr et al., 2019; Rossi, 2019), and theoretically the complementarities with a quality dimension for the spouses (Reynoso, 2019). However, as in the case of men’s utilities $u(x, y, n_Y, r)$, this assumption keeps the model tractable and allows us to focus on the main mechanism of interest. η_{wxn_Yr} follows a Gumbel distribution and is independent between (x, n_Y, r) . Each woman w is therefore allowed to have idiosyncratic preferences for the husband’s attributes x , the number of co-wives n_Y and the rank r , but not allowed to have preferences for individual men.

Because the sympathy shocks ε_{mY} and η_{wxn_Yr} are i.i.d. and follow a Gumbel distribution, the distribution of men’s wives $\pi(Y|x)$ and the distribution of women’s husbands $\pi(x, n_Y, r|y)$ both follow a multinomial logit at the equilibrium (McFadden, 1974):

$$\pi(Y|x) = \frac{\exp(U(x, Y))}{\sum_{Y' \in \mathcal{Y}} \exp(U(x, Y'))} \quad (8)$$

$$\pi(x, n_Y, r|y) = \frac{\exp(v(x, y, n_Y, r))}{\sum_{x' \in \mathcal{X}, n'_Y \in \mathbb{N}, r' \leq n'_Y} \exp(v(x', y, n'_Y, r'))} \quad (9)$$

Like all multinomial logit frameworks, the model imposes Independence of Irrelevant Alternatives (IIA) — the choice between two partners is not affected by the availability of other potential partners.

For an equilibrium to exist, these two equations need to be compatible, that is, for every (x, y, n, r) , the number of men of type x who choose to marry n spouses with a spouse of type y in rank r needs to equal the number of women willing to accept this situation. In other words, calling f and g the distribution of men's and women's types, for every (x, y, n, r) :

$$f(x) \sum_{Y|n_Y=n, Y_r=y} \pi(Y|x) = g(y)\pi(x, n_Y = n, r|y) \quad (10)$$

where Y_r denotes the wife of rank r in marriage Y .

Many-to-one matching models tend to be unstable and have no equilibrium.⁵⁴ In our case, we show that there is a unique equilibrium (Theorem 1). For transparency, we underline three features of the model — already present in the monogamous model of Choo and Siow (2006) — contributing to the existence of a unique equilibrium: (1) the presence of transferable utilities, (2) the assumption of large markets where many potential spouses with identical attributes are perfect substitutes,⁵⁵ and (3) the absence of friction in matching. Admittedly, the last two assumptions are somewhat more demanding for polygamous marriages than for monogamous ones.

Theorem 1. *For any utility functions following (6) and (7), and for any distribution of female and male attributes, there is a unique equilibrium following (8), (9), and (10). This equilibrium maximizes the total surplus and is characterized by the matching function*

$$\pi(x, y, n_Y, r) = \exp\left(\frac{\Phi(x, y, n_Y, r) - b(y) - a(x, n_Y, r)}{2}\right) \quad (11)$$

where $\Phi(x, y, n_Y, r) = u(x, y, n_Y, r) + v(x, y, n_Y, r)$ is the total systematic utility generated by a match (x, y, n_Y, r) .

The proof of Theorem 1 is given in Appendix D.

Observing the distribution of marriages allows identification of the joint utility Φ up to

⁵⁴See Hatfield et al. (2013) for a theoretical example.

⁵⁵This creates an equilibrium where marginal individuals are (nearly) indifferent between several types of marriages. In a very different, much more general context, Azevedo and Hatfield (2018) show the existence of an equilibrium in large markets modeled by a continuum of agents.

two separately additive functions, $b(y)$ and $a(x, n_Y, r)$. We can identify here the second derivatives of Φ with respect to the attributes of men and women. Because it appears in the function $a(x, n_Y, r)$ but not in the function $b(y)$, the number of co-wives and the rank can be considered as husband attributes.

In the monogamous case, Dupuy and Galichon (2014) adopt the simple parametrization $\Phi(x, y) = x' Ay$, where A is a $d_x \times d_y$ matrix (d_x and d_y are the number of attributes of x and y). A is the Hessian of Φ : $\Phi: \frac{\partial^2 \Phi}{\partial \mathbf{x} \partial \mathbf{y}} = A$. We follow this simple parametrization, but we adapt it to the polygamous case by considering a longer vector of men's characteristics \mathbf{x} , containing the vector x , as well as information on polygamy. In a first specification, we ignore the rank and consider $\mathbf{x} = (x, n_Y)$. In a second specification, we fully interact the rank with the other husband characteristics: we consider $\mathbf{x} = (x \times \mathbf{1}_{r=1}, n_Y \times \mathbf{1}_{r=1}, x \times \mathbf{1}_{r>1}, n_Y \times \mathbf{1}_{r>1})$. This means we estimate affinities between the attributes y of a wife and the attributes (x, n_Y) of a husband separately for matches where the wife is of rank 1 and for matches where the wife is of rank 2 or higher. We opt for this parametrization in light of our finding that the effect of education on the likelihood that women are in a polygamous union differs by rank (Table 7), and in light of the literature on assortative mating in polygamous marriages: Matz (2016) finds, in Ethiopia, differences between the average characteristics of first wives and higher-order wives; Reynoso (2019) explicitly models assortative mating on education for first wives, but not for second wives.

Having modeled $\Phi(x, y, n_Y, r) = \mathbf{x}' Ay$, we can identify the elements $a_{\mathbf{x}, y}$ of the affinity matrix A using observations on the distribution of marriages. Let us consider as an example $a_{x_E \times \mathbf{1}_{r=1}, y_E}$ — the affinity between the education of a husband and the education of his first wife. It is the second derivative of the joint utility of the match with respect to education of the husband and education of the wife. If we assume that education of the wife increases the joint utility of the match, then $a_{x_E \times \mathbf{1}_{r=1}, y_E} > 0$ means that the increase in utility brought by the education of the first wife is higher in marriages where the husband is educated. Because individuals are maximizing their utility, it also means that we are more likely to observe a match between an educated man and an educated woman as a first wife than a match between an educated man and an uneducated woman

as a first wife.⁵⁶ Each affinity takes as given the affinity between all other attributes of the match. However, it should be noted that estimating the matrix A does not allow to establish the overall impact of a change in the number of educated men and women on polygamy, or aggregate effects in general.⁵⁷

4.2. Logit Estimation of the Model on Pairs of Couples

We propose a new way of estimating this matching model of the marriage market, following Charbonneau’s (2014) approach to estimating logit models with two dimensions of fixed effects.

We consider a pair of couples, Mr. 1 and Mrs. 1, and Mr. 2 and Mrs. 2 and we show that the probability that Mr. 1 is with Mrs. 1 and Mr. 2 with Mrs. 2 rather than the opposite can be written in a logistic form.

Theorem 2. *Let’s consider a pair of couples, two women, $w = 1$ and $w = 2$, whose respective husbands, $h(1) = h_1$ and $h(2) = h_2$, are (respectively or not) $m = 1$ and $m = 2$. if we assume that the fact that one particular man is already married does not affect the overall probability of a woman getting married, the probability that the couples are (1,1) and (2,2) rather than the opposite is written as follows:*

$$P(h_1 = 1 | \{h_1, h_2\} = \{1, 2\}) = \frac{\exp(\Phi_{11} + \Phi_{22} - \Phi_{12} - \Phi_{21})}{1 + \exp(\Phi_{11} + \Phi_{22} - \Phi_{12} - \Phi_{21})} \quad (12)$$

where Φ_{mw} is the total systematic utility generated by a match between m and w .

The proof of Theorem 2 is given in Appendix D. Equation (12) is easy to interpret: Mrs. 1 is more likely to be married with Mr. 1 than with Mr. 2 when $\Phi_{11} + \Phi_{22} > \Phi_{12} + \Phi_{21}$; that is, the sum of the systematic utilities of matches is higher when woman 1 is married with man 1 and woman 2 with man 2. Equation (12) is defined only for married individuals,

⁵⁶If we assume that education of the wife decreases the joint utility, then a positive affinity means that it decreases utility more in marriage where the husband is uneducated. Our model does not allow us to identify the attractiveness of each type of individual on the marriage market.

⁵⁷This is because in equation (11), $b(y)$ and $a(x, n_Y, r)$ depend on the distribution of characteristics of men and women, including education. Estimating a and b would be computationally very demanding, because there is no closed-form solution.

which means that single people do not contribute to the estimation. This comes from the IIA assumption imposed by the multinomial logit framework: the attributes of unmarried people are irrelevant for estimating the trade-offs faced by married people.

We adopt parametrization $\Phi(\mathbf{x}, y) = \mathbf{x}'Ay$ and apply it to equation (12), which gives, after simplification:

$$P(h_1 = 1 | \{h_1, h_2\} = \{1, 2\}) = \frac{\exp((\mathbf{x}_1 - \mathbf{x}_2)'A(y_1 - y_2))}{1 + \exp((\mathbf{x}_1 - \mathbf{x}_2)'A(y_1 - y_2))} \quad (13)$$

To estimate the affinity matrix A , we compute the sum of the log-likelihoods defined by (13) over a sample of potential pairs of couples — hence we estimate the affinity matrix by maximum of pseudo-likelihood.⁵⁸ Our dataset contains about 500,000 couples: considering every possible pair of couples would mean considering more than 10^{11} potential pairs, which is not feasible. In each village, we randomly divide all couples into clusters of roughly 20, and we consider every possible pair of couples within each cluster (190 pairs of couples per cluster).⁵⁹ This procedure also eliminates village fixed effects from the equations.

The logit in equation (13) has no constant: when $\mathbf{x}_1 = \mathbf{x}_2$ or $y_1 = y_2$, $(h_1 = 1, h_2 = 2)$ is as likely as $(h_1 = 2, h_2 = 1)$ (for the econometrician). The dependent variable of the logit is always 1, as Mr. 1 is always the husband of Mrs. 1. So, the model is identified when the matching is imperfect. For example, assume that education is the only dimension of \mathbf{x} and y . If the assortative matching on education was perfect (the more-educated man is always with the more-educated woman), then $\mathbf{x}_1 - \mathbf{x}_2$ and $y_1 - y_2$ would always share the same sign, and increasing A would always increase the likelihood of the model. However, if the matching is imperfect, there are some couples for which $\mathbf{x}_1 - \mathbf{x}_2$ and $y_1 - y_2$ have different signs, so that increasing A does not systematically increase the likelihood.

⁵⁸We consider only women with different husbands; that is, we never consider two couples such that one man is married to both women.

⁵⁹The size of the cluster is not always exactly 20 because the number of couples per village is not always a multiple of 20. Our results are robust to considering smaller clusters of 5 or 10 couples (not reported).

4.3. Endogeneity of Education

Our structural model allows us to estimate the “affinity” between a certain number of attributes of wives and husbands. These attributes are what we can observe in the census: age, whether the husband is a polygamist, the rank of the wife in the marriage, and education. We are interested in the exogenous part of education, the one we instrument using the stock of public schools in the village when the individual was of school age. We are also motivated to use public schools as an instrument because we are interested in education received in public schools rather than in private Christian schools, and the census does not say where individuals received their education.

To account for the endogeneity of education, we use a two-step control-function approach. Rivers and Vuong (1988) prove the consistency of control-function approaches in the probit case. We are not aware of any paper focusing on logit endogenous variables, but Wooldridge (2015) discusses control-function approaches in econometrics along a large class of models, including nonlinear dependent variables in general. In this case, control-function approaches require independence between the distribution of the error terms and of the instruments.

In a first step, we estimate:

$$\begin{aligned} \Delta E_{w_{jv}} = & \Delta Q(A_{w_{jv}}) + \gamma_w \Delta N_{w_{jv}}^{public,7} \\ & + \Delta Q(A_{w_{jv}}) \times x'_v \theta_w + \phi_w \Delta N_{w_{jv}}^{private,7} + \Delta e_{w_{jv}} \end{aligned} \quad (14)$$

$\Delta E_{w_{jv}} = E_{1jv} - E_{2jv}$ is the difference in women’s education for couple pair j in village v , $\Delta Q(A_{w_{jv}})$ is the difference in a quartic polynomial in age,⁶⁰ and $\Delta N_{w_{jv}}^{public,7}$ is the difference in the number of public schools in the village at age 7 (and $\Delta N_{w_{jv}}^{private,7}$ the difference in the number of private schools). We also interact the quartic in age difference with a vector x_v of village attributes.⁶¹ This equation is intended to be as close as possible to the first stage in equation (2), with a couple of caveats: there are no village fixed effects,

⁶⁰ $q_1(A_{1jv} - A_{2jv}) + q_2(A_{1jv}^2 - A_{2jv}^2) + q_3(A_{1jv}^3 - A_{2jv}^3) + q_4(A_{1jv}^4 - A_{2jv}^4)$

⁶¹As in equation (2), a dummy for belonging to British Cameroon, precipitation, temperature, elevation, ruggedness and agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, Roome mission station, 1913 German mission school, and 1913 German government school.

because we consider only pairs of couples within the same village. There are no cohort fixed effects either. Because we consider the interactions between all attributes of the wives and husbands, adding cohort fixed effects would require estimating thousands of additional coefficients (we consider 35 different cohorts). For this reason, cohort fixed effects are replaced by a quartic function of age.⁶²

We estimate a similar first-step equation for men, as follows:

$$\begin{aligned}\Delta E_{m_jv} = & \Delta Q(A_{m_jv}) + \gamma_m \Delta N_{m_jv}^{public,13} + \lambda \Delta P_{m_jv} \\ & + \Delta Q(A_{m_jv}) \times x'_v \theta_m + \phi_m \Delta N_{m_jv}^{private,13} + \Delta e_{m_jv}\end{aligned}\quad (15)$$

$\Delta P_{m_jv} = P_{1jv} - P_{2jv}$ is the difference⁶³ between the number of wives of husband 1 and husband 2. In the specification where we interact husband's characteristics with the rank of his wife, we estimate two separate first stage equations:⁶⁴

$$\begin{aligned}\Delta(\mathbf{1}_{m,r=1} E_{m_jv}) = & \Delta Q(\mathbf{1}_{m,r=1} A_{m_jv}) + \gamma_{m1} \Delta(\mathbf{1}_{m,r=1} N_{m_jv}^{public,13}) + \lambda_1 \Delta(\mathbf{1}_{m,r=1} P_{m_jv}) \\ & + \Delta Q(\mathbf{1}_{m,r=1} A_{m_jv}) \times x'_v \theta_{m1} + \phi_{m1} \Delta(\mathbf{1}_{m,r=1} N_{m_jv}^{private,13}) + \Delta(\mathbf{1}_{m,r=1} e_{m_jv})\end{aligned}\quad (16)$$

$$\begin{aligned}\Delta(\mathbf{1}_{m,r>1} E_{m_jv}) = & \Delta Q(\mathbf{1}_{m,r>1} A_{m_jv}) + \gamma_{m2} \Delta(\mathbf{1}_{m,r>1} N_{m_jv}^{public,13}) + \lambda_2 \Delta(\mathbf{1}_{m,r>1} P_{m_jv}) \\ & + \Delta Q(\mathbf{1}_{m,r>1} A_{m_jv}) \times x'_v \theta_{m2} + \phi_{m2} \Delta(\mathbf{1}_{m,r>1} N_{m_jv}^{private,13}) + \Delta(\mathbf{1}_{m,r>1} e_{m_jv})\end{aligned}\quad (17)$$

$\mathbf{1}_{m,r=1}$ is a binary equal to one if husband m is the husband of a first wife and $\mathbf{1}_{m,r>1}$ is a binary equal to one if husband m is the husband of a higher-order wife. Note that a polygamous man will be the husband of a first wife in some pairs of couples and the husband of a higher-order wife in others.

