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The Stability Paradox: Why Expansion of Women's Education Has Not Delayed Early Union Formation or Childbearing in Latin America

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Despite substantial improvements in women's education, the age at which Latin American women marry (cohabit) or become mothers for the first time has barely decreased over the past four decades. We refer to this as the "stability paradox." We examine the relationship between years of schooling and transitions to first union or child, analyzing retrospective information from 50 cohorts of women born between 1940 and 1989 in 12 Latin American countries. Absolute and relative measures of schooling are compared. Data is drawn from 38 Demographic Health Surveys (DHS) conducted between 1986 and 2012 in these countries. Results show that expected postponement in family transitions due to educational expansion was offset by a rise in union formation and childbearing within strata of absolute education, but stayed approximately constant with strata of relative education. The relative measure of education retains the stratifying power of education but neutralizes any effect attached to a specific number of years of schooling and the learning skills associated with them. This is consistent with the idea that access to education in Latin America reproduces existing patterns of socioeconomic advantage, rather than creating a more equitable distribution of learning opportunities and outcomes.

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The expansion of women's education has had a dramatic impact on family dynamics worldwide and, in particular, on the decline of fertility (Caldwell 1980; Hirschman 1994; Castro 1995; Shapiro et al.

2013). The postponement of ages at first union (marriage or cohabitation) and at first child is widely documented as being one of the main ways in which education reduces fertility (Cochrane 1979; Jejeebhoy 1995; Bongaarts 2003). Education transforms women's ideals, expectations, and opportunity costs with regard to early marriage and childbearing, lessens the desire for large families, and improves access to and use of modern contraception. The relationship between education and age at first union or childbearing at the individual (micro) level is commonly mirrored at the societal (macro) level (Lloyd 2005), implying that educational expansion results in delays in union formation and childbearing at the population level. However, the experience of Latin America does not conform to this pattern (Esteve, López-Ruiz, and Spijker 2013; Bongaarts, Mensch, and Blanc 2017).

Although there have been substantial improvements in female (and male) schooling and marked declines in fertility, the age at which Latin American women marry (cohabit) or become mothers for the first time has barely changed over the past four decades (Rosero-Bixby 1996; Binstock 2010; Guzmán et al. 2006; Bozon, Gayet, and Barrientos 2009; Perez Amador 2008; Fussell and Palloni 2004; Cesare and Vignoli 2006; Rodríguez-Vignoli 2014). We refer to this as the "stability paradox," a trait shared by all countries in the region, and it suggests that the extra years of schooling gained by younger cohorts have not triggered the expected postponement effect on reproductive transitions (Lloyd 2005; United Nations 2002). In this context, we examine the relationship between years of schooling and the transitions to first union or child, analyzing retrospective information from 50 cohorts of women born between 1940 and 1989 in 12 Latin American countries. We compare absolute and relative measures of schooling; both measures should retain the stratifying effect of education, but the latter neutralizes any effect attached to additional years of schooling. We run a series of three logistic regression models on each outcome (union formation and childbearing) to examine the effect of raw cohort change (Model 1), cohort change adjusted for absolute strata of years of schooling (Model 2), and cohort change adjusted for relative strata of years of schooling (Model 3). Data are drawn from 38 Demographic and Health Surveys (DHS) conducted between 1986 and 2012 in these countries. We focus on the union and childbearing statuses of these cohorts at age 18, which is a critical threshold for signaling the onset of the postponement.

BACKGROUND

Link between education and reproductive decisions in Latin America

The link between education and fertility is one of the most robust across contemporary societies (Cochrane 1979; Hirschman 1994; Castro 1995; Jejeebhoy 1995; United Nations 2002; Shapiro 2012; Lloyd 2005; Bongaarts, Mensch, and Blanc 2017). Education reduces the quantum (intensity) and delays the tempo (calendar) of fertility. It alters the foundations of the family economy (Caldwell 1980) and redefines women's domestic and labor roles (Hoffman and Centeno 2003). The ways in which education reduces fertility include delay in union formation or childbearing, preference for smaller families, and wider access to contraception (Cochrane 1979; Bongaarts 1978). Education increases women's autonomy and enhances the chances of premarital employment (Jejeebhoy 1995), a combination of factors that tends to delay marriage and childbearing (Becker 1973; Oppenheimer 1988; Castro 1995). Highly educated women are also more likely to want fewer children because they are more self-reliant, have less need for the potential labor and domestic contribution of children, and are more aware of the cost to raise a child (Ainsworth, Beegle and Nyamete 1996; Caldwell 1980; Hirschman 1994). Finally, highly educated women have more knowledge of, access to, and motivation for using modern contraceptive methods than less educated women.

Latin America is no exception to the close link between education and reproductive transitions (Weinberger, Lloyd, and Klimas-Blac 1989). Castro (1995) attributed 33–60 percent of the fertility decline to the structural effects of the education expansion. At the individual level, this relationship has been confirmed in several studies (Heaton and Forste 1998; Heaton, Forste, and Otterstrom 2002; Castro and Juárez 1995; Bozon, Gayet, and Barrientos 2009; Bongaarts, Mensch, and Blanc 2017), and by the mid 1990s was proven to be stronger than in any developing region of the world (Castro and Juárez 1995).

Despite the influence of education in timing of marriage (and cohabitation) and childbearing at the individual level, the expected correspondence at the macro level is not observed. This sets

Latin America apart from other regions in the world (Bongaarts, Mensch, and Blanc 2017). Age at first union and first child has remained constant over the last 40 years, and only recently has shown slight signs of postponement among the highest-educated women in richer countries, such as Uruguay, Chile, and Argentina (Rosero-Bixby, Martín, and García 2009; Binstock 2010; Guzmán et al. 2006; Lima et al. 2016). The broken micro-macro link, which we have termed the "stability paradox," has been examined only tangentially in the literature. Some researchers came up against the paradox when they realized that trends over time in age at first child were not showing any signs of delay (Castro and Juárez 1995; Heaton and Forste 1998; Bongaarts, Mensch, and Blanc 2017). The reason for this is that changes (rejuvenation of family transitions) in rates within educational strata have offset compositional changes due to improvements in educational attainment (Esteve, López-Ruiz, and Spijker 2013; Bongaarts, Mensch, and Blanc 2017).

The stability paradox lacks a thorough theoretical explanation. Fussell and Palloni (2004) suggested that women's ability to reduce fertility without delaying union formation was possible thanks to the "stabilizing institution" of marriage during periods of rapid social change, as in the turbulent years of the 1980s and the early 1990s. However, even in those decades, marriage has undergone severe deinstitutionalization because of the rise in unmarried cohabitation, female-headed households, and single motherhood (Esteve, Lesthaeghe, and López-Gay 2012a; Laplante et al. 2015; Liu, Esteve, and Treviño 2016). Moreover, although the region's major economies revived in the 1990s, women's age at union formation and childbearing remained as precocious as it had been over the previous two decades.

