HHS Public Access
Author manuscript
Popul Res Policy Rev. Author manuscript; available in PMC 2016 June 01.

Published in final edited form as:
Popul Res Policy Rev. 2015 June 1; 34(3): 417–435. doi:10.1007/s11113-014-9351-8.

Educational Attainment and Timing to First Union across Three Generations of Mexican Women

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Abstract

We use data from Wave 3 of the Mexican Family Life Survey (N = 7276) and discrete-time regression analyses to evaluate changes in the association between educational attainment and timing to first union across three generations of women in Mexico, including a mature cohort (born between 1930 and 1949), a middle cohort (born between 1950 and 1969), and a young cohort (born between 1970 and 1979). Mirroring prior research, we find a curvilinear pattern between educational attainment and timing to first union for women born between 1930 and 1969, such that once we account for the delaying effect of school enrollment, those with the lowest (0–5 years) and highest levels of education (13+ years) are characterized by the earliest transition to a first union. For women born between 1970 and 1979, however, we find that the pattern between education and first union formation has changed. In contrast to their peers born in earlier cohorts, highly educated women in Mexico are now postponing first union formation relative to the least educated. We draw on competing theories of educational attainment and timing to first union to help clarify these patterns in the context of Mexico.

Keywords

union formation; Mexico; educational attainment; marriage

Introduction

For women in Mexico, marriage is nearly universal and occurs early in the adult life course (Echarri Cánovas and Pérez Amador 2007). While the average age at first union has risen slightly, it remains remarkably young, particularly for women. Over a 40-year period (women born in 1940–49 compared to women born 1975–1979), the median age at first union for women increased one year from 20 to 21 (Solís and Puga 2008). Such trends have also remained remarkably stable over time, even as Mexican women have achieved higher

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levels of education than at any other point in the history of the country, including educational parity or near parity with men at all levels of education (Behrman, Gaviria and Szekely 2001; Giorguli Saucedo 2010; Terrazas, Papademetriou and Rosenblum 2011; Parrado and Zenteno 2005; Creighton and Park 2010). Stability in first union timing in spite of educational gains presents an anomaly in the context of other countries experiences where higher educational levels are typically associated with marital delays (Mensch, Singh, and Casterline 2005). Frequently, Mexico's exceptionalism is interpreted as cultural, i.e. powerful familistic forces structure young women's early life course transitions even in spite of higher educational attainment and lower expectations for childbearing that have led to changes in women's status (Fussell and Palloni 2004). In this paper, we challenge the one-size-fits-all explanation for the anomalous association between educational attainment and timing to first union in Mexico. In particular, we use data from Wave 3 of the Mexican Family Life Survey (MxFLS), collected between 2009 and 2012, to consider whether the associations between educational attainment and timing to first union have changed over time. We are especially interested in the experience of women coming of age in the late 1980s and 1990s, a time of remarkable educational expansion as well as increasing levels of social inequality.

Competing Theories of Educational Attainment and Timing of Union Formation in Mexico

Educational Expansion in Mexico

Over the last few decades, Mexico, along with much of Latin America, experienced intense educational expansion (Esteve, López-Ruiz, and Spijker 2013). In 1993, the number of years of compulsory education in Mexico increased from 6 to 9 years (up to age 15) (Mier, Rocha, and Romero 2003). During the 1970s, 91 percent of the adult population had not completed nine years of basic education. By 2000, this number had been cut nearly in half, with gender differences completely eliminated (Creighton and Park, 2010; Santibanez, Vernez, and Razquin 2005). Yet while the expansion of basic education has been large, in 2010, the average educational attainment of the population aged 15 and older was still only 8.6 years, with 41 percent of the population completing less than secondary education and less than one-fifth of the population 25 years and older holding a college degree (INEGI 2010). It also remains the case that the expansion has been uneven across different sectors of society, e.g. lower in rural areas versus urban, and that the system is still characterized by unequal access and differential quality across social groups (Mier, Rocha, and Romero 2003).

Notwithstanding gaps in the expansion of education, the puzzle remains that increasing levels of education, no matter how uneven, have not affected the relatively constant early age at first marriage in Mexico[1]. Typically, educational gains are associated with marital delays (Mensch et al. 2005). Although the causal ordering is not completely clear, three perspectives dominate the literature on educational attainment and timing of marriage (Lindstrom and Paz 2001).

[1]In our review of the literature, we refer broadly to the institution of marriage. In practice, we operationalize marriage as including both legal marriages and consensual unions. We use the term "first union" to describe our dependent variable.