In a second step, when estimating equation (13), we consider — instead of education of women — the predicted education and residual from equation (14), $\Delta \hat{E}_w$ and $\Delta \hat{e}_w$, and

⁶²Results are robust to including a polynomial of degree 3 or 5 in age (not reported).

⁶³Or, in another specification, the difference between a dummy equal to 1 if husband 1 is a polygamist and a dummy equal to 1 if husband 2 is a polygamist.

⁶⁴The estimation of the first equation (15) gives $\widehat{\Delta e_{m_jv}}$. This does not predict the residuals of (16) and (17) as long as $\widehat{\Delta e_{m_jv} \Delta r_{m_jv}}$ cannot predict $\Delta e_{m_jv} r_{m_jv}$: $\Delta e_m \Delta r_m \neq \Delta(e_m r_m)$.

— instead of education of men — the predicted education and residual from equation (15), or, when interacting with the rank, from equation (16) (if the wife is a first wife) or (17) (if the wife is a higher-order wife). We could also have actual education and the residuals. This does not affect the affinity between education and the other attributes, but it makes the affinity between wife’s and husband’s education and the residuals easier to interpret.

Finally, the precision of the estimated matrix A must account for the two-stage procedure. The precision of A is estimated using a sandwich-like estimator with the Hessian of the joint log-likelihood

$$l_j = \log \left\{ \Lambda[\Delta \mathbf{x}'_j A \Delta y_j] \varphi(\hat{e}_{mj}^2 / \hat{\sigma}_m^2) \varphi(\hat{e}_{wj}^2 / \hat{\sigma}_w^2) \right\}$$

where Λ is the cumulative of the logistic distribution, φ is the density of the normal distribution, and $\hat{\sigma}_m^2$, and $\hat{\sigma}_w^2$ are the estimated variances of \hat{e}_m^2 , and \hat{e}_w^2 . In the specification where the husband’s attributes are fully interacted with rank, the joint log-likelihood writes:

$$l_j = \log \left\{ \Lambda[\Delta \mathbf{x}'_j A \Delta y_j] \varphi(\widehat{\mathbf{1}_{m,r=1} e_{mj}^2} / \hat{\sigma}_{m,r=1}^2) \varphi(\widehat{\mathbf{1}_{m,r>1} e_{mj}^2} / \hat{\sigma}_{m,r>1}^2) \varphi(\hat{e}_{wj}^2 / \hat{\sigma}_w^2) \right\}$$

Because we consider a dyadic dataset of possible pairs of couples, a given couple appears several times in the dataset and the pseudo-likelihood function, and this needs to be accounted for in the estimation of standard errors — this is discussed in Jochmans (2017). This does not really matter in our case because we opt to cluster standard errors by a larger unit, the village.

4.4. Results: Affinity Matrices

In this section, we estimate matrices A of affinity parameters between attributes of husbands and wives. These matrices are d_x by d_y , where d_x is the number of male attributes considered, and d_y the number of female attributes considered. Element a_{kl} of A is the

affinity between husband’s attribute \mathbf{x}_k and wife’s attribute y_ℓ . It is the second derivative of the joint utility of a match with respect to \mathbf{x}_k and y_ℓ .

For now, we set aside the potential differences in affinities by rank in the marriage. In a first step, we estimate the affinity between a woman’s education and the number of wives of her husband without accounting for the matching on education (nor the rank in the marriage). In this way, we replicate in the structural framework the central result of the quasi difference-in-differences estimation (Table 6): educated women are more likely to mate with polygamous men. On a sample of couple pairs indexed by j , we estimate the following equation:

$$P_j = \Lambda \left[\Delta \mathbf{x}'_j A \Delta y_j \right]$$

where the vector $\Delta \mathbf{x}_j$ contains the difference between the number of wives of husband 1 and husband 2 ($\Delta P_{m_{jv}} = P_{1jv} - P_{2jv}$), as well as a quartic polynomial in the age difference. Δy_j is a vector containing all the Δ variables in the first-step equation (14), as well as the residual $\Delta \hat{e}_{wj}$ (the control function). The main difference with the nonstructural approach is that we are able to flexibly account for the matching on age by including a quartic polynomial in age for both husbands and wives. The wife’s vector contains 58 attributes and the husband’s vector contains five, so we estimate a total of 290 parameters, most of which have no meaningful interpretation because their role is to make the first step of the control function approach as close as possible to the first stage of our 2SLS estimation. In Table 8, we display only the meaningful affinity parameters.

The top panel of Table 8 displays the matrix A , not accounting for assortative matching on education. The first cell of the matrix displays the affinity between the education of a wife and the number of co-wives of her husband: the affinity is positive (0.15) and statistically significant at the 5% level. Our logit estimation procedure on pairs of couples helps us interpret the magnitude of this affinity, which is an odds ratio (see equation 13). Let’s imagine a pair of couples composed of a woman with zero years of education, a woman with one year of education, a monogamous man (with zero other wives) and a polygamous man with exactly one other wife. An affinity of 0.15 means that the educated woman is $\exp(0.15) = 1.16$ times more likely to be married to the polygamous man than

to the monogamous man. This is, of course, when we do not account for matching on education.

Table 8: Matrix A (Husband’s Number of Wives)

Matrix A, Without Accounting for Matching on Education				
	Husband no. of wives	Husband education	Husband education control function	Other husband characteristics
Wife education	0.15** (0.07)			✓
Wife education control function	0.01*** (0.00)			✓
Other wife characteristics	✓			✓
Matrix A, Accounting for Matching on Education				
	Husband no. of wives	Husband education	Husband education control function	Other husband characteristics
Wife education	0.06 (0.09)	1.24*** (0.26)	0.01 (0.04)	✓
Wife education control function	0.00 (0.01)	0.01 (0.05)	0.06*** (0.00)	✓
Other wife characteristics	✓	✓	✓	✓

Notes: Observations: 3,582,030. Each observation is a pair of couples within the same village (women 25–60 and their husband). Standard errors clustered at the village level are in parentheses. Other husband and wife characteristics are the number of private schools in the village at age 7 (for women) and age 13 (for men), a quartic (fourth-degree) polynomial in age, and a quartic polynomial in age interacted with a vector of time-invariant village variables.

Next, we add to the vector of husbands’ attributes $\Delta \mathbf{x}_j$ all the Δ variables in equation (15) and the residual $\Delta \hat{\epsilon}_{mj}$ (the control function). Thus, we account for the matching between education of the husband and education of the wife.⁶⁵ We estimate a strong affinity between husband and wife education (1.24, significant at the 1% level).⁶⁶ This strong affinity reduces the affinity between wife education and the number of wives of the husband, which is divided by three and loses statistical significance. This shows that our

⁶⁵The wife’s vector contains 58 attributes, and the husband’s vector now contains 59 characteristics, so we estimate a total of 3,422 affinity parameters. Again, we display only the meaningful affinity parameters.

⁶⁶Again, the logit estimation procedure on pairs of couples helps us interpret the magnitude of this affinity. Let’s imagine a pair of couples composed of a man and a woman with zero years of schooling, and a man and a woman with one year of schooling. An affinity of 1.24 means that it is $\exp(1.24) = 3.46$ times more likely that schooling levels are the same within each marriage rather than the opposite.

IV results in Section 3 were, in large part, explained by assortative matching on education and the fact that educated men are more likely to be polygamous.

Appendix Table C.12 is like Table 8, but the number of wives of the husband is replaced by a binary variable equal to one if the husband is a polygamist, and to zero if he has only one wife. Here again, the positive affinity of 0.33 we estimate between a woman’s education and whether her husband is a polygamist is reduced (to 0.27) when we account for the strong matching on education (1.23), though it retains statistical significance at the 10% level.

Table 9: Matrix A (Husband’s Number of Wives), With Wife Rank

Matrix A, Without Accounting for Matching on Education							
	Husband no. of wives ×(rank=1)	Husband no. of wives ×(rank≥ 2)	Husband education ×(rank=1)	Husband education ×(rank≥ 2)	Husband educ. control function ×(rank=1)	Husband educ. control function ×(rank≥ 2)	Other husband characteristics
Wife education	0.23* (0.12)	0.05 (0.12)					✓
Wife education control function	0.00 (0.00)	0.00 (0.00)					✓
Other wife characteristics	✓	✓					✓
Matrix A, Accounting for Matching on Education							
	Husband no. of wives ×(rank=1)	Husband no. of wives ×(rank≥ 2)	Husband education ×(rank=1)	Husband education ×(rank≥ 2)	Husband educ. control function ×(rank=1)	Husband educ. control function ×(rank≥ 2)	Other husband characteristics
Wife education	0.11 (0.13)	-0.03 (0.12)	0.89*** (0.22)	0.79*** (0.24)	-0.02 (0.05)	0.01 (0.09)	✓
Wife education control function	-0.01 (0.00)	-0.01** (0.01)	0.05*** (0.01)	0.07*** (0.02)	0.07*** (0.00)	0.05*** (0.00)	✓
Other wife characteristics	✓	✓	✓	✓	✓	✓	✓

Notes: Observations: 3,582,030. Each observation is a pair of couples within the same village (women 25–60 and their husband). Standard errors clustered at the village level are in parentheses. Other husband and wife characteristics are the number of private schools in the village at age 7 (for women) and age 13 (for men), a quartic (fourth degree) polynomial in age, and a quartic polynomial in age interacted with a vector of time-invariant village variables.

In light of the results of Table 7 — showing that the effect of women’s education on polygamy differs by rank in the marriage — we then consider the possibility that affinities between husband and wife attributes differ by rank. Table 9 replicates Table 8, but it intersects all the male attributes with wife rank. The first panel confirms the results of Table 7: the affinity between women’s education and polygamy is mostly explained by first wives. The affinity between wife education and their husband’s number of wives is equal to 0.24 (significant at the 10% level) for first wives, versus 0.05 (not significant) for

higher-order wives. When we consider, instead of number of wives, a binary variable equal to one if the husband is polygamous, we find a strongly negative affinity of -0.67 between women's education and husband polygamy for higher-order wives (Appendix Table C.13), but it is imprecisely estimated and not distinguishable from zero at conventional levels.

In the bottom panel of Table 9, we add education to the vector of husbands' attributes. Matching on education appears to be of a similar magnitude for first wives (0.89) and for higher-order wives (0.79), contrary to what one might expect from the literature (Matz, 2016; Reynoso, 2019). Accounting for this matching on education reduces the affinity between women's education and the number of co-wives for both first wives and higher-order wives. For first wives, the affinity is divided by 2 (0.11) and loses statistical significance. For higher-order wives, the affinity turns negative (-0.03), but is not statistically different from zero. In Appendix Table C.13, when considering a binary for polygamy rather than the number of wives of the husband, we estimate a strong negative affinity of -1.20 between the education of higher-order wives and polygamy, but this negative affinity is not statistically different from zero at conventional levels.

Though our results point towards the possibility of a negative affinity between women's education and polygamy for second and higher-order wives once we account for the matching on education, our estimates are unfortunately too imprecise for us to draw a strong conclusion.

5. Conclusion

This paper focuses on nonreligious public schools during the colonial period in Cameroon. Contrary to missionary schools, these secular schools did not explicitly fight polygamy. Secular education, however, was not without consequences for the marriage prospects of the women receiving it. Women who were of school age when a public school opened in their village were able to marry men who were younger, more educated, and more likely to earn a wage. They were themselves more likely to earn a wage, but their labor-market returns to education were much lower than for men. However, they were not more likely

to marry monogamously. In fact, women who received public education were more likely to end up in a polygamous marriage, as a first wife. We have provided evidence that this was largely explained by assortative matching on education: educated women were able to marry more-educated men, but these educated men, because they were doing well in the labor market, were also more likely to take a second wife later.

Education did not increase the probability of being married to a polygamist as a second or higher-order wife. When we account for matching on education, we even find a negative affinity between education and marrying a polygamist as a second or higher-order wife, but this negative affinity is not significantly different from zero, and our results are overall very noisy, making it hard to draw a firm conclusion.

Our paper has implications for the literature on education and polygamy, and, to an extent, for the literature on the marriage-market returns to education, which needs to account for the role of matching on the marriage market when interpreting reduced form estimates.

As in previous papers, for example Ashraf et al. (2020), we show that the effect of an educational reform is mediated by local cultural norms and customs. In a setting where polygamous unions are frequent, it is perhaps not that surprising that women who receive education and thus become more attractive on the marriage market end up marrying polygamous men. Norms and customs, however, can be changed by education if it is set up with the express aim of transforming them. In Africa in general, and in Cameroon in particular, Christian missionaries had a goal of promoting the monogamous model of marriage, and our paper provides some evidence that, unlike public secular education, private Christian education decreased the likelihood that a woman would marry a polygamist. This confirms the findings of previous literature and point towards cultural change and religious conversion as important channels for explaining the decline in African polygamy.

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Appendix

A. Data Appendix

A.1. 1976 Population Census of Cameroon

Our main data source is the Cameroonian population census of 1976, for which we have the whole population, with the exception of three districts out of 138 that were missing in the raw data.⁶⁷ For each individual, the census gives us sex, age (with some imprecision in the form of age heaping), education (last grade attended), marital status (whether an individual is single, divorced, widowed or married — and the number of wives for men), and some very scarce information about occupation.

The census does not directly give the line identifier of the spouse for married individuals, but we were able to match 91.7% of married women with their husband living in the same household from information on marital status (including the number of wives for men) and relationship to the household head. In most households, there was no ambiguity about the pairing of spouses (for instance, a household with one household head, two spouses of the household head, one married son of the household and one married other member of the household); however, in large, complex households, we were not always able to match all spouses (for instance when there were several married men and several married women listed as “other household members”). This also means that we were not able to match spouses living in separate households. We show in Appendix Table C.2 that education does not affect the probability of observing a married woman’s husband.

Men could over-report or under-report their number of wives. To check the accuracy of this variable in the census, we compare the average number of wives declared by men to the total number of married men divided by the total number of married women. These two figures should in theory be equal. We find that men declare an average of 1.356 wives, while the total number of married women divided by the total number of married men is 1.352. Men might slightly overstate their number of wives, but only by a very

⁶⁷The districts of Mvengue, Dzeng and Kribi, see Appendix Figure A.2.

small amount.

The census contains no observation on wages. To estimate labor-market returns to education, we use information on employment status, more specifically whether and individual is working for a wage. Only 14% of workers in the 1976 census are recorded as earning a wage (22% of men and 3% of women). Though a tiny minority of non-wage earners are owners of formal businesses, and though some wage earners might be employed in non-registered firms, the wage-earner/non-wage-earner divide largely reflects the formal/informal divide. In the context of a developing country in the 1970s, earning a wage rather than working as an independent should be interpreted as a sign of a higher socioeconomic status.