Other authors stress that the persistence of early motherhood is a result of the rise and rejuvenation of sexual activity (Rodriguez-Vignoli 2014), which has loosened the bond between sexuality and union formation (Bozon, Gayet, and Barrientos 2009: Quilodrán and Castro 2009; Juárez and Gayet 2014) in a context where use of contraception is not widespread. Nevertheless, evidence shows that access to and use of contraception has increased dramatically over recent decades (UN 2015). Our own analysis of DHS data reveals that the spread of contraceptive use among young women goes beyond the expansion of schooling (results available upon request). The effect of education on contraceptive use is noticeable even among women with very low levels of schooling. The gap between primary and no schooling already reflects a 10–20 percent higher use of contraception (Bozon, Gayet, and Barrientos 2009).

Expansion of education in Latin America: Quantity versus quality

A potential explanation for the stability paradox is that expanded schooling or enrollment rates may not always have improved the learning profiles and skills of students. Latin American governments have made extraordinary, widely successful efforts to universalize access to primary education and to extend this access to the secondary and tertiary levels (Glewwe and Kremer 2006). According to the World Bank (2016), net enrollment rates rose from 81 percent in 1970 to 95 percent in 2010 at the primary level, from 53 percent in 1986 to 74 percent in 2010 at the secondary level, and from 6 percent in 1970 to 40 percent in 2010 at the tertiary level. These rates are higher than in any other middle- or low-income region of the world, and not far from those observed among countries that belong to the Organisation for Economic Co-operation and Development (OECD). On the other hand, whereas the expansion of education in the Western world took place in a context of “old” and “well-established” educational systems, with national coverage and following an “escalated” design, in Latin America it has been a relatively recent process (López Segrera 2009). It has been especially fast-paced at the tertiary level, in a context of educational systems that have not yet reached national coverage (Ferreyra et al. 2017; Gregorutti and Delgado 2015; Gregorutti et al. 2016; Ossenbach 2000). This has led to differentiation and stratification of educational opportunities (Brunner 1993; Chiroleu 2013; Dubet 2011; Marteleto et al. 2011; Mizala and Torche 2012; Wakeling and Savage 2015).

Regarding quality of education, Latin American countries rank among the lowest performing of the 68 countries that participated in the 2012 Programme for International Student Assessment (PISA). The PISA test for Latin America also revealed a lack of equity within educational strata. Socially and economically privileged families are able to provide higher-quality education for their children than less privileged ones. According to the OECD, “schools tend to reproduce existing patterns of socioeconomic advantage, rather than create a more equitable distribution of learning opportunities and outcomes" (OECD 2013: p. 104). Latin America shows a polarized educational system that does not offer the same educational opportunities to people from less privileged backgrounds by comparison with the upper classes, whose opportunities are enhanced through private institutions (Hoffman and Centeno 2003). The strong mediating role of education creates a situation of "inherited meritocracy," which is legitimized through access to education (Torche 2014).

On this account, it is arguable whether the growth of the educational offer has meant democratization of educational opportunities or just a change in the pattern of intergenerational transmission of social (dis)advantage. To shed light on this question, we contrast two alternative measures of schooling: one based on absolute years of schooling, thus assuming the transformative power of education, and one based on relative years of schooling within cohorts, which retains the stratifying dimension of education but neutralizes the effect of the extra years at school.

Research aims, questions, and hypotheses

This article documents the stability paradox. We adopt a long cohort perspective embracing the 50 cohorts of women born between 1940 and 1989 in 12 Latin American countries. We focus on the union formation and childbearing status of these women at age 18, which is a critical age for observation of delays in family transition. These two transitions are closely connected (Rodriguez-Vignoli 2014) and we therefore expect similar results. The broad temporal and spatial perspectives used in this analysis allow inquiry into how the stability paradox has unfolded over time as well as an update of trends in union formation and childbearing vis-à-vis previously published research. Consistent with earlier findings, we expect a strong positive relationship between absolute and relative strata of years of schooling and age at first union or childbearing at the individual level in every cohort of women. We also expect that if stability is to occur, rate changes on union formation and first childbearing within absolute strata of education must offset the expected effects due to compositional change.

We hypothesize that a relative measure of education ranking each woman's education within the overall educational performance of her cohort will be more consistent with the pattern of stability. In other words, the least educated women in each cohort, regardless of the number of years of schooling attained, will marry (or cohabit) or become mothers at similar ages. The relative measure of education should retain the stratifying power of education but neutralize any effect attached to a specific number of years of schooling and the learning skills associated with them. This would be consistent with the idea that access to education reproduces existing patterns of socioeconomic advantage, rather than creating a more equitable distribution of learning opportunities and outcomes.

DATA AND METHOD

We use data from 38 Demographic Health Surveys carried out in 12 Latin American countries between 1986 and 2012 (see Table 1 for characteristics). For Brazil 2006, we included the *Pesquisa Nacional de Demografia e Saúde*. DHS data have been widely used to study transitions to adulthood in developing countries as they have the advantages of using a standardized questionnaire, providing multiple rounds of data, and asking retrospective questions on union formation and childbearing. We have included all the available DHS in the region except, due to data limitations, El Salvador (1985) and Trinidad and Tobago (1987). DHS are nationally representative surveys of women aged 15 to 49, with sample sizes varying from 4,000 to 53,000 women. From these samples, we selected women aged 20 to 49 born between 1940 and 1989. Birth cohorts were grouped in ten-year intervals: 1940–49, 1950–59, 1960–69, 1970–79, and 1980–89. To maximize the size of the sample and the number of cohorts it includes, we pooled all the DHS data available in every country, thus for some cohorts, there is more than one observation point in time.

Demographic and Health Surveys include retrospective questions on age at first union (marriage or cohabitation) and first child. Two dummy variables were created to distinguish between women who had experienced these transitions by age 18 and those who had not. Alternative age thresholds (e.g., 20 and 22) yielded very similar results because the outcome variable (percent of mothers at a given age) was highly correlated between ages. The percentage of women who had, at some point, been in union or a mother increases steadily with age and there is no evidence of a bimodal distribution of the age at union formation or childbearing in our group of countries.

Retrospective questions in cross-sectional surveys are subject to selection and recall biases because potential respondents who died or migrated before the date of the survey are not included and some may not recall precisely some key dates from the past. However, it is unlikely that the age stability paradox would arise from such potential biases. First, stability in age at first union or first child has been corroborated by other statistical sources that are not subject to this limitation, as happens, for example, in the analysis of population censuses (Esteve, López-Ruiz, and Spijker 2013). Second, women of the same cohorts offer consistent responses when two or more DHSs of the same country are compared over time. As a precautionary measure, we controlled for respondent's age and the results indicate that there are no inconsistencies regarding transitions to first union or to first child.

Concerning education, we constructed an absolute and a relative measure, both based on years of schooling. In the absolute measure, years of schooling were grouped into five categories: "5 or less," "6 to 8," "9 to 12," and "13+". In the relative one, we classified each cohort of women into four categories, "Low," "Medium-low," "Medium-high," and "High." To set the thresholds for these categories, we calculated quartiles Q_1, Q_2, and Q_3 of years of schooling for each country and cohort. The women who fall into the "Low" group always represent the 25 percent of their cohort who are least educated. Women with "High" education represent the 25 percent most educated. This turns out to be the most heterogeneous group because, depending on the country, it may include women with quite different levels of educational attainment. Aware of this limitation, we mainly focus on the Low and Medium educational groups.