Role Incompatibility

Perhaps the most straightforward explanation for the inverse relationship between women's educational attainment and timing of marriage is the role incompatibility hypothesis, which posits that education delays marriage primarily through school enrollment (Lindstrom and Paz 2001). Because the role of a student is hypothesized to be incompatible with marriage and childbearing, more education is expected to translate into delays in marriage formation. In general, one of the major challenges to understanding the inverse relationship between women's educational attainment and the timing of marriage has been the failure of past analyses to separate out the effects of school enrollment from educational attainment (Raymo 2003). In the studies that have done so successfully, evidence in support of role incompatibility has been strong across a range of societies, including Japan (Raymo 2003, Shirahase 2000), China (Tian 2013) and Mexico (Lindstrom and Paz 1992, Parrado and Zenteno 2002), among others. In the case of Mexico, Lindstrom and Paz (2001) use data from the 1992 Women's Status and Fertility Survey and find that role incompatibility explains a significant portion of education's negative effect on first union formation. Once school enrollment is accounted for, the only increase in schooling that remains associated with a lower risk of marriage occurs at the lower end of the educational distribution (from 0–2 years to 3–6 years of education). At the aggregate level, the authors conclude that growth in female education has not substantially altered the mean age at marriage in Mexico because it has involved ages at which schooling does not run into conflict with union formation.

Female Independence Model

First formulated by Becker (1981) as the economic independence hypothesis (also referred to as the female independence model), the second perspective hypothesizes that as women's educational opportunities increase, marriage will become deferred and, to some extent, forgone, as women become less dependent on men and their economic gains with marriage decrease (Becker 1981; Dominguez-Folgueras and Castro-Martin 2008; Raymo 2003). In this way, investment in human capital decreases women's expected dependence on their husband's earnings. A related hypothesis is that higher educated women marry later and less because of the transformative power of schools where women are exposed to non-traditional ideas and alternative role models to those of wife and mother (Lindstrom and Paz 2001; Gyimah 2009).

The strongest empirical support of the female independence theory has come from countries with lower gender equality. According to Blossfeld (1995), in societies in which gender roles make it difficult to combine work and family, women with higher education are more likely to experience less and later marriage. Indeed, a negative relationship between education and marriage has been found in the cases of Japan (Raymo 2003), Italy (Billari et al. 2002), and Spain and Portugal (Dominguez-Folgueras and Castro-Martin 2008). In the case of Latin America, and Mexico specifically, there has been little evidence in support of the female independence model. In an analysis of the 1998 Mexican Retrospective Demographic Survey, Parrado and Zenteno (2002) find that, once the delaying effect of enrollment is controlled for, highly educated women are actually *more* likely to marry.

Specialization and Trading Model

In addition to Becker's economic independence model, Oppenheimer's (1988) specialization and trading model (also referred to as the marital search model) is the other dominant perspective on educational attainment and the timing of marriage (Gyimah 2009). In Oppenheimer's formulation, timing to marriage is related to the uncertainties surrounding the transition to adult economic roles, with the symmetrical quality of economic dependence more important than each individual's contribution. In this way, marriage serves both as a buffer against economic risk for those who are disadvantaged and as a method of preserving wealth for those with more resources (Fussell and Palloni 2004). In the case of education, the model would expect that women with low levels of education are more likely to enter into marriage than women with intermediate levels of education who face more uncertainty with respect to their future earnings potential (Parrado and Zenteno 2002). Similarly, highly educated women would also display an increased likelihood for earlier marriage because the educational elite (higher than 13 years of education) make attractive partners with secure economic futures and therefore, in contrast to the patterns observed in other countries (e.g. Japan), do not delay marriage (net of school enrollment) (Raymo 2003).

Applying Oppenheimer's model to the case of Mexico, Parrado and Zenteno (2002) use data from the 1998 Mexican Retrospective Demographic Survey (Encuesta Demográfica Retrospectiva, EDER) and find a curvilinear pattern between education and the transition to marriage, with an increased likelihood of earlier marriage for women at the lowest (0–6 years) and highest (13+ years) levels of education in Mexico. They argue that women with both low and high levels of education are more likely to enter into marriage than women with intermediate levels of education who face more uncertainty with respect to their future earnings potential. At the high end of the educational distribution they find that improvements in women's education did not "reduce the desirability of marriage in Mexico, as female independence theories predict" (2002: 769).

Yet whether this pattern continues to hold among more recent cohorts of women is not yet known. The "young" cohort analyzed by Parrado and Zenteno was born between 1965–69 and was only observed until 1998. The possibility that the relationship between higher education and timing to marriage may have changed in more recent times is particularly salient given recent trends in the transition to adulthood described below.