In Appendix Table A.1, we check the accuracy of 1976 census data by comparing the average of various characteristics for the cohort born between 1916 and 1951 (aged between 25 and 50 in 1976) with two more recent data sources: the 1987 census and the 1991 Demographic and Health Survey (DHS). Even for immutable characteristics, like birth year, or characteristics that rarely change after the age of 25, like education, we do not expect the average to be equal across data sources because of selective mortality. However, very large differences in averages might point toward a serious reporting problem in the census. Reassuringly, this is not the case. There are some minor discrepancies, like the fact average years of schooling for women goes from 0.64 in 1976 to 0.86 in 1987 and 0.65 in 1991 (a similar pattern is observed for men). This might be explained by selective mortality or by a problem with the reporting of years of schooling in 1987 (the school attendance variable is less impacted). The percentage of women in a polygamous union is very similar in 1976 (46%) and in 1987 (45%). The percentage of men in a polygamous union increases from 24% to 28%, which is expected if men take additional wives as they become older. The DHS gives marital status only for women aged between 15 and 49, we therefore focus our comparison with the censuses on the cohort born between 1942 and 1951. The percentage of married women in a polygamous union in this cohort in the DHS is 49%, slightly higher than in the 1976 and 1987 censuses, but the DHS sample is small (430 women).

Table A.1: Cross-Validation of 1976 Census Data

	Women Born 1916–1951			Men Born 1916–1951		
	Mean 1976 Census	Mean 1987 Census	Mean 1991 DHS	Mean 1976 Census	Mean 1987 Census	Mean 1991 DHS
Age	38.87	49.39	53.32	39.60	49.45	52.65
Years of schooling	0.64	0.86	0.65	2.38	2.72	2.57
Ever attended school	0.14	0.16	0.14	0.41	0.43	0.43
Married	0.76	0.62		0.76	0.84	
Single	0.06	0.05		0.18	0.09	
Widow	0.15	0.28		0.02	0.03	
Divorced	0.04	0.05		0.03	0.04	
Observations	1,109,860	881,718	1,745	962,622	836,679	1,623
	Married Women Born 1916–1951			Married Men Born 1916–1951		
	Mean 1976 Census	Mean 1987 Census		Mean 1976 Census	Mean 1987 Census	
In a polygamous union	0.46	0.45		0.24	0.28	
No. of spouses	1.00	1.00		1.35	1.39	
Observations	761,383	544,176		730,788	700,113	
	Married Women Born 1942–1951					
	Mean 1976 Census	Mean 1987 Census	Mean 1991 DHS			
In a polygamous union	0.45	0.46	0.49			
Observations	333,335	283,751	430			

Notes: This table compare various aggregate statistics for the cohort born between 1916 and 1951 in three data sources: the 1976 population census, the 1987 population census, and the 1991 Demographic and Health Survey (DHS). There are missing districts in both the 1976 census and the 1987 census. For the sake of comparability, we restrict the sample to districts present both in 1976 and 1987. In the 1991 DHS, information on age and schooling is available for all household members (men and women), but information on the type of union (polygamous/monogamous) is available only for women aged 15–49 (born between 1942 and 1976). For our comparison with the 1976 census, we therefore focus on the sample of women born between 1942 and 1951.

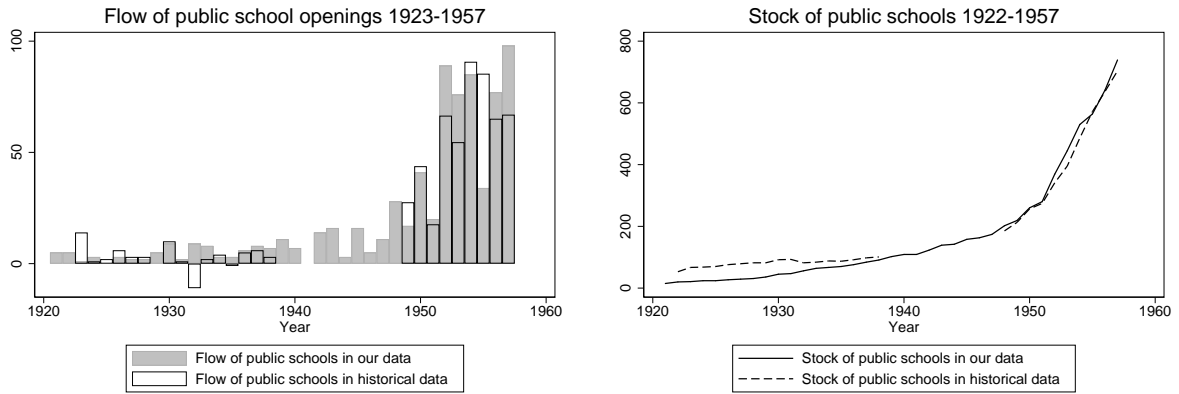
A.2. Cross-Validation of 2016 Administrative School Data

We use 2016 administrative school data giving the date of opening of every school in Cameroon to infer information about historical school supply. This relies on the assumption of no attrition — schools, once opened, did not close down. To show that attrition is not too much of a problem, we cross-validate our source with historical data for the colonial period. We could not find historical data giving the yearly supply of schools at the regional level, let alone at the village level, but the reports sent by France and Britain to the League of Nations/United Nations give the total number of schools in Cameroon for the period 1922–1938 and 1948–1957.⁶⁸ Figure A.1 shows the total flow and stock of public schools in Cameroon between 1923 and 1957, according to 2016 administrative data and to historical data. Though there is some discrepancy, especially before World War II, the two series are reasonably close to one another. Reassuringly, from the middle of the 1930s onwards, the two sources agree on the total stock of public schools in Cameroon. If attrition were a major problem and a large number of public schools had closed between their opening date and 2016, then the total stock of schools given by 2016 administrative data would be lower than the stock given by historical sources, but it is not the case. For private (missionary) schools, attrition is much more of a problem: historical data always show many more private schools than what can be inferred from 2016 administrative data.⁶⁹ This is one of the reasons for using only the variation in the number of public schools to identify the effects of education on marriage markets.

⁶⁸Sources: Great Britain, Colonial office (1922–1938, 1949–1959); France, Ministère des Colonies (1921–1938, 1947–1957). French reports also give the number of schools in 1921 and 1947. Between 1948 and 1959, British reports give only the total number of students; to infer the total number of schools, we use the average number of students per school in 1938 (110).

⁶⁹For instance, in 1955, we count 564 private schools in Cameroon while there were more than 2,000 private schools according to historical data. That said, the source for the number of private schools in historical data is the missions themselves, and their figures might be inflated (Dupraz, 2019). However, while 2016 administrative data seem to give us a reasonably accurate measure of the historical stock and flow of *public* schools, we need to think of the *private* schools in these data as the ones that were high-quality enough to stay open until 2016. The French report to the UN in 1949 gives the number of private schools with three classes or more; at about 119, it is reasonably close to the number of private schools that opened before 1949 in 2016 administrative data (79).

Figure A.1: Comparison of 2016 Administrative Data with Historical Data



Sources: Great Britain, Colonial office (1922–1938, 1949–1959); France, Ministère des Colonies (1921–1938, 1947–1957).

A.3. Geolocation of Villages and Schools

We geolocated villages and schools from each database from the name of the locality, using a variety of gazetteers.⁷⁰ In the 16 districts of the Bamiléké region, village-level geolocation was impossible; we therefore excluded from our estimations individuals born in these districts.⁷¹ We also excluded from our estimations individuals born in Yaoundé (the administrative capital) and Douala (the economic capital) because of the difficulty of precisely geolocating enumeration areas within these agglomerations.⁷² We were able to geolocate 99.9% of the remaining villages in the census and 98.3% of individuals — even though we geolocated almost every village, errors in village code entry prevented geolocation for some individuals. Appendix Figure A.2 maps these villages and the districts where geolocation was impossible. We geolocated all 3,765 schools in the administrative school database opened before 1976 from the name of the locality.⁷³ For 40 schools (1.1%)

⁷⁰We used the Fallingrain Global Gazetteer (<http://www.fallingrain.com>), the GeoNames geographical database (<http://www.geonames.org>), the website of the Cameroonian Ministry of Energy and Water (<http://www.mng-cameroon.org/SIG/>) and the Wiki World Map OpenStreetMap (<https://openstreetmap.org>). Geographical information about non-located villages was inferred by taking the mean of located villages in the same canton (a canton is a group of about 10 villages).

⁷¹These districts represent about 14% of the population. In the Bamilékés, village codes in the raw data did not match village codes in the locality file. The 16 districts are Mbouda, Batcham, Galim, Bafang, Bana, Bandja, Kekem, Dcshang, Penka-Michel, Bafoussam, Bandjoun, Bamendjou, Bangou, Bazou, Tonga and Bangangte.

⁷²These cities represent eight districts, corresponding to roughly 5% of the population in 1976.

⁷³We were able to locate the village for 3,333 schools. For 392 schools, we use the location of the town (“ville”), a geographical division larger than the village.

that could not be geolocated from the name of the locality, we used the centroid of the district.

Finally, we combined both geolocated sources in a GIS software to build the stock and flows of schools (total, public and private) in a radius of 10 km around each village at each date. Figure A.3 gives a graphical illustration of the procedure.

A.4. Other Sources of Data

Temperature and precipitation. Temperature and precipitation data are from WorldClim (<http://www.worldclim.org>). They are averages over the period 1950–2000.

Elevation and ruggedness. Elevation data come from NASA Shuttle Radar Topography Mission (<http://srtm.csi.cgiar.org>). Ruggedness is computed from elevation data using the slope tool in ArcGIS.

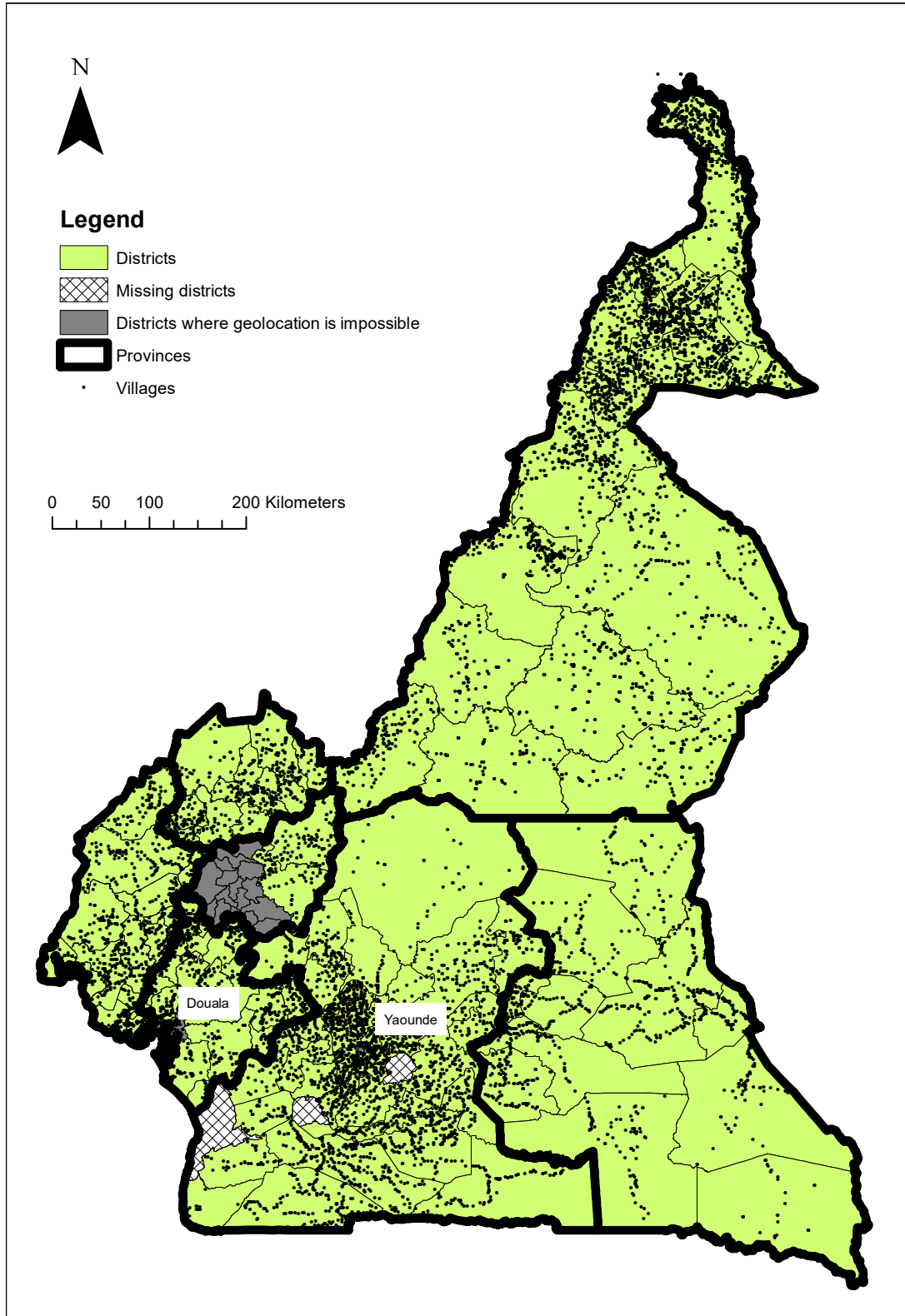
Malaria stability index. The malaria stability index comes from Kiszweski et al. (2004) and was downloaded from <https://sites.google.com/site/gordoncmccord/datasets>.

Agricultural suitability. Agricultural suitability is suitability for rainfed crops excluding forest ecosystems, an index ranging from 0 to 11. It comes from Global Agro-Ecological Zones: <http://webarchive.iiasa.ac.at/Research/LUC/GAEZ/index.htm>.

Railroads in 1922. The main railroads in 1922 were a narrow-gauge line from Victoria (now Limbe) to Soppo in British Cameroon, the Northern railway from Douala to Nkongsamba and the Central railway from Douala to Eseka in French Cameroon. The exact delineation of railroads comes from <http://diva-gis.org>.

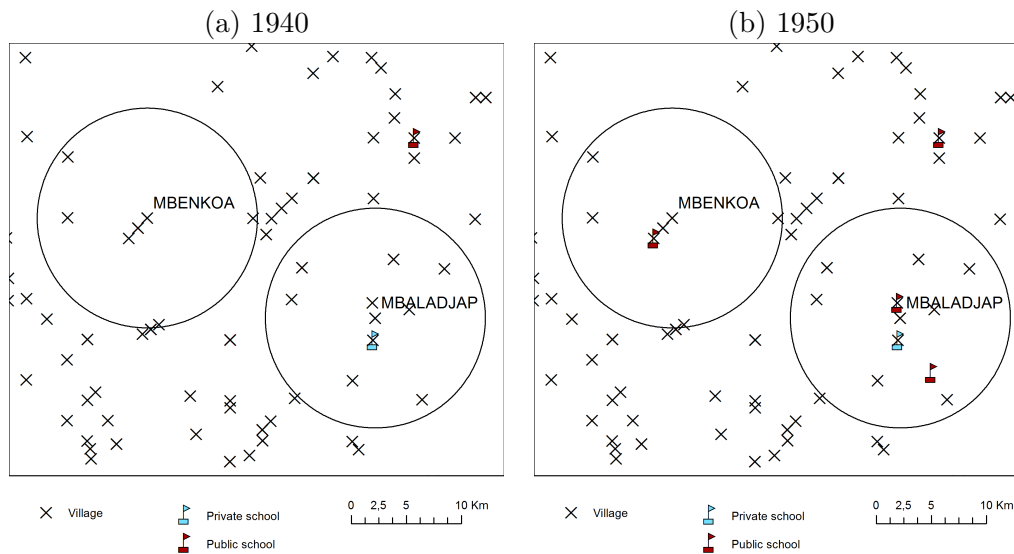
Main rivers. Data on the location of rivers come from www.naturalearthdata.com/downloads/10m-physical-vectors/.