To examine cohort changes in union formation and first childbearing, we rely on logistic regression analysis. We run a series of three logistic regression models on each outcome. The main focus is on changes across cohorts with and without adjustments for years of schooling. Model 1 examines the effect of raw (unadjusted) cohort change. In Model 2, cohort differences are adjusted for absolute years of schooling to show the cohort change that would have occurred if rates of union formation and first childbearing remained constant over time within absolute strata of education. In Model 3, cohort differences are adjusted for relative years of schooling to show the cohort change would have occurred if within-stratum rates increased over time. Consistent with our main hypothesis, we expect that the odds ratios on cohort change from Model 3 would be more similar to

Model 1 (unadjusted change) than those from Model 2 would be to Model 1. Additionally, all models control for respondent's age and rural-urban residence.

RESULTS

Trends in age at first union and first child

The two panels of Figure 1 provide information on the percentage of women who had experienced first union and first child by age 18. The stability across cohorts becomes obvious at first glance in each of the two transitions. To give some examples, the percentage of mothers at 18 was 28 percent among Colombian women born in the 1940s and 29 percent among women born in the 1980s. A similar pattern holds for countries in Central America (e.g., Guatemala and Nicaragua), the Caribbean (e.g., Dominican Republic and Haiti), and the Andean region (e.g., Bolivia, Ecuador, and Peru). The flatness of the trend is shared by all countries despite differences in levels between them. Women in Central America and the Caribbean (but not in Haiti) are the most precocious. In these regions, the percentage of women who had, at some point, been in union and even mothers by age 18 is close to 50 percent, whereas in the other countries it is around or below 35 percent. Cohort trends remain equally stable in countries with early family transitions as they do in those with later family transitions.

Expansion of education

The two panels of Figure 2 depict an extraordinary rise in education in Latin America using the two alternative indicators of education. The first panel shows the percentage of women with nine or more years of schooling by birth cohort and country. The second panel represents the median years of schooling, which is to say the number of years of schooling separating the lower half from the upper half within each cohort and country. Both indicators offer analogous views of the expansion of education. The percentage of women with nine or more years of schooling has doubled and, in most cases, tripled between the oldest and youngest birth cohorts. The median years of schooling has likewise risen: On average, recent cohorts have five years of schooling more than the 1940s cohort.

Transitions to first union and first child by years of schooling

Transitions to first union and first child are clearly stratified by years of schooling. In Figures 3 and 4 we show, by years of schooling and cohort, the percentage of women who had their first union and first child by age 18. We selected Colombia, Peru, and the Dominican Republic and show trends by absolute and relative years of schooling. In both indicators, women with high levels of education experience family transitions at later ages than women with low levels of education. The educational gradient exists for all events, cohorts, and countries (even for those not shown in Figures 3 and 4). Cohort percentages by absolute years of schooling have increased in the two events, in particular among women with secondary or less education. This increase indicates that women with analogous years of schooling but born in the 1980s were forming unions and having children earlier in life than women born a few decades earlier. When women are classified according to relative years of schooling, cohort trends are more stable (see Figure 4). There is little evidence of any drop in age across educational groups and, in the case of Peru, there are hardly any differences between the oldest and youngest cohorts. Hence, the least-educated women in every cohort are experiencing transitions to first union and first child at similar ages. The exception is the highly educated group, which is the most heterogeneous in composition because it includes the widest range of years at school.

Logistic regression models confirm the main findings reached to this point. Table 2 shows the odds ratios of having experienced first union or first child by the age of 18 between the youngest and oldest cohort in each country. These ratios come from logistic regression which, in addition to birth cohort, controls for the respondent's age, years of schooling, and region (rural-urban) (complete models are shown in Appendixes 1 and 2[1]). The rural-urban effect disappears when educational attainment is taken into consideration. Values above 1 indicate that the odds of having experienced the transition by age 18 are higher in the youngest cohort than in the oldest. There is, therefore, a rejuvenation of the event. Values below 1 indicate the contrary, which is postponement, and values not significantly different from 1 indicate stability.

As mentioned in the Methods section, Model 1 shows the raw (unadjusted) cohort differences. Hence, this model does not take into account any measure of educational attainment.

[1] Appendixes are available at the supporting information tab at wileyonlinelibrary.com/journal/sfp.

The odds ratios between the earliest and latest birth cohorts do not differ substantially from 1 in 8 of the 12 countries. In three countries, Haiti, Peru, and Honduras, the likelihood of experiencing first union and first child by age 18 among the latest cohort was slightly lower than among the earliest one. Brazil was the only country in which women born in the 1980s were more likely to have experienced first union and first child by age 18 than women born in the 1940s. In Model 2, cohort changes are adjusted for absolute years of schooling. As a result, cohorts' differences are amplified. This model shows the cohort differences that would have occurred if rates of union formation and first childbearing remained constant within absolute strata of education. Controlling for absolute strata of years of schooling, women born in the 1980s are more likely to have experienced first union and first child by age 18 than women born in the 1940s. In this case, odds ratios on cohorts are measuring the overall offsetting effect or rate change within educational strata that would have had to take place to avoid the expected postponement due to expansion of education. In Model 3, cohort change is adjusted for relative strata of education. This shows the cohort differences that would have pertained if the within-stratum rates of union formation and first childbearing increased over time, while the educational gradients remained more or less constant. The odds ratios on cohorts from Model 3 are more similar to the unadjusted odds ratios from Model 1 than those from Model 2 are to Model 1.

In Models 2 and 3, statistically significant differences exist in union formation and childbearing across educational groups. The higher the absolute or relative level of education, the lower the probability of a woman's having been in union or a mother by age 18. The correlation between years of schooling and transitions to first union and first child is noticeable from the lowest to the highest levels (results shown in Appendix 1).

CONCLUSIONS

Latin American societies have undergone many social, demographic, and economic changes over recent decades, but age at first union and first child has remained unchanged. Although, at the individual level, women's years of schooling are strongly correlated with age at first union and first child, the striking expansion of schooling that has occurred over the past few decades has hardly had any impact on women's transitions to first marriage/cohabitation or motherhood: What we refer to as the stability paradox.

We have documented the stability paradox for 12 Latin American countries among women born between 1940 and 1989. In all of these countries, the percentage of women who experienced first union and first child remained stable across cohorts. Transitions to adulthood were clearly stratified by years of schooling, but during the process of educational expansion women were forming unions and becoming mothers at earlier ages, in particular women with secondary education or less. Hence, consistent with earlier findings (Esteve, López-Ruiz, and Spijker 2013; Bongaarts, Mensch, and Blanc 2017), the expected postponement in family transitions due to educational expansion was wholly offset by a rejuvenation of union formation and childbearing within educational groups.