Figure A.2: Villages, Districts and Provinces in the 1976 Census



Note: Authors' map from 1976 Cameroonian population census data.

Figure A.3: Construction of the School-Supply Variable: Example



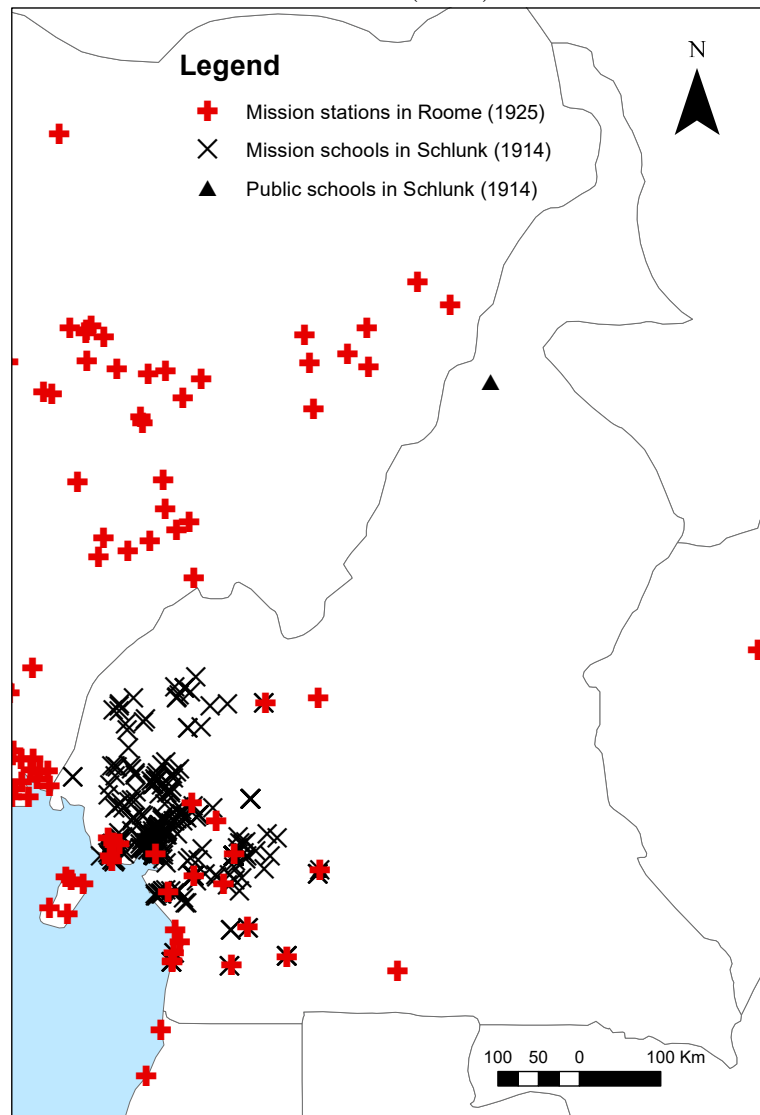
Note: A circle represents a 10-km radius around a given village. In 1940, the village of Mbenkoa has zero schools in a radius of 10 km and the village of Mbaladjap has one (private) school. In 1950, the village of Mbenkoa has one (public) school and the village of Mbaladjap has three schools (one private, two public).

Towns in 1922. Data on the location of towns in 1922/1923 come from France, Ministère des Colonies (1922) and Great Britain, Colonial office (1923).

Roome mission stations. The mission station map of Roome (1925) was digitized by Nunn (2010) and is available at <https://scholar.harvard.edu/nunn/pages/data-0>.

German mission/public schools in 1913. Data on German mission schools and public schools in 1913 comes from Schlunk (1914) and was digitized by Dupraz (2019). See Figure A.4 below.

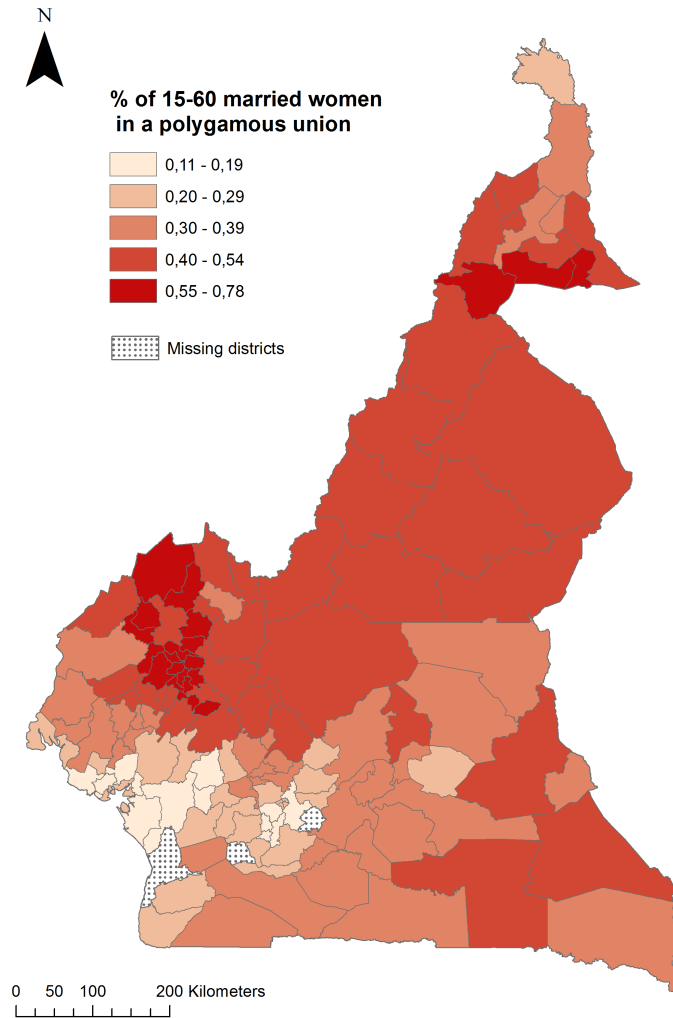
Figure A.4: Mission Stations in Roome (1925) and Schools in Schlunk (1914)



Note: Authors' map from data in Roome (1925) and Schlunk (1914).

B. Additional Descriptive Statistics

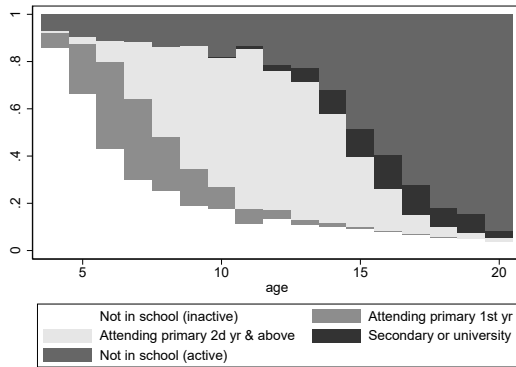
Figure B.1: Share of Married Women Aged 15–60 in a Polygamous Union in 1976



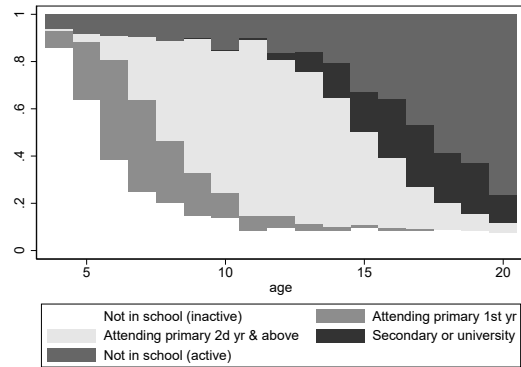
Note: Authors' map from 1976 Cameroonian population census data.

Figure B.2: School Attendance by Age in the 1976 Census

(a) Women

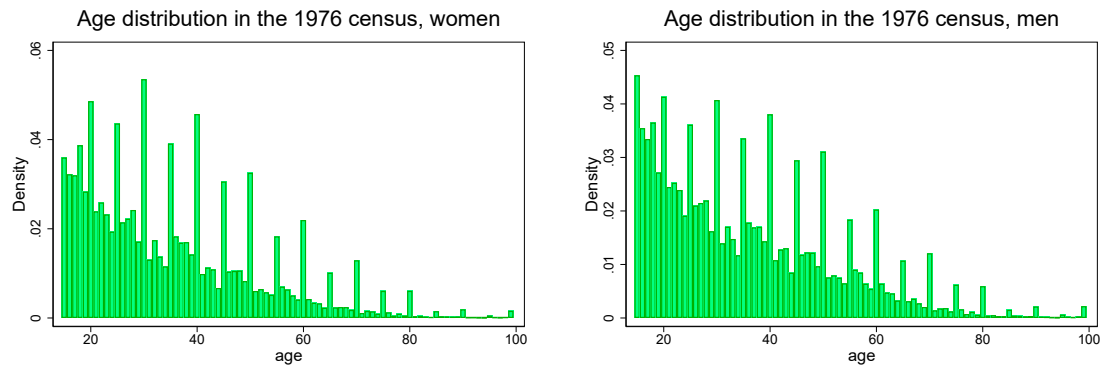


(b) Men



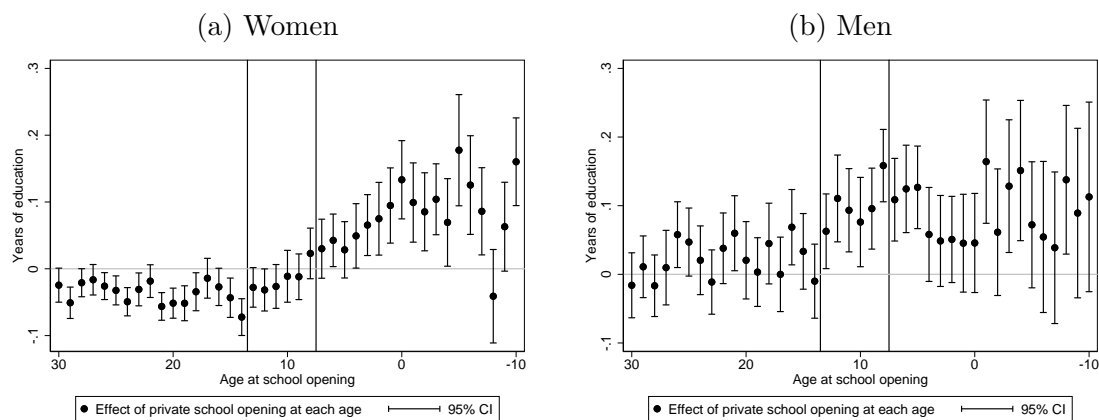
Notes: this figure represents, for men and women separately, the percentage of each age group attending various grades of education in 1976. Secondary and higher education are grouped together for readability, but this category corresponds mainly to students attending secondary school (only 0.09% of women and 0.36% of men aged 18-20 attended higher education in 1976). The “Not in school (active)” category includes individuals not attending school and involved in household work. Very few men and women attend higher education in 1976 Cameroon.

Figure B.3: Age Heaping in the 1976 Census



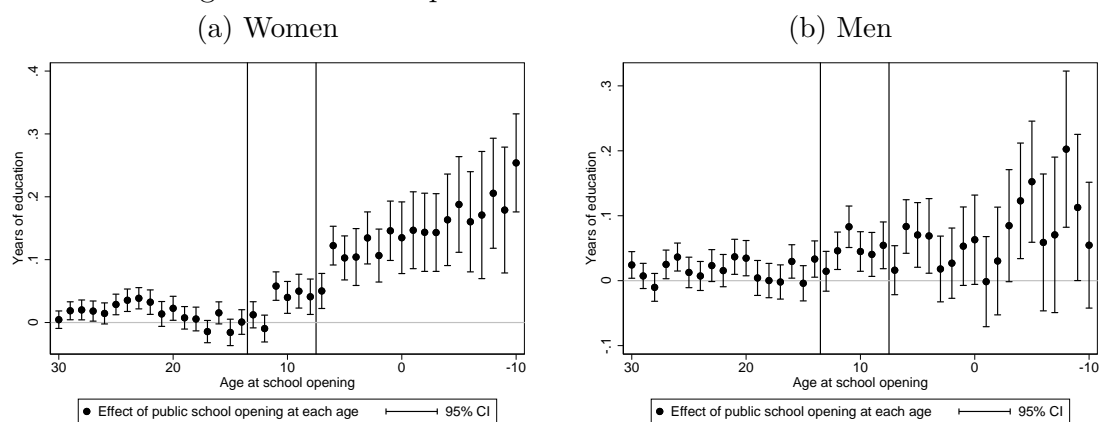
C. Additional Results

Figure C.1: Event-Study Graphs: Effect of Private School Openings on Education



Note: Both figures display the $\beta_{private}$ coefficients of equation (1), estimated separately for men and women. Standard errors are clustered at the village level.

Figure C.2: Event-Study Graphs: Effect of Public School Openings on Education, Including Schools that Opened After 1960



Note: Both figures display the β_{public} coefficients of equation (1), estimated separately for men and women. Contrary to Figure 2, we also consider the schools that opened after 1960.

Figure C.3: Selection of the Age Maximizing the First-Stage F-test

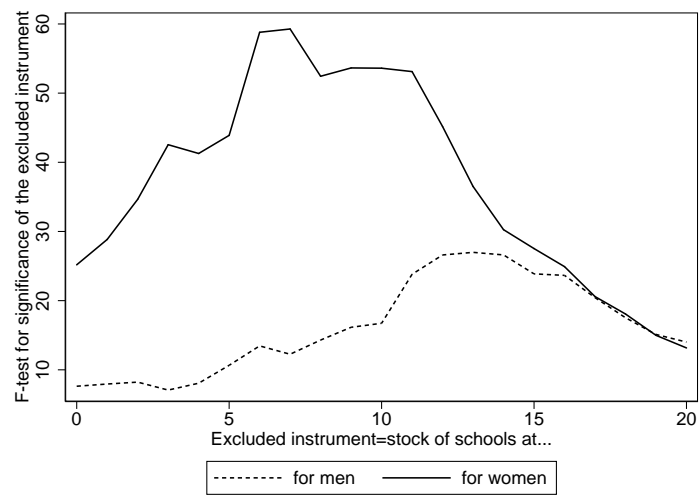


Table C.1: Invariance of First Stage and Reduced Form to Interacting Cohort Trend with Ethnic-Group-Level Controls

	(1)	(2)	(3)	(4)
Panel A: Married Women				
	Years of Education		No. of Co-Wives	
No. public schools at age 7	0.0939*** (0.0115)	0.0906*** (0.0115)	0.0154 (0.0094)	0.0164* (0.0096)
No. private schools at age 7	0.1046*** (0.0206)	0.1053*** (0.0200)	-0.0403*** (0.0112)	-0.0380*** (0.0110)
Observations	322,983	322,983	322,983	322,983
Panel B: Married Men				
	Years of Education		No. of Wives	
No. public schools at age 13	0.0327** (0.0132)	0.0337** (0.0134)	0.0113*** (0.0034)	0.0102*** (0.0035)
No. private schools at age 13	0.0744*** (0.0229)	0.0756*** (0.0226)	0.0086 (0.0054)	0.0079 (0.0059)
Observations	319,865	319,865	319,865	319,865
Village fixed effects	✓	✓	✓	✓
Cohort fixed effects	✓	✓	✓	✓
Cohort fixed effects × Br. Cameroon	✓	✓	✓	✓
Village controls × cohort quartic trend	✓	✓	✓	✓
Ethnic controls × cohort quartic trend		✓		✓

Notes: Source for ethnic group level control: *Ethnographic Atlas*, corrected version by Gray (1999). We overlay the map of villages over Murdock's map of African ethnic groups (Murdock, 1959) to join each location with ethnic group variables from the *Ethnographic Atlas*: number of levels beyond the local community in the jurisdictional hierarchy (political centralization), and whether polygyny is general or limited. These variables are missing for a third of villages, which is why we do not include these controls in the main specification. This table shows that the first stage and reduced form coefficients are invariant to interacting cohort trends with these ethnic group level controls. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.2: Returns to Education: Additional Results

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women						
	In Labor Force	Wage Earner (Given in LF)	Is Married	In a Polygamous Union ^(a)	No. of Co-Wives ^(a)	Husband Observed
<i>Reduced Form</i>						
No. public schools at age 7	-0.003 (0.003)	0.004*** (0.001)	0.004* (0.002)	0.006** (0.002)	0.023*** (0.008)	0.002 (0.002)
<i>OLS</i>						
Years of schooling	0.0002 (0.0013)	0.0217*** (0.0017)	-0.0149*** (0.0009)	-0.0159*** (0.0009)	-0.0476*** (0.0025)	-0.0099*** (0.0005)
<i>2SLS</i>						
Years of schooling	-0.028 (0.033)	0.033*** (0.008)	0.038* (0.023)	0.053** (0.025)	0.217** (0.086)	0.019 (0.018)
K-P F-stat	59.20	86.44	59.59	53.27	53.27	88.09
Mean dep. variable	0.567	0.008	0.764	0.353	1.381	0.917
Observations	696,119	386,450	698,736	654,906	654,906	533,876
No. clusters	9,213	7,307	9,214	9,178	9,178	9,093
Panel B: Men						
	In Labor force	Wage Earner (Given in LF)	Is Married	In a Polygamous Union ^(a)	No. of Wives ^(a)	
<i>Reduced Form</i>						
No. public schools at age 13	0.0012 (0.0008)	0.0053*** (0.0015)	0.0015 (0.0020)	0.0042*** (0.0014)	0.0089** (0.0036)	
<i>OLS</i>						
Years of Schooling	0.002*** (0.000)	0.037*** (0.001)	0.017*** (0.001)	0.005*** (0.000)	0.027*** (0.001)	
<i>2SLS</i>						
Years of schooling	0.024 (0.015)	0.104*** (0.034)	0.029 (0.039)	0.081** (0.032)	0.173** (0.077)	
K-P F-stat	26.45	29.68	28.43	28.37	28.37	
Mean dep. variable	0.953	0.079	0.781	0.201	1.060	
Observations	604,931	558,430	604,670	602,576	602,576	
No. clusters	9,203	9,166	9,201	9,199	9,199	

Notes: Panel A sample: In columns (1) and (3)through (5), all nonmigrant women aged 25–60 in 1976 (the discrepancy in number of observations is explained by missing information on labor force participation, marital status and number of wives of the husband for married women); in column (2), all nonmigrant women aged 25–60 in the labor force in 1976; in column (6) all nonmigrant married women aged 25–60 in 1976. Panel B sample: in columns (1) and (3)through (5), all nonmigrant men aged 25–60 in 1976 (the discrepancy in number of observations is explained by missing information on labor force participation and marital status); in column (2), all nonmigrant men aged 25–60 in the labor force in 1976. (a) The variables “In a polygamous union” and “number of (co)-wives” are *not* conditional on being currently married. Single, divorced, and widowed men and women are not currently polygamous, the variable “In a polygamous union” is zero for them, and their number of wives (co-wives for women) is zero. All regressions control for the stock of private schools at age 7 (for women) or age 13 (for men), village fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant village controls interacted with a quartic cohort trend: precipitation, temperature, elevation, ruggedness, a malaria stability index, agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, Roome mission station, 1913 German mission school, and 1913 German government school (see Data Appendix). Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.3: Returns to Education: Geographic Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women						
	Christian-Majority Regions (South)			Muslim-Majority Regions (North)		
	<i>First Stage</i>	<i>Reduced Form</i>		<i>First Stage</i>	<i>Reduced Form</i>	
	Years of Schooling	Husband Polygamous	Husband No. of Wives	Years of Schooling	Husband polygamous	Husband No. of Wives
No. public schools at age 7	0.116*** (0.012)	0.001 (0.002)	0.015** (0.007)	0.029*** (0.009)	0.007 (0.004)	0.006 (0.012)
K-P F-stat	90.42			10.15		
	<i>2SLS</i>			<i>2SLS</i>		
Years of schooling		0.010 (0.021)	0.135** (0.061)		0.254 (0.166)	0.226 (0.409)
Mean dep. variable	0.436	0.458	1.843	0.061	0.498	1.852
Observations	319,800	319,800	319,800	170,245	170,245	170,245
No. clusters	5,730	5,730	5,730	3,309	3,309	3,309
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B: Men						
	Christian-Majority Regions (South)			Muslim-Majority Regions (North)		
	<i>First Stage</i>	<i>Reduced Form</i>		<i>First Stage</i>	<i>Reduced Form</i>	
	Years of Schooling	Polygamous	No. of Wives	Years of Schooling	Polygamous	No. of wives
No. public schools at age 13	0.047*** (0.011)	0.003* (0.002)	0.007** (0.003)	0.066*** (0.022)	0.003 (0.003)	0.002 (0.005)
K-P F-stat	18.70			9.48		
	<i>2SLS</i>			<i>2SLS</i>		
Years of schooling		0.065* (0.040)	0.146* (0.075)		0.041 (0.047)	0.026 (0.081)
Mean dep. variable	1.987	0.245	1.340	0.359	0.283	1.393
Observations	313,540	313,540	313,540	156,707	156,707	156,707
No. clusters	5,775	5,775	5,775	3,316	3,316	3,316

Notes: Sample: in Panel A, columns (1) to (3), nonmigrant married women aged 25–60 in 1976 and living in Christian-majority *départements*; in Panel A, columns (4) to (6), nonmigrant married women aged 25–60 in 1976 and living in Muslim-majority *départements*; in Panel B, columns (1) to (3), nonmigrant married men aged 25–60 in 1976 and living in Christian-majority *départements*; in Panel B, columns (4) to (6), nonmigrant married men aged 25–60 in 1976 and living in Muslim-majority *départements*. Muslim-majority *départements* are *départements* where Muslims represented the majority of the population in 2005 (earlier censuses do not give information on religion). These correspond to the 1976 *départements* of Adamaoua, Benoue, Diamare, Logone et Chari, and Bamoun. All other *départements* are Christian-majority *départements*. All regressions control for the stock of private schools at age 7, village fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant village controls interacted with a quartic cohort trend. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.4: Returns to Education, District-Level Instrument: Women

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Sample of Nonmigrants			Full Sample			
	<i>First Stage</i>			<i>First Stage</i>			
	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling
Average No. public schools at age 7	0.331*** (0.059)	0.297*** (0.058)	0.297*** (0.058)	0.282*** (0.056)	0.283*** (0.056)	0.262*** (0.055)	0.262*** (0.055)
K-P F-stat	31.10	26.37	26.37	25.49	25.92	22.66	22.66
	Wage Earner	Husband Polygamous	Husband No. of wives	Migrant	Wage Earner	Husband Polygamous	Husband No. of wives
	<i>Reduced Form</i>			<i>Reduced Form</i>			
Average No. public schools at age 7	0.004*** (0.001)	0.012* (0.07)	0.058*** (0.019)	0.012 (0.007)	0.005*** (0.002)	0.004 (0.016)	0.036** (0.017)
	<i>Two-Stage Least Squares</i>			<i>Two-Stage Least Squares</i>			
Years of schooling	0.013*** (0.004)	0.040* (0.024)	0.199*** (0.072)	0.042* (0.023)	0.017*** (0.004)	0.016 (0.024)	0.140** (0.065)
Mean dep. variable	0.004	0.471	1.846	0.271	0.014	0.439	1.768
Observations	706,199	499,221	499,221	979,715	968,102	680,232	680,232
No. clusters	112	112	112	112	112	112	112

Notes: Samples: In column (1), nonmigrant women aged 25–60 in 1976; in columns (2) and (3), nonmigrant married women aged 25–60 in 1976; in columns (4) and (5), all women aged 25–60 in 1976 (the slight discrepancy in number of observations is explained by missing employment status); in columns (6) and (7), all married women aged 25–60 in 1976. The average number of public schools at age 7 for a woman born in year c in district d is the average number of public schools in the village at age 7 for nonmigrant women born in year c in district d . All regressions control for the district average number of private schools at age 7, district fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant district controls interacted with a quartic cohort trend. The time-invariant district controls are the time-invariant village controls of Table 6 averaged over all villages of a district: precipitation, temperature, elevation, ruggedness, a malaria stability index, agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, Roome mission station, 1913 German mission school, and 1913 German government school. Standard errors clustered at the district level are in parenthesis. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.5: Returns to Education, District-Level Instrument: Men

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Sample of Nonmigrants			Full Sample			
	<i>First Stage</i>			<i>First Stage</i>			
	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling	Years of Schooling
Average No. public schools at age 13	0.090** (0.035)	0.081** (0.036)	0.081** (0.036)	0.059* (0.031)	0.06* (0.031)	0.041 (0.032)	0.041 (0.032)
K-P F-stat	6.56	5.09	5.09	3.63	3.73	1.66	1.66
	Wage Earner	Polygamous	No. of Wives	Migrant	Wage Earner	Polygamous	No. of Wives
	<i>Reduced Form</i>			<i>Reduced Form</i>			
Average No. public schools at age 13	0.006* (0.003)	0.010** (0.005)	0.022** (0.009)	0.012** (0.005)	0.004 (0.004)	0.004 (0.005)	0.011 (0.009)
	<i>Two-Stage Least Squares</i>			<i>Two-Stage Least Squares</i>			
Years of schooling	0.066 (0.049)	0.118 (0.081)	0.266 (0.166)	0.197 (0.123)	0.060 (0.070)	0.103 (0.151)	0.268 (0.321)
Mean dep. variable	0.074	0.257	1.357	0.287	0.183	0.237	1.324
Observations	613,602	479,319	479,319	869,901	855,921	654,568	654,568
No. clusters	112	112	112	112	112	112	112

Notes: Samples: In column (1), nonmigrant men aged 25–60 in 1976; in columns (2) and (3), nonmigrant married men aged 25–60 in 1976; in columns (4) and (5), all men aged 25–60 in 1976 (the slight discrepancy in number of observations is explained by missing employment status); in columns (6) and (7), all married men aged 25–60 in 1976. The average number of public schools at age 13 for a man born in year c in district d is the average number of public schools in the village at age 13 for nonmigrant men born in year c in district d . All regressions control for the district average number of private schools at age 13, district fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant district controls interacted with a quartic cohort trend. The time-invariant district controls are the time-invariant village controls of Table 6 averaged over all villages of a district: precipitation, temperature, elevation, ruggedness, a malaria stability index, agricultural suitability, and distances to the nearest 1922 railroad, river, 1922 town, Roome mission station, 1913 German mission school, and 1913 German government school. Standard errors clustered at the district level are in parenthesis. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.6: Returns to Education Controlling for the Stock of Schools when Potential Spouses were of School Age

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women						
	Wage Earner	Ever Married	Husband Education	Husband Polygamous	Husband No. of Wives	Husband's Age ^(a)
<i>2SLS</i>						
Years of schooling	0.014*** (0.004)	-0.018 (0.014)	0.558*** (0.112)	0.066** (0.030)	0.221*** (0.083)	-1.613*** (0.581)
K-P F-stat	58.51	59.59	79.29	79.03	79.03	80.05
<i>2SLS Controlling for No. of Schools at Age 3</i>						
Years of schooling	0.009* (0.005)	-0.010 (0.015)	0.526*** (0.133)	0.051 (0.036)	0.240** (0.096)	-3.120*** (0.874)
No. public schools at age 3	0.001 (0.001)	-0.002 (0.003)	0.006 (0.016)	0.003 (0.005)	-0.004 (0.012)	0.318*** (0.100)
K-P F-stat	44.90	44.91	54.08	53.95	53.95	54.51
Mean dep. variable	0.004	0.941	0.950	0.472	1.846	12.234
Observations	693,045	698,736	491,152	490,045	490,045	491,806
No. clusters	9,205	9,214	9,040	9,039	9,039	9,042
Panel B: Men						
	Wage Earner	Ever Married	Wife(s) Education	Polygamous	No. of Wives	Wife(s) Age ^(a)
<i>2SLS</i>						
Years of schooling	0.087*** (0.031)	0.040 (0.037)	1.339*** (0.246)	0.101*** (0.039)	0.181** (0.070)	0.344 (0.549)
K-P F-stat	26.56	28.43	26.24	25.71	25.71	25.85
<i>2SLS Controlling for No. of Schools at Age 17</i>						
Years of schooling	0.052 (0.036)	-0.035 (0.050)	0.981*** (0.243)	0.067 (0.046)	0.087 (0.077)	-1.062* (0.624)
No. public schools at age 17	0.002* (0.001)	0.005*** (0.002)	0.022** (0.010)	0.002 (0.002)	0.006 (0.003)	0.085*** (0.031)
K-P F-stat	18.69	19.70	18.79	16.32	16.32	18.92
Mean dep. variable	0.073	0.840	0.637	0.258	1.358	-9.740
Observations	602,210	604,670	438,328	470,247	470,247	439,070
No. clusters	9,202	9,201	9,041	9,091	9,091	9,043

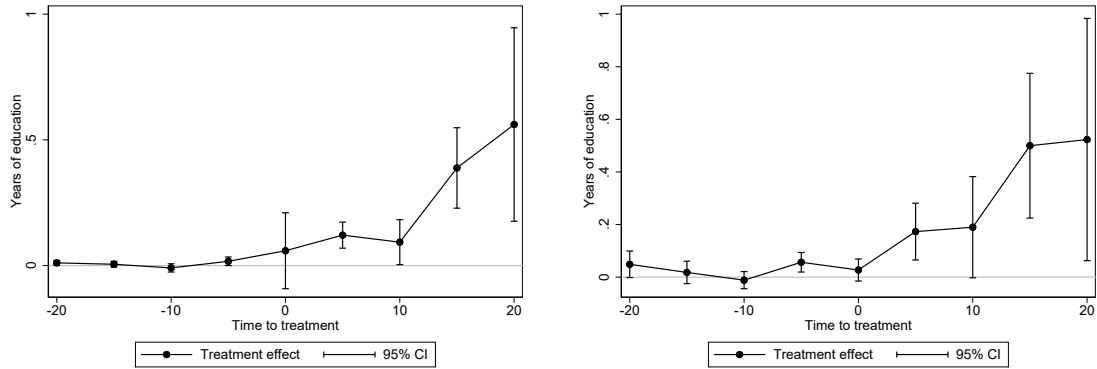
Notes: Samples and controls like in Table 6. Panel A reproduces the 2SLS estimates of Table 6, where woman's education is instrumented by the number of public schools in the village at age 7, before presenting estimates controlling for the number of public schools in the village at 3, that is when the cohort of men 10 years older was 13. Panel B reproduces the 2SLS estimates of Table 6, where man's education is instrumented by the number of public schools in the village at age 13, before presenting estimates controlling for the number of public schools in the village at 17, that is when the cohort of women 10 years younger was 7. (a) the dependent variable is spouse age minus own age. Because of the birth-year fixed effects, it does not matter for the coefficient of interest, but the mean of the dependent variable is more informative, giving average age difference. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.7: Replication of Main Results Grouping Individuals into Five-Year Cohorts