We hypothesized that a relative measure of education ranking each woman's schooling within the overall performance of her cohort would be more consistent with the pattern of stability. Our results suggest that women with low to medium levels of schooling had very similar calendars across cohorts, regardless of the number of years of schooling attained. The least educated women in all cohorts were forming unions and having children at similar ages. The relative measure of education had the same stratifying power as the absolute one but showed lower levels of variation across cohorts. This implies that the extra years of schooling gained by Latin American women have not triggered the postponement mechanisms as would seem to have happened in other societies, not only in the western world but also in East Asian countries. The lack of correspondence between the micro and macro dimensions of schooling deserves further attention. We suggest two avenues of future research.

First, researchers should investigate whether expanded schooling has contributed to higher learning skills. Several reports have identified Latin American countries as low performing in terms of education and, moreover, have revealed a lack of equity in educational opportunities (OECD 2013: p.104). The quality of education may have confounded (or neutralized) the cause-effect relationship between education and postponement of marriage and childbearing. One way of investigating that is to compare the reproductive decisions of women who have similar amounts of schooling but who attend schools of different quality. Even better would be direct measures of learning skills and their association with early union formation and childbearing. This would almost certainly be a challenging task given the high data requirements of this kind of research.

Second, future research should consider the interaction between education and the labor market, and the opportunities that Latin American societies offer by way of capitalizing on individual investments in education in terms of work opportunities and formal, stable jobs. The situation may be similar to that argued by Grant (2015) in the context of Malawi, which is to say that given the lack of meaningful employment in Latin America, women relegated to the informal sector may not have perceived the benefits of postponing childbearing. Women's (and men's) demographic behavior will not change until education offers opportunities for higher-paying jobs (Hoffman and Centeno 2003) and inequality will therefore continue to shape the life transitions of young adults, resulting in very diverse paths strongly stratified by socioeconomic status and class (Juárez and Gayet 2014). Additionally, despite recent gains in female labor market participation (Novta and Wong 2017), the region still exhibits marked gender differences (Contreras 2002; Duryea et al. 2007; Gamboa and Waltenberg 2012).

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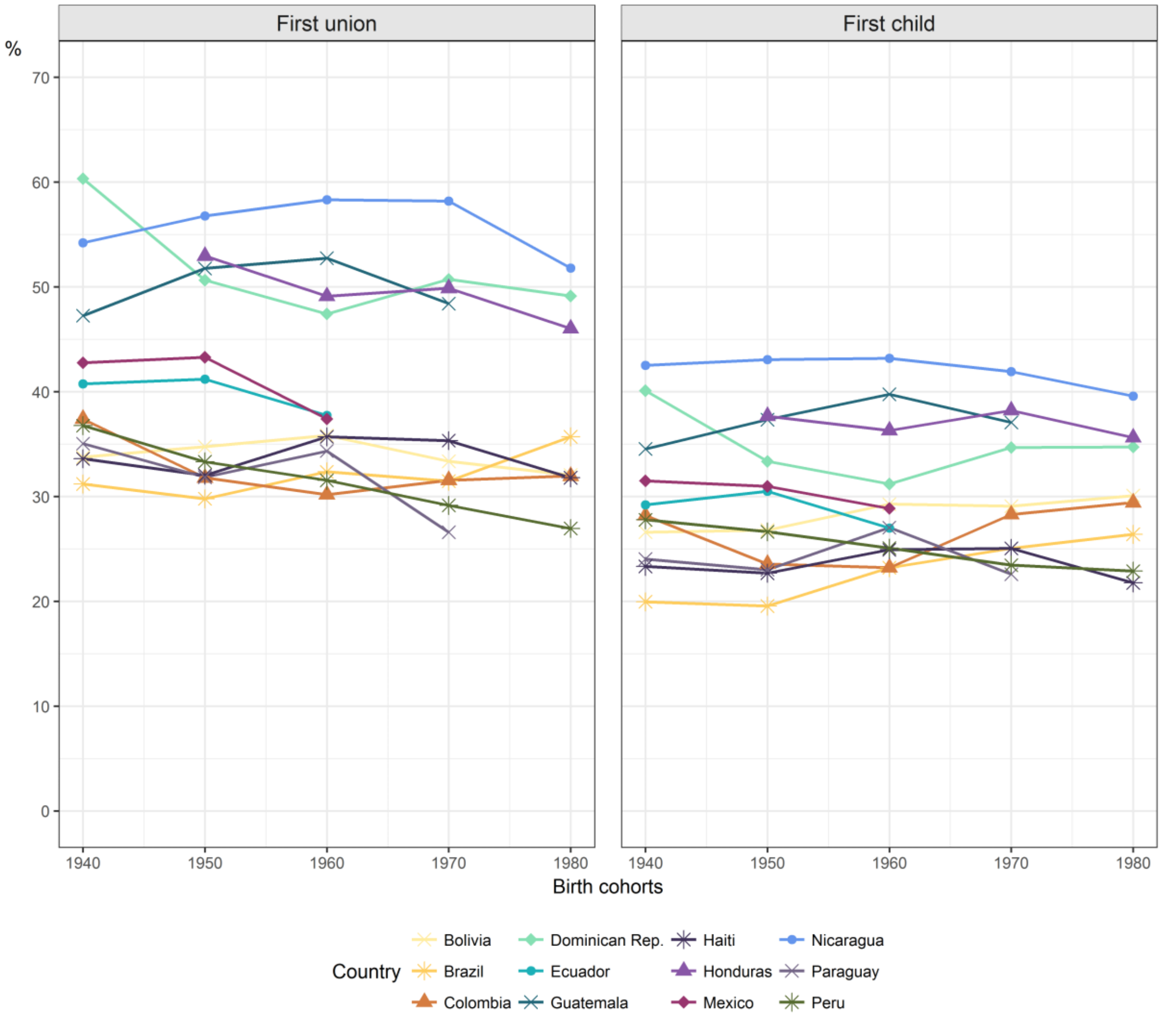
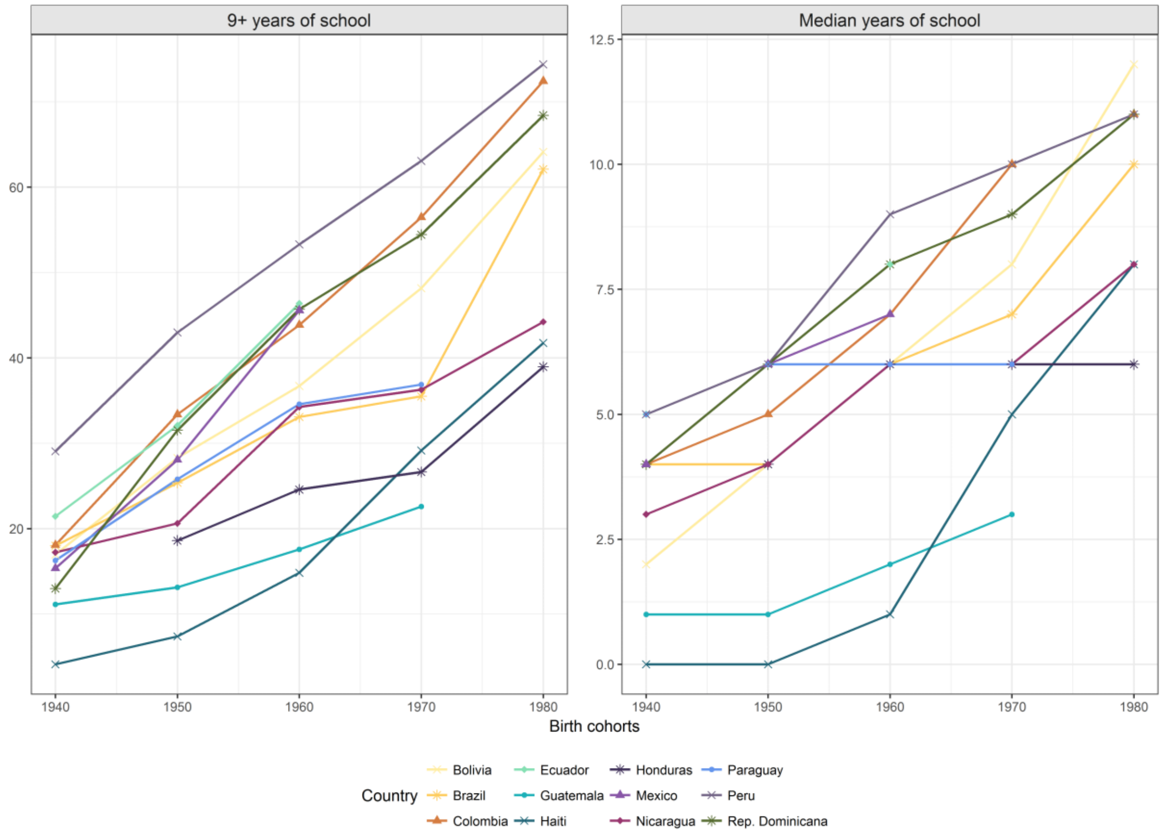
Figure 1. Cohort trends in percentages of women who experienced first union and child by age 18 in 12 Latin American countries



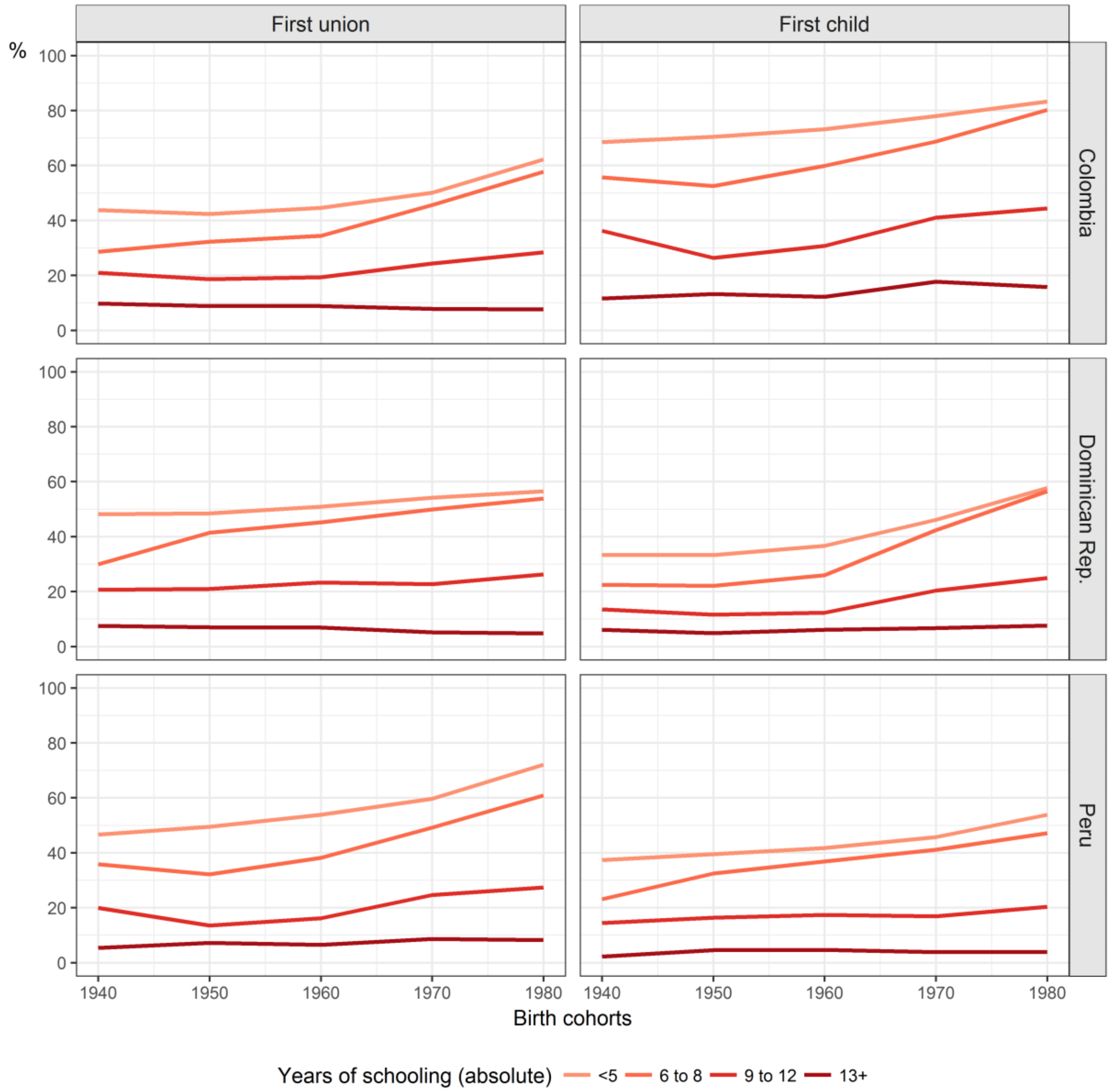
Source: DHS data.