	(1)	(2)	(3)	(4)	(5)	(6)
	Women			Men		
	Years of Schooling	Husband's No. of Wives	Husband's No. of Wives	Years of Schooling	No. of Wives	No. of Wives
No. public schools at age a	0.1306*** (0.0151)	0.0135* (0.070)		0.073*** (0.0107)	0.0054* (0.0032)	
Years of schooling			0.1421* (0.0749)			0.0769 (0.0480)
Estimation	OLS	OLS	IV	OLS	OLS	IV
Sample	All women 25-60	Married women 25-60	Married women 25-60	All men 25-60	Married men 25-60	Married men 25-60
K-P F-stat			100.57			45.24
Observations	588,019	408,917	406,522	503,937	388,896	387,598

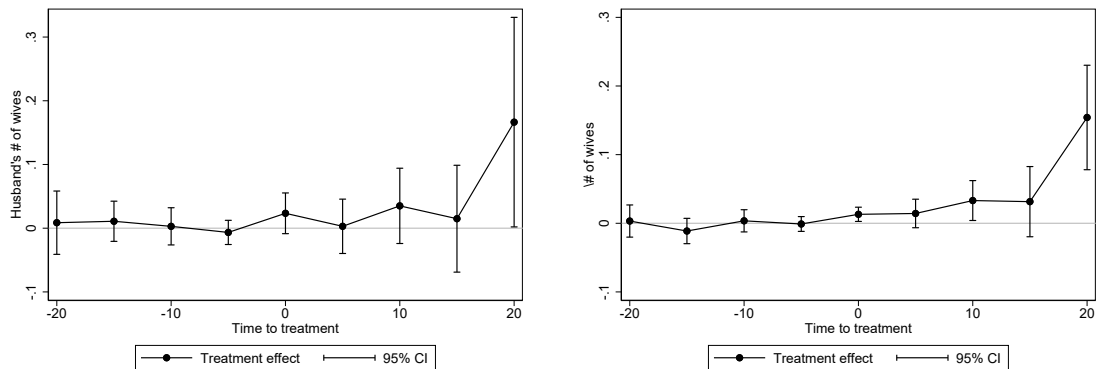
Notes: Standard errors clustered at the village level are in parentheses. This table replicates our main results grouping individuals into five-year cohorts centered around ages ending in 0 and 5. Each individual is attributed the central birth date of their cohort and the corresponding number of schools at age a (7 for women and 13 for men). All regressions control for cohort and village fixed effects, cohort fixed effects interacted with a binary for British Cameroon, time-invariant village controls interacted with a village quartic trend, and the number of private schools in the village at a .

Figure C.4: DID_M Estimator, Placebo and Dynamic Effects: First Stage
 (a) Women (b) Men



Notes: Placebo and dynamic effects estimated using the Stata command provided by de Chaisemartin and D'Haultfoeulle (2020). Individuals grouped into five-year cohorts around ages ending in 0 and 5 (see explanation in main text). Added controls are like in the main estimation: cohort fixed effects interacted with a dummy for British Cameroon, village time-invariant controls interacted with cohort quadratic trends, and number of private schools at age 7.

Figure C.5: DID_M Estimator, Placebo and Dynamic Effects: Reduced Form
 (a) Women (b) Men



Notes: Placebo and dynamic effects estimated using the Stata command provided by de Chaisemartin and D'Haultfoeulle (2020). Individuals grouped into five-year cohorts around ages ending in 0 and 5 (see explanation in main text). Added controls are like in the main estimation: cohort fixed effects interacted with a dummy for British Cameroon, village time-invariant controls interacted with cohort quadratic trends, and number of private schools at age 7.

Table C.8: Results Using the DID_M Estimator

	Women				Men			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Years of Schooling	Years of Schooling	Husband's No. of wives	Ratio (3)/(2)	Years of Schooling	Years of Schooling	# of Wives	Ratio (7)/(6)
Average of 5 dynamic effect	0.225*** (0.049)	0.244*** (0.049)	0.049* (0.027)	0.201	0.217*** (0.067)	0.283*** (0.080)	0.049*** (0.014)	0.173
Sample	All women 25-60	Married women 25-60	Married women 25-60		All men 25-60	Married men 25-60	Married men 25-60	
Observations	542,203	381,056	383,365		461,474	350,929	352,098	

Notes: Dynamic effects estimated using the Stata command provided by de Chaisemartin and D'Haultfoeuille (2020). Bootstrapped standard errors are in parentheses. Added controls are like in the main estimation: cohort fixed effects interacted with a dummy for British Cameroon, village time-invariant controls interacted with cohort quadratic trends, and number of private schools at age 7.

Table C.9: Labor-Market and Marriage-Market Returns to Education Restricting the Sample to Men and Women Aged 25–45 in 1976

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women						
	Wage Earner	Ever Married	Husband Education	Husband Polygamous	Husband No. of Wives	Husband's Age ^(a)
<i>2SLS on the Full Sample (25–60)</i>						
Years of schooling	0.014*** (0.004)	-0.018 (0.014)	0.558*** (0.112)	0.066** (0.030)	0.221*** (0.083)	-1.613*** (0.581)
K-P F-stat	58.51	59.59	79.29	79.03	79.03	80.05
Mean dep. variable	0.004	0.941	0.950	0.472	1.846	12.234
Observations	693,045	698,736	491,152	490,045	490,045	491,806
No. clusters	9,205	9,214	9,040	9,039	9,039	9,042
<i>2SLS on the Young Sample (25–45)</i>						
Years of schooling	0.012** (0.005)	-0.029* (0.015)	0.381*** (0.126)	0.083** (0.035)	0.258*** (0.090)	-1.137* (0.691)
K-P F-stat	59.10	59.20	62.16	62.16	62.16	62.65
Mean dep. variable	0.005	0.938	1.026	0.478	1.844	12.830
Observations	514,620	519,036	399,679	398,665	398,665	400,095
No. clusters	9,062	9,074	8,881	8,880	8,880	8,883
Panel B: Men						
	Wage Earner	Ever Married	Wife(s) Education	Polygamous	No. of Wives	Wife(s) Age ^(a)
<i>2SLS on the Full Sample (25–60)</i>						
Years of schooling	0.087*** (0.031)	0.040 (0.037)	1.339*** (0.246)	0.101*** (0.039)	0.181** (0.070)	0.344 (0.549)
K-P F-stat	26.56	28.43	26.24	25.71	25.71	25.85
Mean dep. variable	0.073	0.840	0.637	0.258	1.358	-9.740
Observations	602,210	604,670	438,328	470,247	470,247	439,070
No. clusters	9,202	9,201	9,041	9,091	9,091	9,043
<i>2SLS on the Young Sample (25–45)</i>						
Years of schooling	0.025 (0.036)	0.049 (0.064)	1.665*** (0.529)	0.119* (0.070)	0.253** (0.123)	1.976* (1.153)
K-P F-stat	10.25	11.62	8.46	8.90	8.90	8.20
Mean dep. variable	0.089	0.802	0.837	0.230	1.302	-7.791
Observations	413,728	415,673	288,737	312,693	312,693	289,250
No. clusters	8,998	8,999	8,719	8,805	8,805	8,723

Notes: This table reproduces the 2SLS estimates of Table 6 for women (Panel A) and men (Panel B) along with estimates obtained on the restricted sample of women and men aged 25–45 in 1976. This sample modification aside, samples and controls are the same as in Table 6. (a) The dependent variable is spouse age minus own age. Because of the birth-year fixed effects, it does not matter for the coefficient of interest, but the mean of the dependent variable is more informative, giving average age difference. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.10: Sensitivity to Violation of the Exclusion Restriction

	(1)	(2)	(3)	(4)	(5)	(6)
	Women			Women		
	Wage- Earner	Husband's No. of Wives	In a Polygamous Union	Wage- Earner	No. of Wives	In a Polygamous Union
Largest γ for significance at 5%	0.50	0.30	0.13	0.42	0.30	0.31
Largest γ for significance at 10%	0.58	0.43	0.28	0.52	0.44	0.44

Notes: This table is based on the “plausibly exogenous” approach of Conley et al. (2012). We allow the instrument (stock of schools) to enter directly in the equation of interest with a coefficient γ . This table reports the largest value of γ that can be assumed while maintaining the effect of education statistically different from zero, as a percentage of the effect of the stock of schools in the reduced form.

Table C.11: Labor-Market and Marriage-Market Returns to Education, 2SLS Estimation with Two Instruments

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women						
	Wage Earner	Ever Married	Husband Education	Husband Polygamous	Husband No. of wives	Husband's Age ^(a)
<i>2SLS (Instrument: No. Public Schools at Age 7)</i>						
Years of schooling	0.014*** (0.004)	-0.018 (0.014)	0.558*** (0.112)	0.066** (0.030)	0.221*** (0.083)	-1.613*** (0.581)
K-P F-stat	58.51	59.59	79.29	79.03	79.03	80.05
<i>2SLS (Instruments: No. Public Schools at Age 7 and No. Public School Openings Between Ages 8 and 13)</i>						
Years of schooling	0.014*** (0.004)	-0.010 (0.014)	0.569*** (0.115)	0.054* (0.030)	0.217** (0.086)	-1.999*** (0.603)
K-P F-stat	30.70	31.36	40.05	39.99	39.99	40.47
Mean dep. variable	0.004	0.941	0.950	0.472	1.846	12.234
Observations	693,045	698,736	491,152	490,045	490,045	491,806
No. clusters	9,205	9,214	9,040	9,039	9,039	9,042
Panel B: Men						
	Wage Earner	Ever Married	Wife(s) Education	Wife(s) Polygamous	No. of Wives	Wife(s) Age ^(a)
<i>2SLS (Instrument: No. Public Schools at Age 13)</i>						
Years of schooling	0.087*** (0.031)	0.040 (0.037)	1.339*** (0.246)	0.101*** (0.039)	0.181** (0.070)	0.344 (0.549)
K-P F-stat	26.56	28.43	26.24	25.71	25.71	25.85
<i>2SLS (Instruments: No. Public Schools at Age 7 and No. Public School Openings between Ages 8 and 13)</i>						
Years of schooling	0.096*** (0.035)	0.028 (0.038)	1.340*** (0.250)	0.106*** (0.040)	0.178** (0.072)	0.044 (0.541)
K-P F-stat	14.32	15.34	13.18	13.04	13.04	12.97
Mean dep. variable	0.073	0.840	0.637	0.258	1.358	-9.740
Observations	602,210	604,670	438,328	470,247	470,247	439,070
No. clusters	9,202	9,201	9,041	9,091	9,091	9,043

Notes: Samples like in Table 6. This table reproduces the 2SLS estimates of Table 6, along with estimates obtained using the same set of instruments for men and women: the number of public schools in the village at age 7 and the number of public school openings between ages 8 and 13. All regressions control for the stock of private schools at age 7 and the number of private school openings between ages 8 and 13, village fixed effects, cohort fixed effects, cohort fixed effects interacted with a British Cameroon binary, and a vector of time-invariant village controls (the same as in Table 6) interacted with a quartic cohort trend. (a) The dependent variable is spouse age minus own age. Because of the birth-year fixed effects, it does not matter for the coefficient of interest, but the mean of the dependent variable is more informative, giving average age difference. Standard errors clustered at the village level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table C.12: Matrix A (Husband Polygamous)

Matrix A, Without Accounting for Matching on Education				
	Husband polygamous	Husband education	Husband education control function	Other husband characteristics
Wife education	0.33** (0.14)			✓
Wife education control function	-0.00 (0.00)			✓
Other wife characteristics	✓			✓

Matrix A, Accounting for Matching on Education				
	Husband polygamous	Husband education	Husband education control function	Other husband characteristics
Wife education	0.27* (0.14)	1.23*** (0.25)	0.02 (0.04)	✓
Wife education control function	0.00 (0.01)	0.02 (0.05)	0.06*** (0.00)	✓
Other wife characteristics	✓	✓	✓	✓

Notes: Observations: 3,582,035. Each observation is a pair of couples within the same village (women 25–60 and their husband). Standard errors clustered at the village level are in parentheses. Other husband and wife characteristics are the number of private schools in the village at age 7 (for women) and age 13 (for men), a quartic (fourth degree) polynomial in age, and a quartic polynomial in age interacted with a vector of time-invariant village variables.

Table C.13: Matrix A (Husband Polygamous), With Wife Rank

Matrix A, Without Accounting for Matching on Education							
	Husband polygamous ×(rank=1)	Husband polygamous ×(rank≥ 2)	Husband education ×(rank=1)	Husband education ×(rank≥ 2)	Husband educ. control function ×(rank=1)	Husband educ. control function ×(rank≥ 2)	Other husband characteristics
Wife education	0.37** (0.18)	-0.67 (0.80)					✓
Wife education control function	-0.01 (0.01)	-0.12*** (0.02)					✓
Other wife characteristics	✓	✓					✓
Matrix A, Accounting for Matching on Education							
	Husband polygamous ×(rank=1)	Husband polygamous ×(rank≥ 2)	Husband education ×(rank=1)	Husband education ×(rank≥ 2)	Husband educ. control function ×(rank=1)	Husband educ. control function ×(rank≥ 2)	Other husband characteristics
Wife education	0.27 (0.19)	-1.20 (0.78)	0.90*** (0.22)	0.85*** (0.25)	-0.02 (0.04)	0.02 (0.09)	✓
Wife education control function	-0.02*** (0.01)	-0.17*** (0.03)	0.05*** (0.01)	0.09*** (0.02)	0.07*** (0.00)	0.05*** (0.00)	✓
Other wife characteristics	✓	✓	✓	✓	✓	✓	✓

Notes: Observations: 3,582,030. Each observation is a pair of couples within the same village (women 25–60 and their husband). Standard errors clustered at the village level are in parentheses. Other husband and wife characteristics are the number of private schools in the village at age 7 (for women) and age 13 (for men), a quartic (fourth degree) polynomial in age, and a quartic polynomial in age interacted with a vector of time-invariant village variables.

D. Mathematical Appendix

D.1. Proof of Theorem 1

The proof is structured in four parts; each of them proving an assertion of Theorem 1. We first prove the existence of an equilibrium, before characterizing the matching function. We then prove the uniqueness of the equilibrium, and we finally prove that the unique equilibrium maximizes total surplus.

Proof of existence. This proof requires a few notations. For any transferable utility function following (6) and (7), let us define:

$$\begin{cases} u(x, y, n, r) = u_0(x, y, n, r) - t(x, y, n, r) \\ v(x, y, n, r) = v_0(x, y, n, r) + t(x, y, n, r) \end{cases} \quad (18)$$

where $t(x, y, n, r)$ is the (positive or negative) utility transfer given by the husband to his wife of type y . $\pi_m(x, y, n, r) = f(x)\pi(y, n, r|x)$ is the number of men of type x who want a marriage with n spouses with a woman of type y in rank r ; similarly, $\pi_w(x, y, n, r) = g(y)\pi(x, n, r|y)$ is the number of women of type y who want to become the wife of rank r of a man of type x with $n - 1$ co-spouses. There is an equilibrium when $\pi_m(x, y, n, r) - \pi_w(x, y, n, r) = 0$. $\pi_m(x, y, n, r)$ and $\pi_w(x, y, n, r)$ are continuous functions of $t(x, y, n, r)$. Besides, for every (x, n, y, r) :

$$\begin{cases} \lim_{t(x,y,n,r) \rightarrow +\infty} \pi_m(x, y, n, r) - \pi_w(x, y, n, r) = 0 - g(y) \\ \lim_{t(x,y,n,r) \rightarrow -\infty} \pi_m(x, y, n, r) - \pi_w(x, y, n, r) = f(x) - 0 \end{cases} \quad (19)$$

Thus, the generalization of the intermediate values theorem to the multidimensional case, the Poincaré-Miranda theorem, proves there is a solution to the problem $\pi_m - \pi_w = 0$. (Technically, the Poincaré-Miranda theorem applies only to bounded sets. However, changing the variable to $t' = \tanh^{-1}(t)$ trivially solves this issue.)