Figure 2. Cohort trends in women's education in 12 Latin American countries using absolute and relative measures of years of schooling



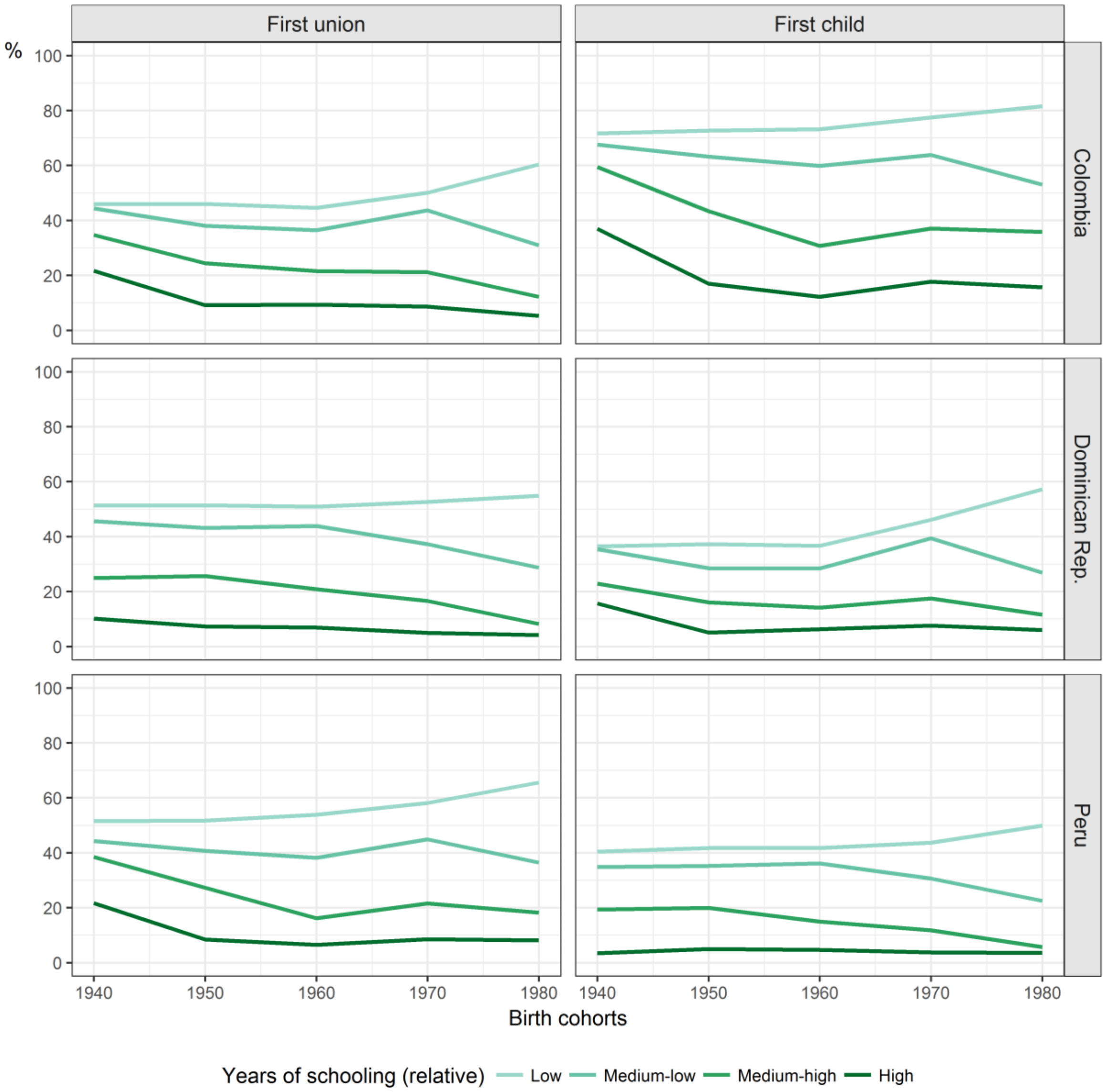
Source: DHS data.

Figure 3. Cohort trends in percentages of women who, by absolute years of schooling, experienced first union and child by age 18 in 3 selected Latin American countries



Source: DHS data.

Figure 4. Cohort trends in percentages of women who, by relative years of schooling, experienced first union and child by age 18 in 3 selected Latin American countries



Source: DHS data.

Table 1 Characteristics of samples included in analysis

Country	Survey year						Birth Cohorts	N
	1985–1989	1990–1994	1995–1999	2000–2004	2005–2009	2010–2012		
Bolivia	1989	1993–94	1998	2003–04	2008	—	1940–89	48,879
Brazil	1986	—	1996	—	2006	—	1940–89	27,732
Colombia	1986	1990	1995	2000	2004–05	2009–10	1940–89	100,176
Dominican Rep.	1986	1991	1996	2002	2007	—	1940–89	58,087
Ecuador	1987	—	—	—	—	—	1940–69	3,535
Guatemala	1987	—	1995	—	—	—	1940–79	13,448
Haiti	—	1994–95	—	2000	2005–06	2012	1940–89	29,527
Honduras	—	—	—	—	2005–06	2011–12	1950–89	31,378
Mexico	1987	—	—	—	—	—	1940–69	6,754
Nicaragua	—	—	1998	2001	—	—	1940–89	20,246
Paraguay	—	1990	—	—	—	—	1940–69	4,565
Peru	1986	1991–92	1996	2000	2003–08	2012	1940–89	111,793
Total Surveys/N	8	6	7	6	7	4		456,120

TABLE 2 Selected odds ratios from a logistic regression model of women's first sexual intercourse, union and child by age 18 in 12 Latin American countries

Country & Birth cohorts	First union			First child		
	M1	M2	M3	M1	M2	M3
Bolivia (1980s vs. 1940s)	0.9	1.5**	0.7**	1.2*	1.8**	0.9
Brazil (1980s vs. 1940s)	1.7**	3.9**	2.1**	1.6**	3.9**	2.0*
Colombia (1980s vs. 1940s)	0.9**	2.1**	0.7**	1.2*	2.9**	1.0
Dominican Rep. (1980s vs. 1940s)	0.9**	2.8**	0.7**	0.9	2.8**	0.9**
Ecuador (1960s vs. 1940s)	1.0	1.1	0.6*	0.9	1.0	0.5**
Guatemala (1970s vs. 1940s)	0.9	1.2	0.9	0.9	1.2	1.0
Honduras (1980s vs. 1950s)	0.7**	0.7**	0.5**	0.8	1.0	0.7**
Haiti (1980s vs. 1940s)	0.7**	1.1	0.4**	0.8	1.5**	0.6**
Mexico (1960s vs. 1940s)	0.8	0.8	0.5**	0.9	1.1	0.6**
Nicaragua (1980s vs. 1940s)	1.0	1.5*	0.7	0.9	1.3	0.7*
Peru (1980s vs. 1940s)	0.7**	1.4**	0.6**	0.8**	1.7**	0.8**
Paraguay (1970s vs. 1940s)	0.9	0.9	0.7	1.1	1.0	0.8

*Significant at < 0.05; **< 0.01.

SOURCE: DHS data.

Appendix 1. Estimated odds ratios from a logistic regression model of women’s first union by age 18 in 12 Latin American countries

	Bolivia			Brazil			Colombia			Dominican Rep.			Ecuador			Guatemala		
	M1	M2	M3	M1	M2(1)	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3
Birth Cohorts																		
1940	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
1950	1,0	1,1 *	1,0	1,1	1,3 **	1,3 **	0,8 **	1,0	0,7 **	0,7 **	1,0	0,6 **	1,0	1,0	0,7	1,0	1,0	1,0
1960	1,1	1,2 **	1,0	1,3 **	1,8 **	1,4 **	0,8 **	1,1	0,7 **	0,7 **	1,2 **	0,5 **	1,0	1,1	0,6 *	0,9	1,1	1,0
1970	1,0	1,2 **	0,8 **	1,5 **	2,5 **	1,9 **	0,8 **	1,4 **	0,8 **	0,8 **	1,9 **	0,7 **				0,9	1,2	0,9
1980	0,9	1,5 **	0,7 **	1,7 **	3,9 **	2,1 **	0,9 *	2,1 **	0,7 **	0,9 *	2,8 **	0,7 **						
Respondent's age																		
20-24	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
25-29	1,1 *	1,1 *	1,1 *	1,2 **	1,3 **	1,3 **	1,1 **	1,1 **	1,1 **	1,2 **	1,3 **	1,3 **	1,5 **	1,5 **	1,5 **	1,0	1,0	1,0
30-34	1,1 **	1,1 **	1,1 *	1,2 **	1,4 **	1,4 **	1,1 **	1,2 **	1,2 **	1,3 **	1,3 **	1,3 **	1,3	1,2	1,2	1,2 **	1,3 **	1,2 **
35-39	1,0	1,0	1,0	1,4 **	1,5 **	1,5 **	1,1 **	1,2 **	1,2 **	1,3 **	1,4 **	1,4 **	1,4	1,1	1,1	1,2	1,2	1,1
40-44	1,0	1,0	0,9	1,3 **	1,5 **	1,5 **	1,2 **	1,2 **	1,2 **	1,4 **	1,4 **	1,4 **	1,4	1,0	1,0	0,9	0,9	0,8
45-49	1,0	0,9 *	0,9 **	1,4 **	1,7 **	1,7 **	1,1 **	1,2 **	1,1 **	1,4 **	1,3 **	1,3 **	1,1	0,7	0,6	0,7 **	0,8 *	0,7 **
Region																		
Urban	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
Rural	1,4 **	0,9 **	0,9 **	1,6 **	1,1	1,0	1,9 **	1,1 **	1,0 *	1,7 **	1,1 **	1,1 **	1,5 **	0,8 **	0,8 *	2,1 **	1,3 **	1,3 **
Years of school																		
< 5		ref.			ref.			ref.			ref.			ref.			ref.	
6 to 8		1,0			0,6 **			0,8 **			0,6 **			0,7 **			0,4 **	
9 to 12		0,4 **			0,3 **			0,3 **			0,2 **			0,3 **			0,2 **	
13 +		0,1 **			0,1 **			0,1 **			0,1 **			0,1 **			0,1 *	
Quartiles of years of school																		
Low			ref.			ref.			ref.			ref.			ref.			ref.
Meium-low			0,9 **			0,7 **			0,6 **			0,5 **			0,7 **			0,9 *
Medium-high			0,6 **			0,3 **			0,3 **			0,2 **			0,4 **			0,6 **
High			0,2 **			0,1 **			0,1 **			0,1 **			0,2 **			0,2 **
Intercept	0,4 **	0,7 **	0,9	0,3 **	0,4 **	0,5 **	0,4 **	0,6 **	1,2 **	0,9 *	1,6 **	3,8 **	0,4 **	1,1	1,6	0,7 **	1,2	1,5 **

** sig<0.01 *sig<0.05

Source: DHS data.

Appendix 1. (continued) Estimated odds ratios from a logistic regression model of women’s first union by age 18 in 12 Latin American countries

	Honduras			Haiti			Mexico			Nicaragua			Peru			Paraguay		
	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3
Birth Cohorts																		
1940				ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
1950	ref.	ref.	ref.	0,8	0,8	0,8	0,9	0,9	0,5 **	1,0	1,1	1,0	0,8 **	1,0	0,9 *	1,1	1,0	0,8
1960	0,9 *	0,9	0,8 **	0,9	1,0	0,7 **	0,8	0,8	0,5 **	1,2	1,6 **	1,1	0,8 **	1,1 **	0,8 **	1,2	1,2	0,8
1970	0,8 **	0,8 *	0,7 **	0,8 *	1,0	0,6 **				1,3	1,7 **	1,1	0,7 **	1,2 **	0,7 **	0,9	0,8	0,7
1980	0,7 **	0,7 **	0,5 **	0,7 **	1,1	0,4 **				1,0	1,5 *	0,7	0,7 **	1,4 **	0,6 **			
Respondent's age																		
20-24	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
25-29	1,1 *	1,0	1,0	1,0	1,1	1,1	1,1	1,0	1,0	1,0	1,1	1,0	1,1 **	1,2 **	1,3 **	1,1	1,1	1,1
30-34	1,0	0,9 **	0,9 **	0,9 *	0,9 *	0,9 **	1,5 **	1,1	1,1	1,0	1,0	1,0	1,1 **	1,2 **	1,3 **	1,2	1,0	1,0
35-39	1,0	0,8 **	0,8 **	0,9	0,9 **	0,8 **	0,9	0,6 **	0,6 **	1,1	1,0	1,0	1,2 **	1,2 **	1,4 **	1,0	0,8	0,8
40-44	0,9 *	0,7 **	0,7 **	0,8 **	0,7 **	0,7 **	1,0	0,6 **	0,5 **	1,4 **	1,2	1,1	1,1 **	1,2 **	1,2 **	1,3	0,9	0,9
45-49	0,9	0,6 **	0,7 **	0,7 **	0,6 **	0,5 **	1,2	0,7	0,6 *	1,1	0,9	0,8	1,1 **	1,1 **	1,1 **	1,3	0,8	0,8
Region																		
Urban	ref.	ref.	ref.				ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
Rural	1,8 **	1,0	1,0	1,5 **	1,0	0,9 **				1,8 **	0,9 **	0,9 **	2,5 **	1,1 **	1,1 **	1,9 **	1,1	1,1
Years of school																		
< 5		ref.			ref.			ref.			ref.			ref.			ref.	
6 to 8		0,6 **			0,5 **			0,5 **			0,6 **			0,8 **			0,5 **	
9 to 12		0,2 **			0,2 **			0,2 **			0,2 **			0,3 **			0,2 **	
13 +		0,1 **			0,1 **			0,1 **			0,1 **			0,1 **			0,1 **	
Quartiles of years of school																		
Low			ref.						ref.			ref.			ref.			ref.
Meium-low			0,7 **			0,6 **			0,6 **			0,8 **			0,6 **			0,6 **
Medium-high			0,4 **			0,4 **			0,2 **			0,4 **			0,2 **			0,5 **
High			0,1 **			0,1 **			0,1 **			0,1 **			0,1 **			0,1 **
Intercept	0,9	2,9 **	3,5 **	0,6 **	0,9	2,1 **	1,4 *	2,2 **	3,6 **	0,9	2,1 **	3,0 **	0,4 **	0,8 **	1,1	0,3 **	0,9	1,2

** sig<0.01 *sig<0.05

Source: DHS data.