Characterization of the equilibrium matching function. We start by writing the unconditional density functions $\pi(Y, x)$ and $\pi(x, y, n_Y, r)$ from the conditional density functions in equations (8) and (9):

$$\pi(Y, x) = \pi(Y|x)f(x) = \exp(U(x, Y) - a(x)) \quad (20)$$

$$\pi(x, y, n_Y, r) = \pi(x, n_Y, r|y)g(y) = \exp(v(x, y, n_Y, r) - b(y)) \quad (21)$$

where $\begin{cases} \exp a(x) = f(x)^{-1} [\sum_{Y' \in \mathcal{Y}} \exp(U(x, Y'))] \\ \exp b(y) = g(y)^{-1} [\sum_{x' \in \mathcal{X}, n'_Y \in \mathbb{N}, r' \leq n'_Y} \exp(v(x', y, n'_Y, r'))] \end{cases}$

In order to derive an expression for $\pi(x, y, n_Y, r)$ from the density function derived from men's behavior in equation (20), we start by writing the distribution of men of type x married with n_Y wives:

$$\begin{aligned} \pi(n_Y, x) &= \sum_{Y' \in \mathcal{Y}, n'_Y = n_Y} \pi(Y', x) \\ &= \sum_{Y' \in \mathcal{Y}, n'_Y = n_Y} \exp\left(\sum_{y' \in Y'} u(x, y', n'_Y, r_{y'}^{Y'}) - a(x)\right) \\ &= \exp(-a(x)) \sum_{Y' \in \mathcal{Y}, n'_Y = n_Y} \prod_{y' \in Y'} \exp u(x, y', n'_Y, r_{y'}^{Y'}) \end{aligned}$$

Because men marry an ordered set of women W , we can factorize the sum and write:

$$\begin{aligned} \pi(n_Y, x) &= \exp(-a(x)) \prod_{r'=1}^{n'_Y} \left[\sum_y \exp u(x, y, n_Y, r') \right] = \exp(-a(x)) \prod_{r'=1}^{n'_Y} C_{xn_Y r'} \\ C_{xn_Y r} &:= \sum_y \exp u(x, y, n_Y, r) \end{aligned}$$

To derive an expression for $\pi(x, y, n_Y, r)$, we need an expression of $\pi(y|n_Y, x, r)$ for each rank $r \leq n_Y$. The density of women of type y in position r , given that there are n_Y

co-spouses and that the husband is of type x , is:

$$\begin{aligned}
\pi(y|x, n_Y, r) &= \sum_{Y' \in \mathcal{Y}, n_Y, Y_r=r} \pi(Y'|x, n_Y, r) \\
&= \frac{\sum_{Y' \in \mathcal{Y}, n_Y, Y_r=r} \exp(U(x, Y'))}{\sum_{Y' \in \mathcal{Y}, n_Y} \exp(U(x, Y'))} \\
&= \frac{\sum_{Y' \in \mathcal{Y}, n_Y, Y_r=r} \prod_{y' \in Y'} \exp(u(x, y', n_Y, r_{y'}^{Y'}))}{\sum_{Y' \in \mathcal{Y}, n_Y} \prod_{y' \in Y'} \exp(u(x, y', n_Y, r_{y'}^{Y'}))} \\
&= \frac{\exp(u(x, y, n_Y, r)) \prod_{r'=1, r' \neq r}^{n_Y} \sum_{y'} \exp(u(x, y', n_Y, r'))}{\sum_{y'} \exp(u(x, y', n_Y, r)) \prod_{r'=1, r' \neq r}^{n_Y} \sum_{y'} \exp(u(x, y', n_Y, r'))} \\
&= \frac{\exp u(x, y, n_Y, r)}{\sum_{y'} \exp u(x, y', n_Y, r)} \\
&= \frac{\exp u(x, y, n_Y, r)}{C_{xn_Yr}}
\end{aligned}$$

Then we write $\pi(x, y, n_Y, r) = \pi(y|n_Y, x, r)\pi(n_Y, x)$ for each rank r and we obtain the following system, with two expressions for the matching function $\pi(x, y, n_Y, r)$, the first one coming from men's optimizing behavior, and the second one from women's:

$$\pi(x, y, n_Y, r) = \exp(u(x, y, n_Y, r) - a(x)) \frac{\prod_{r'=1}^{n_Y} C_{xn_Yr'}}{C_{xn_Yr}} \quad (22)$$

$$\pi(x, y, n_Y, r) = \exp(v(x, y, n_Y, r) - b(y)) \quad (23)$$

We note $\Phi(x, y, n_Y, r) = u(x, y, n_Y, r) + v(x, y, n_Y, r)$ the total systematic utility of a match and we write the systematic utility of a match for men and women as functions of total utility: $u(x, y, n_Y, r) = \frac{1}{2}(\Phi(x, y, n_Y, r) - \tau(x, y, n_Y, r))$ and $v(x, y, n_Y, r) = \frac{1}{2}(\Phi(x, y, n_Y, r) + \tau(x, y, n_Y, r))$. Substituting in (22) and (23) we find:

$$\pi(x, y, n_Y, r) = \exp \left[\frac{1}{2}(\Phi(x, y, n_Y, r) - \tau(x, y, n_Y, r)) - a(x) \right] \frac{\prod_{r'=1}^{n_Y} C_{xn_Yr'}}{C_{xn_Yr}} \quad (24)$$

$$\pi(x, y, n_Y, r) = \exp \left[\frac{1}{2}(\Phi(x, y, n_Y, r) + \tau(x, y, n_Y, r)) - b(y) \right] \quad (25)$$

Solving for the function $\tau(x, y, n_Y, r)$, we find:

$$\exp \tau(x, y, n_Y, r) = \exp(b(y) - a(x)) \frac{\prod_{r'=1}^{n_Y} C_{xn_Yr'}}{C_{xn_Yr}} \quad (26)$$

Substituting for $\tau(x, y, n_Y, r)$ in (25) we find:

$$\pi(x, y, n_Y, r) = \exp \left[\frac{\Phi(x, y, n_Y, r) - b(y) - a(x)}{2} \right] \frac{\prod_{r'=1}^{n_Y} C_{xn_Yr'}^{\frac{1}{2}}}{C_{xn_Yr}^{\frac{1}{2}}} \quad (27)$$

This expressions still includes the C_{xn_Yr} which are a function of $u(x, y, n_Y, r)$. To write C_{xn_Yr} as a function of only Φ , a and b , we start by substituting for $u(x, y, n_Y, r) = \frac{1}{2} (\Phi(x, y, n_Y, r) - \tau(x, y, n_Y, r))$ in the definition of C_{xn_Yr} :

$$\begin{aligned} C_{xn_Yr} &= \sum_y \exp u(x, y, n_Y, r) \\ &= \sum_y \exp \left[\frac{1}{2} (\Phi(x, y, n_Y, r) - \tau(x, y, n_Y, r)) \right] \end{aligned}$$

We then substitute for $\tau(x, y, n_Y, r)$ using (26) and solve for C_{xn_Yr} to find:

$$\begin{aligned} C_{xn_Yr} &= \sum_y \exp \left(\frac{\Phi(x, y, n_Y, r) - b(y) + a(x)}{2} \right) \frac{C_{xn_Yr}^{\frac{1}{2}}}{\prod_{r'=1}^{n_Y} C_{xn_Yr'}^{\frac{1}{2}}} \\ \tilde{C}_{xn_Yr} &:= \sum_y \exp \left(\frac{\Phi(x, y, n_Y, r) - b(y) + a(x)}{2} \right) \\ C_{xn_Yr} &= \tilde{C}_{xn_Yr}^2 \prod_{r'=1}^{n_Y} C_{xn_Yr'}^{-1} \\ \prod_{r=1}^{n_Y} C_{xn_Yr} &= \prod_{r=1}^{n_Y} \tilde{C}_{xn_Yr}^2 \prod_{r=1}^{n_Y} C_{xn_Yr}^{-n_Y} \\ \prod_{r=1}^{n_Y} C_{xn_Yr} &= \prod_{r=1}^{n_Y} \tilde{C}_{xn_Yr}^{\frac{2}{n_Y+1}} \\ C_{xn_Yr} &= \tilde{C}_{xn_Yr}^2 \prod_{r'=1}^{n_Y} \tilde{C}_{xn_Yr'}^{\frac{-2}{n_Y+1}} \end{aligned}$$

With this expression of C_{xn_Yr} , (27) is now only a function of the \tilde{C}_{xn_Yr} , which are functions

of Φ , a and b . Finally, substituting for C_{xn_Y} in (27) we find:

$$\begin{aligned}
\pi(x, y, n_Y, r) &= \exp \left[\frac{\Phi(x, y, n_Y, r) - b(y) - a(x)}{2} \right] \frac{\prod_{r'=1}^{n_Y} \tilde{C}_{xn_Y r'}^{\frac{1}{n_Y+1}}}{\tilde{C}_{xn_Y r} \prod_{r'=1}^{n_Y} \tilde{C}_{xn_Y r'}^{\frac{-1}{n_Y+1}}} \\
&= \exp \left[\frac{\Phi(x, y, n_Y, r) - b(y) - a(x)}{2} \right] \frac{\prod_{r'=1}^{n_Y} \tilde{C}_{xn_Y r'}^{\frac{2}{n_Y+1}}}{\tilde{C}_{xn_Y r}} \\
&= \exp \left[\frac{\Phi(x, y, n_Y, r) - b(y)}{2} - \frac{a(x)}{2 + n_Y} \right] \frac{\prod_{r'=1}^{n_Y} \left[\sum_{y'} \exp \left(\frac{\Phi(x, y', n_Y, r') - b(y')}{2} \right) \right]^{\frac{2}{n_Y+1}}}{\sum_{y'} \exp \left(\frac{\Phi(x, y', n_Y, r) - b(y')}{2} \right)}
\end{aligned}$$

Defining $\exp a(x, n_Y, r) = \exp a(x) \frac{\tilde{C}_{xn_Y r}^{\frac{2}{n_Y+1}}}{\prod_{r'=1}^{n_Y} \tilde{C}_{xn_Y r'}^{\frac{2}{n_Y+1}}}$, we write the more concise expression:

$$\pi(y, n_Y, x, r) = \exp \left[\frac{\Phi(x, y, n_Y, r) - b(y) - a(x, n_Y, r)}{2} \right]$$

Uniqueness of the equilibrium To show that the equilibrium is unique, we study the densities of men and women implied by the functions $a(x)$ and $b(y)$ defined above (20)/(21). The proof is more intuitively understood if we reason in terms of vectors: the function $b(y)$ defines a vector b in a space where each dimension is a distinct type of woman $y = y_w$, and the function $a(x)$ defines a vector a in a space where each dimension is a distinct type of man $x = x_m$. We want to show that for any distribution of men's attribute $f(x)$ and women's attributes $g(y)$, there is a single set of vectors (a, b) such that 1) the distribution of men $f(x)$ is equal to the distribution of men implied by the matching function $\pi(x, y, n_Y, r)$ and 2) the distribution of women $g(y)$ is equal to the distribution of women implied by the matching function $\pi(x, y, n_Y, r)$. That is, we want to show that for any $f(x)$ and $g(y)$ and for any utility function $\Phi(x, y, n_Y, r)$, there is a single set of vectors (a, b) such that:

$$g(y) = \pi(y, \emptyset) + \sum_{x, n_Y, r} \pi(x, y, n_Y, r) \tag{28}$$

$$f(x) = \pi(x, \emptyset) + \sum_{n_Y} \frac{1}{n_Y} \sum_{y, r} \pi(x, y, n_Y, r) \tag{29}$$

where $\pi(x, \emptyset) = \exp(-a(x))$ and $\pi(y, \emptyset) = \exp(-b(y))$ are respectively the number of single men and women⁷⁴ and we divide in (29) the density of women married with a man of type x and n_y co-wives by the number of wives $n_Y + 1$ to obtain the density of married men.

Firstly, we can show that for every vector b , there is a single vector a such that equation (29) is true. We write the density of men married with n_Y wives as, for each rank $r \leq n_Y$:

$$\begin{aligned}\pi(x, n_Y) &= \sum_y \pi(x, y, n_Y, r) \\ &= \sum_y \exp \left[\frac{\Phi(x, y, n_Y, r) - b(y)}{2} - \frac{a(x)}{1 + n_Y} \right] \frac{\prod_{r'=1}^{n_Y} \left[\sum_{y'} \exp \left(\frac{\Phi(x, y', n_Y, r') - b(y')}{2} \right) \right]^{\frac{2}{n_Y+1}}}{\sum_{y'} \exp \left(\frac{\Phi(x, y', n_Y, r) - b(y')}{2} \right)} \\ &= \exp \left[-\frac{a(x)}{1 + n_Y} \right] \prod_{r'=1}^{n_Y} \left[\sum_{y'} \exp \left(\frac{\Phi(x, y', n_Y, r') - b(y')}{2} \right) \right]^{\frac{2}{n_Y+1}}\end{aligned}$$

Note the derivatives:

$$\begin{aligned}\frac{\partial \pi(x, n_Y)}{\partial b(y)} &= -\frac{1}{1 + n_Y} \pi(x, n_Y) \sum_r \pi(y|x, n_Y, r) \\ \frac{\partial \pi(x, n_Y)}{\partial a(x)} &= -\frac{1}{1 + n_Y} \pi(x, n_Y)\end{aligned}$$

The total density of men implied by the matching function is:

$$\begin{aligned}\pi(x) &:= \pi(x, \emptyset) + \sum_{n_Y > 0} \pi(x, n_Y) \\ &= \exp(-a(x)) + \sum_{n_Y > 0} \exp \left[-\frac{a(x)}{1 + n_Y} \right] \prod_{r=1}^{n_Y} \left[\sum_y \exp \left(\frac{\Phi(x, y, n_Y, r) - b(y)}{2} \right) \right]^{\frac{2}{1+n_Y}}\end{aligned}$$

Note the derivatives:

$$\begin{aligned}\frac{\partial \pi(x)}{\partial b(y)} &= \sum_{n_Y > 0} \frac{\partial \pi(x, n_Y)}{\partial b(y)} = -\sum_{n_Y > 0} \frac{1}{1 + n_Y} \pi(x, n_Y) \sum_{r=1}^{n_Y} \pi(y|x, n_Y, r) \\ \frac{\partial \pi(x)}{\partial a(x)} &= -\exp(-a(x)) + \sum_{n_Y > 0} \frac{\partial \pi(x, n_Y)}{\partial a(x)} = -\sum_{n_Y \geq 0} \frac{1}{1 + n_Y} \pi(x, n_Y)\end{aligned}$$

⁷⁴This is because we normalize the utility of singles to 0.