Appendix 2. Estimated odds ratios from a logistic regression model of women's first child by age 18 in 12 Latin American countries

	Bolivia			Brasil			Colombia			Dominican Rep.			Ecuador			Guatemala		
	M1	M2	M3	M1	M2(1)	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3
Birth Cohorts																		
1940	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
1950	1,0	1,1	1,0	1,0	1,2 **	1,2 **	0,8 **	1,0	0,7 **	0,8 **	1,0	0,7 **	1,0	1,0	0,7	1,0	1,0	1,0
1960	1,1	1,2 **	1,0	1,2 **	1,7 **	1,3 **	0,8 **	1,1 **	0,7 **	0,7 **	1,2 **	0,6 **	0,9	1,0	0,5 **	1,0	1,2	1,0
1970	1,1	1,4 **	1,0	1,5 **	2,5 **	1,8 **	1,1	1,9 **	1,0	0,9 *	1,9 **	0,8 **				0,9	1,2	1,0
1980	1,2 **	1,8 **	0,9	1,6 **	3,9 **	2,0 **	1,2 **	2,9 **	1,0	1,0	2,8 **	0,9 **						
Respondent's age																		
20-24	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
25-29	1,1 **	1,1 **	1,1 **	1,2 **	1,3 **	1,4 **	1,1 **	1,1 **	1,1 **	1,2 **	1,3 **	1,3 **	1,3	1,2	1,3	1,0	1,0	1,0
30-34	1,1 **	1,1 *	1,1 *	1,1 *	1,2 **	1,2 **	1,1 **	1,2 **	1,2 **	1,2 **	1,2 **	1,2 **	1,1	0,9	1,0	1,1	1,2	1,1
35-39	1,0	1,0	1,0	1,1 *	1,2 **	1,2 **	1,1 **	1,2 **	1,2 **	1,3 **	1,2 **	1,2 **	1,1	0,9	0,9	1,0	1,0	1,0
40-44	1,0	1,0	0,9	1,1	1,2 **	1,2 **	1,1 **	1,2 **	1,2 **	1,3 **	1,2 **	1,2 **	1,1	0,8	0,8	0,8	0,8	0,8 *
45-49	0,9	0,9 *	0,9 **	1,1	1,2 **	1,3 **	1,1 **	1,1 **	1,1 **	1,2 **	1,1 **	1,1 *	0,8	0,5 *	0,4 **	0,8	0,8	0,8
Region																		
Urban	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
Rural	1,5 **	1,0 *	1,0	1,4 **	0,9 *	0,9 **	1,8 **	1,0	1,0	1,6 **	1,0	1,0	1,6 **	0,9	0,9	2,0 **	1,2 **	1,2 **
Years of school																		
< 5		ref.			ref.			ref.			ref.			ref.			ref.	
6 to 8		1,0			0,5 **			0,7 **			0,6 **			0,6 **			0,5 **	
9 to 12		0,5 **			0,2 **			0,3 **			0,2 **			0,2 **			0,2 **	
13 +		0,1 **			0,1 **			0,1 **			0,1 **			0,1 **			0,1 **	
Quartiles of years of school																		
Low			ref.			ref.			ref.			ref.			ref.			ref.
Meium-low			0,8 **			0,6 **			0,6 **			0,5 **			0,6 **			0,9 **
Medium-high			0,6 **			0,3 **			0,2 **			0,2 **			0,4 **			0,7 **
High			0,2 **			0,1 **			0,1 **			0,1 **			0,1 **			0,2 **
Intercept	0,3 **	0,5 **	0,6 **	0,2 **	0,3 **	0,4 **	0,3 **	0,4 **	0,8 **	0,4 **	0,8 **	1,6 **	0,3 **	0,9	1,2	0,4 **	0,7 **	0,8

** sig<0.01 *sig<0.05

Source: DHS data.

Appendix 2. (continued) Estimated odds ratios from a logistic regression model of women's first child by the age of 18 in 12 Latin American countries

	Honduras			Haiti			Mexico			Nicaragua			Peru			Paraguay		
	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3	M1	M2	M3
Birth Cohorts																		
1940				ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
1950	ref.	ref.	ref.	0,9	0,9	0,9	0,9	0,9	0,5 **	0,9	1,0	0,9	0,9 *	1,1 **	1,0	1,1	1,1	0,9
1960	1,0	1,0	0,9	1,0	1,1	0,8 *	1,0	1,1	0,6 **	0,9	1,1	0,8	0,8 **	1,2 **	0,9 **	1,3	1,3	0,9
1970	1,0	1,0	0,8 *	1,0	1,3 *	0,7 *				1,0	1,2	0,8	0,8 **	1,3 **	0,8 **	1,0	1,0	0,8
1980	0,8	1,0	0,7 **	0,9	1,5 **	0,6 **				0,9	1,3	0,7 *	0,8 **	1,7 **	0,8 **			
Respondent's age																		
20-24	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
25-29	1,1	1,0	1,0	1,2 **	1,3 **	1,3 **	1,3 *	1,1	1,1	1,1	1,1	1,1	1,1 **	1,2 **	1,3 **	1,0	1,0	1,0
30-34	1,0	0,9 *	0,9	1,2 **	1,2 **	1,2 **	1,6 **	1,2	1,2	1,1	1,1	1,1	1,1 **	1,2 **	1,3 **	1,0	0,8	0,8
35-39	0,9	0,8 **	0,8 **	1,3 **	1,2 **	1,2 **	1,0	0,7 *	0,7 *	1,2 **	1,1	1,1	1,1 **	1,2 **	1,3 **	0,9	0,7	0,7
40-44	0,8 **	0,7 **	0,7 **	1,1	1,0	0,9	1,2	0,7	0,7 *	1,3 **	1,1	1,0	1,1 **	1,1 **	1,2 **	1,0	0,6	0,6
45-49	0,9	0,7 **	0,7 **	1,0	0,9	0,8 **	1,3	0,7	0,7	1,1	0,8	0,8 *	1,0	1,0	1,1 *	1,2	0,8	0,7
Region																		
Urban	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.	ref.
Rural	1,7 **	1,0	1,0	1,6 **	1,0	0,9 **				1,8 **	0,9 **	0,9 **	2,4 **	1,1 **	1,1 **	2,0 **	1,1	1,1
Years of school																		
< 5		ref.			ref.			ref.			ref.			ref.			ref.	
6 to 8		0,6 **			0,4 **			0,4 **			0,5 **			0,8 **			0,5 **	
9 to 12		0,2 **			0,2 **			0,2 **			0,2 **			0,3 **			0,2 **	
13 +		0,1 **			0,1 **			0,0 **			0,1 **			0,1 **			0,0 **	
Quartiles of years of school																		
Low			ref.			ref.			ref.			ref.			ref.			ref.
Meium-low			0,6 **			0,7 **			0,5 **			0,7 **			0,6 **			0,5 **
Medium-high			0,4 **			0,3 **			0,2 **			0,4 **			0,2 **			0,4 **
High			0,1 **			0,1 **			0,1 **			0,1 **			0,1 **			0,1 **
Intercept	0,5 **	1,4 **	1,7 **	0,2 **	0,4 **	0,8	0,7 *	1,0	1,8 **	0,6 **	1,3	2,1 **	0,3 **	0,5 **	0,7 **	0,2 **	0,6	0,8

** sig<0.01 *sig<0.05

Source: DHS data.