(we define $\pi(x, 0) = \pi(x, \emptyset)$)

π is a strictly decreasing function of $a(x)$, $\lim_{a(x) \rightarrow +\infty} \pi(x) = 0$ and $\lim_{a(x) \rightarrow -\infty} \pi(x) = +\infty$. $\pi(x)$ does not depend on $a(x')$, $x' \neq x$. Hence there is a unique solution $a_b(x)$ to $\pi(x) = f(x)$ for every vector b . Note that the derivative of this solution with respect to $b(y)$ is:

$$\frac{\partial a_b(x)}{\partial b(y)} = -\frac{\frac{\partial \pi(x)}{\partial b(y)}}{\frac{\partial \pi(x)}{\partial a_b(x)}} = -\frac{\sum_{n_Y > 0} \frac{1}{1+n_Y} \pi(x, n_Y) \sum_{r=1}^{n_Y} \pi(y|x, n_Y, r)}{\sum_{n_Y \geq 0} \frac{1}{1+n_Y} \pi(x, n_Y)} \quad (30)$$

That is because $\frac{d\pi(x)}{db(y)} = \frac{\partial \pi(x)}{\partial b(y)} + \frac{\partial \pi(x)}{\partial a(x)} \frac{\partial a(x)}{\partial b(y)}$, and because the solution $a_b(x)$ is such that $\pi(x)$ always equals $f(x)$, so that $\frac{d\pi(x)}{db(y)} = 0$. Therefore, $\frac{\partial \pi(x)}{\partial b(y)} + \frac{\partial \pi(x)}{\partial a_b(x)} \frac{\partial a_b(x)}{\partial b(y)} = 0$.

We now focus on the density of women as a function of vectors b and a_b . We are going to prove there is a single vector b such that the density of women implied by the matching function is equal to $g(y)$. This is sufficient to prove the uniqueness of the equilibrium, because the choice of a_b ensures the distribution of men is $f(x)$.

We search for b such that $\pi(y) := \pi(y, \emptyset) + \sum_{x, n_Y, r \leq n_Y} \pi(x, y, n_Y, r) = g(y)$ for every y , which defines a system of equations. As stated in the introduction of this proof, we know that there is a solution, and an application of the Poincaré-Miranda theorem here confirms that. We want to prove the solution is unique. We are going to prove that there is at maximum a solution in every line of the space.

Let us choose vector $\xi \neq 0$ such that $b + \alpha \xi$, $\alpha \in \mathbb{R}$ defines a line in the vector space of b (recall that this is a space where each dimension is a distinct type of woman $y = y_w$). We prove below that for every such line, only one point can potentially solve $\pi(y) = g(y)$ for every y : there is a single point solving the linear equation $\sum_y \xi_y \pi(y) = \sum_y \xi_y g(y)$.

For each x , we define:

$$\mathbb{E} \left(\sum_{y \in Y} \xi_y |x, n_Y \right) := \sum_{r=1}^{n_Y} \sum_y \xi_y \pi(y|x, n_Y, r) \quad (31)$$

$$\mathbb{E} \left(\sum_{y \in Y} \xi_y |x \right) := \sum_{n_Y > 0} \frac{\pi(x, n_Y)}{\pi(x)} \mathbb{E} \left(\sum_{y \in Y} \xi_y |x, n_Y \right) \quad (32)$$

This allows us to base our computations on a sum over the x :

$$\sum_y \xi_y \pi(y) = \sum_y \xi_y \pi(y, \emptyset) + \sum_x \pi(x) \mathbb{E} \left(\sum_{y \in Y} \xi_y | x \right) \quad (33)$$

We need to compute the derivatives of the two terms in (33):

$$\begin{aligned} \text{grad} \left(\sum_y \xi_y \pi(y, \emptyset) \right) \cdot \xi &= - \sum_y \xi_y^2 \pi(y, \emptyset) < 0 \\ \text{grad} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x \right) \cdot \xi &= \sum_{n_Y > 0} \frac{\text{grad} \pi(x, n_Y) \cdot \xi}{\pi(x)} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \\ &\quad + \sum_{n_Y > 0} \frac{\pi(x, n_Y)}{\pi(x)} \text{grad} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \cdot \xi \end{aligned}$$

where there is no need to compute the derivative of $\pi(x) = f(x)$ as long as $a(x) = a_b(x)$.

The first term equality above is obviously negative, and we are going to see why the second term is negative. Firstly, we show that $\text{grad} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \cdot \xi < 0$

From (27), we compute $\pi(y|x, n_Y, r)$:

$$\begin{aligned} \pi(y|x, n_Y, r) &= \frac{\exp \left(\frac{\Phi(x, y, n_Y, r) - b(y)}{2} \right)}{\sum_{y'} \exp \left(\frac{\Phi(x, y', n_Y, r) - b(y')}{2} \right)} \\ \text{grad} \pi(y|x, n_Y, r) \cdot \xi &= -\pi(y|x, n_Y, r) \frac{\xi_y}{2} + \sum_{y'} \pi(y|x, n_Y, r) \pi(y'|x, n_Y, r) \frac{\xi_{y'}}{2} \\ &\quad - \frac{\pi(y|x, n_Y, r)}{2} [\mathbb{E}(\xi|x, n_Y, r) - \xi_y] \end{aligned}$$

where $\mathbb{E}(\xi|x, n_Y, r) := \sum_{y'} \pi(y'|x, n_Y, r) \xi_{y'}$. This gives the gradient of (31):

$$\begin{aligned} \text{grad} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \cdot \xi &= \sum_{r=1}^{n_Y} \sum_y \xi_y \text{grad} \pi(y|x, n_Y, r) \cdot \xi \\ &= \sum_{r=1}^{n_Y} \sum_y \frac{\pi(y|x, n_Y, r)}{2} [\xi_y \mathbb{E}(\xi|x, n_Y, r) - \xi_y \xi_y] \\ &= -\frac{1}{2} \sum_{r=1}^{n_Y} [(\mathbb{E}(\xi|x, n_Y, r))^2 - \mathbb{E}(\xi^2|x, n_Y, r)] \\ &= -\frac{1}{2} \sum_{r=1}^{n_Y} \text{Var}(\xi|x, n_Y, r) < 0 \end{aligned}$$

We are now going to compute $\sum_{n_Y > 0} \frac{\text{grad}\pi(x, n_Y) \cdot \xi}{\pi(x)} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right)$ and prove it is also negative. After that, all the terms of the derivative of (33) will be proved negative.

$$\begin{aligned} \text{grad}\pi(x, n_Y) \cdot \xi &= -\frac{\pi(x, n_Y)}{1 + n_Y} \text{grada}(x) \cdot \xi - \sum_{r'=1}^{n_Y} \frac{1}{n_Y + 1} \sum_y \xi_y \pi(y | x, n_Y, r) \pi(x, n_Y) \\ &= \frac{\pi(x, n_Y)}{1 + n_Y} \left[-\text{grada}(x) \cdot \xi - \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \right] \end{aligned}$$

Hence we need to compute $\text{grada}(x) \cdot \xi$:

$$\begin{aligned} \text{grada}_b(x) \cdot \xi &= \sum_y \xi_y \frac{\partial a_b(x)}{\partial b(y)} = -\frac{\sum_{n_Y > 0} \frac{\pi(x, n_Y)}{1 + n_Y} \sum_{r=1}^{n_Y} \sum_y \xi_y \pi(y | x, n_Y, r)}{\sum_{n_Y \geq 0} \frac{\pi(x, n_Y)}{1 + n_Y}} \\ &= -\sum_{n_Y > 0} w_{xn_Y} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \\ \text{grad}\pi(x, n_Y) \cdot \xi &= \frac{\pi(x, n_Y)}{1 + n_Y} \left[\sum_{n_Y > 0} w_{xn_Y} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) - \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \right] \end{aligned}$$

where $w_{xn_Y} = \frac{\frac{\pi(x, n_Y)}{1 + n_Y}}{\sum_{n_Y > 0} \frac{\pi(x, n_Y)}{1 + n_Y}}$ is the weight of the term in n_Y in the average share of men in the households for men x .

We can now compute the complete term:

$$\begin{aligned} &\sum_{n_Y > 0} \frac{\text{grad}\pi(x, n_Y) \cdot \xi}{\pi(x)} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \\ &= \sum_{n_Y > 0} \left(\frac{\pi(x, n_Y)}{\pi(x)(1 + n_Y)} \right) \left[\left(\sum_{n_Y > 0} w_{xn_Y} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \right)^2 - \sum_{n_Y > 0} w_{xn_Y} \left(\mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \right)^2 \right] \\ &= -\sum_{n_Y > 0} \left(\frac{\pi(x, n_Y)}{\pi(x)(1 + n_Y)} \right) \left[\sum_{n_Y > 0} w_{xn_Y} \left(\mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n_Y \right) \right) - \sum_{n'_Y > 0} w_{xn'_Y} \mathbb{E} \left(\sum_{y \in Y} \xi_y | x, n'_Y \right) \right]^2 < 0 \end{aligned}$$

Finally, all the terms of the derivative of (33) are negative. Hence, this proves that at maximum one point of the line $b + \alpha\xi$ is a candidate for $\pi(y) = f(y)$ for all y .⁷⁵

So for every line of the space, there is at most a single vector b following the distribution $g(y)$. Hence, given that a vector b following the distribution $g(y)$ must exist, it is unique. This vector defines a joint distribution following equations (8), (9) and (11), and

⁷⁵A study of the limits of $\pi(y | x, n_Y, r)$ when $\alpha \rightarrow \pm\infty$ proves this point such that $\sum_y \xi_y \pi(y) = \sum_y \xi_y g(y)$ always exist.

distributions of attributes $g(y)$ and $f(x)$.

The equilibrium maximizes the total surplus The equilibrium of the model is such that, for every man and every woman, the particular marital situation they are in (in a particular marriage, or single) maximizes their post-transfer utility. That is for every man m married with a set of wives of characteristics Y_m (including single men for whom $Y_m = \emptyset$):

$$U(x_m, Y_m) + \varepsilon_{mY_m} > U(x_m, Y') + \varepsilon_{mY'} \text{ for } Y' \neq Y_m \quad (34)$$

where we can include $Y' = \emptyset$ (being single). And for every woman w married with a man of type x_w in rank r_w with n_w co-spouses (including single women for whom $x_w = \emptyset$ and $n_w = 0$):

$$v(x_w, y_w, n_w, r_w) + \eta_{wx_w n_w r_w} > v(x', y_w, n', r') + \eta_{wx' n' r'} \text{ for } x' \neq x_w \text{ or } n' \neq n_w \quad (35)$$

where we can include $x' = \emptyset$ (being single). The total utility is:

$$U_{tot} = \sum_{m, Y_m \neq \emptyset} \left[\varepsilon_{mY_m} + U(x_m, Y_m) + \sum_{w \in W_m} v(x_m, y_w, n_m, r_m) + \eta_{wx_w n_w r_w} \right] + \sum_{m, y_m = \emptyset} \varepsilon_{m\emptyset} + \sum_{w, x_w = \emptyset} \eta_{w\emptyset}$$

According to (34) and (35), any modification of the equilibrium will result in all parties involved obtaining a lower post-transfer utility. For example, if a wife changes marriages, she will obtain a lower post-transfer utility according to (35). Her co-wives will also obtain a lower post-transfer utility according to (35), as their number of co-wives n_w is changing. According to (34), the husband will also obtain a lower post-transfer utility. We can take another example: if a monogamous marriage splits up and the man and woman become single, then according to (34), the man's utility will decrease, and according to (35), the woman's utility will decrease. Therefore, with transfers at their equilibrium level, the equilibrium matching maximizes total utility. But the utility transfers between spouses do not matter for total utility. Therefore, the equilibrium matching maximizes total utility.

D.2. Proof of Theorem 2

We start by writing $P(m, w)$ the probability that man m of type \mathbf{x}_m is matched with woman w of type y_w . Given the independence between the sympathy shocks, each match of type (\mathbf{x}, y) is equiprobable, which means that $P(m, w)$ is simply:

$$P(m, w) = \frac{\pi(\mathbf{x}_m, y_w)}{N_{\mathbf{x}_m} N_{y_w}} = \exp(\Phi(\mathbf{x}_m, y_w) - a_m - b_w) \quad (36)$$

where $N_{\mathbf{x}_m}$ and N_{y_w} are respectively the density of men of type \mathbf{x}_m and the density of women of type y_w . We add some flexibility in the model here, in the sense that a_m and b_w are sets of individual fixed effects, that need not be fully determined by \mathbf{x} and y .

Given that woman w is married, the probability of her husband (denoted $h(w)$) being man m is:

$$P(h(w) = m) = \frac{P(m, w)}{\sum_{m'} P(m', w)} = \frac{\exp(\Phi(\mathbf{x}_m, y_w) - a_m)}{\sum_{m'} \exp(\Phi(\mathbf{x}_{m'}, y_w) - a_{m'})}$$

And her probability of marrying man m given that she's married in a set of men S is:

$$P(h(w) = m | h(w) \in S) = \frac{\exp(\Phi(\mathbf{x}_m, y_w) - a_m)}{\sum_{m' \in S} \exp(\Phi(\mathbf{x}_{m'}, y_w) - a_{m'})} \quad (37)$$

Let us now consider a pair of couples, two women $w = 1$ and $w = 2$ whose husbands $h(1) = h_1$ and $h(2) = h_2$ are (respectively or not) $m = 1$ and $m = 2$. We are interested in the probability that the couples are $(1, 1)$ and $(2, 2)$ rather than the opposite. This probability writes:

$$P(h_1 = 1 | \{h_1, h_2\} = \{1, 2\}) = \frac{P(h_1 = 1, h_2 = 2)}{P(h_1 = 1, h_2 = 2) + P(h_1 = 2, h_2 = 1)} \quad (38)$$

To simplify notations, let's denote $\Phi_{11} = \Phi(\mathbf{x}_1, y_1)$, $\Phi_{12} = \Phi(\mathbf{x}_1, y_2)$. From equation (37)

and Bayes' rule, we have:

$$P(h_1 = 1, h_2 = 2) = \frac{\exp(\Phi_{11} - a_1)}{\sum_{m'} \exp(\Phi(\mathbf{x}_{m'}, y_1) - a_{m'})} \frac{\exp(\Phi_{22} - a_2)}{\sum_{m' \neq 1} \exp(\Phi(\mathbf{x}_{m'}, y_2) - a_{m'})}$$

Similarly, we can write $P(h_1 = 2, h_2 = 1)$. If we assume that $\frac{\sum_{m' \neq 1} \exp(\Phi(\mathbf{x}_{m'}, y_2) - a_{m'})}{\sum_{m' \neq 2} \exp(\Phi(\mathbf{x}_{m'}, y_2) - a_{m'})}$ is sufficiently close to 1 (which means that the fact that one particular man is already married hardly affects the overall probability of a woman getting married), then the probability (38) simplifies and we have:

$$\begin{aligned} P(h_1 = 1 | \{h_1, h_2\} = \{1, 2\}) &= \frac{\exp(\Phi_{11} + \Phi_{22})}{\exp(\Phi_{11} + \Phi_{22}) + \exp(\Phi_{12} + \Phi_{21})} \\ &= \frac{\exp(\Phi_{11} + \Phi_{22} - \Phi_{12} - \Phi_{21})}{1 + \exp(\Phi_{11} + \Phi_{22} - \Phi_{12} - \Phi_{21})} \end{aligned}$$