Birth Cohort Perspective in the Mexican Context

A tacit assumption of much of the prior research on educational attainment and marriage has been that the effect of educational attainment on timing to first marriage has remained constant across successive generations of women. More recently, several studies have empirically evaluated this claim and found violations to this assumption in the case of Japan (Raymo 2003), Ghana (Gyimah 2009) and Latin America as a whole (Esteve et al. 2013). The possibility that the relationship between education and timing to marriage may have changed over time is particularly relevant in the Mexican context. During the 1990s, substantial expansions in education and labor force participation among women were off-set by prevailing high poverty rates, economic inequality and unequal access to the emerging educational and labor opportunities for the young. The result in Mexico, and in Latin

America as a whole, has been an increasing polarization in the transition to adulthood (Esteve et al. 2013; Giorguli Saucedo 2010). Increasing social differentiation points to the possibility of a newly altered relationship between education and first union timing in Mexico for a younger cohort of women coming of age in the 1990s. Our expectation is that it is precisely among this newer cohort that the education-first union timing relationship will have changed. In particular, we test whether the curvilinear relationship observed in previous studies holds for the most recent cohort of women. We hypothesize that among young women a bifurcated pattern has emerged. That is, we expect that first unions will continue to operate as a buffer against the uncertainty of the economy for the majority of women who still enter first unions early, net of school enrollment. But for a small elite group of women, we hypothesize that education will be increasingly associated with a delaying effect as specified by the female independence model.

Data

Our data come from the Mexican Family Life Survey (MxFLS), a nationally- representative longitudinal dataset of 8,440 households in 150 communities. The first wave was collected in 2002 and Waves 2 and 3 were collected in 2005 and 2009–2012, respectively. In this paper we use data from Wave 3, during which investigators followed the baseline households with a 90 percent retention rate (Rubalcava and Teruel 2013). These data are uniquely suited to our study in that they contain detailed retrospective co-residential union and educational attainment histories from adult women (ages 15 and older) interviewed. The data also allow us to include controls for socioeconomic background during childhood, an important feature given the unequal expansion of education across sociodemographic strata.

In order to facilitate comparisons between women across cohorts, it is important to include identical age categories per cohort (Parrado and Zenteno 2002). Because national statistics indicate that the majority of women in Mexico enter their first union by their late 20s (Solís and Puga 2008), following women in each cohort through age 30 allows us to observe the majority of first union transitions within each cohort. Accordingly, our analytic sample includes women who are at least 30 years old at the time of interview. Of the 12,794 women interviewed in the MxFLS-3, 7,824 are at least 30 years of age and therefore form the basis of our analytic sample. From our analytic sample individuals were excluded via listwise deletion if they were missing on year of first marriage/consensual union (n=168), education (n=3), rural household at age 12 (n=90), or poor household infrastructure at age 12 (n=3). An additional 202 women were excluded because they were born before 1930 and were scattered thinly across various cohorts, comprising too select a group for analysis. Finally, 82 women were excluded because they reported their first marriage/consensual union as occurring prior to age 12. In total, 548 women in our analytic sample were excluded from our final sample, with 291 because of missing data. Our final sample consists of 7,276 women.

Key Variables

Year of First Union Formation

Our outcome of interest is the timing to a first marriage or consensual union (i.e. cohabitation). Respondents are not asked to differentiate whether their first union began as a marriage versus cohabitation, and therefore we are not able to account for the difference in our analysis. Traditionally, consensual unions have been a stable form of partnership common in Mexico, particularly in poorer rural areas, and have generally been viewed similarly to marriage (Martin 2002). Other work estimating the relationship between education and union formation in Mexico found no evidence that the effects of education on entry into a union varied between consensual and formal unions (Lindstrom and Paz 2001). For simplicity of language, from this point on we will refer to the dependent variable as first union, though we note that the first union could either refer to a legal marriage or an informal consensual union.

Age Cohorts

We categorize our sample into three age cohorts representative of roughly three successive generations: a mature cohort born between 1930 and 1949, a middle cohort born between 1950 and 1969, and a young cohort born between 1970 to 1979[2]. Women in the mature cohort were approximately 68 years of age at the time of interview, women in the middle cohort were 49 years of age, and women in the young cohort were 35 years of age on average. While we acknowledge the point made by Raymo (2003) that, "any categorization of birth cohort is ultimately arbitrary" (90) and while our pattern of results does not change under a 10-year cohort specification, our reason for splitting the cohorts using these cut points was two-fold. First, we wanted to make comparisons across women from multiple generations and the roughly 20 year time-span in the first two cohorts (and 10 years for the most recent) allows us to do this. Second, the cut points chosen were also methodologically motivated. That is, after first partitioning our analytic sample into 10 year cohorts, we grouped together neighboring cohorts of women who shared proportional baseline hazards concerning age at first union formation (explained in greater detail in the Methods section that follows).

Education

We measure education with two time-varying indicators, one for highest level of education completed and the other for school enrollment. The MxFLS is unique in that it asks the age at which women began and finished school, in addition to schooling interruption questions. We utilized this information in conjunction with the average age of completion for each educational level in Mexico to create our time-varying education and enrollment variables. For instance, for a woman who reports completing high school, she would have a 1 on the school enrollment variable from ages 12 (when she enters the study) to 18. Completed education level includes categories for less than primary education (0 to 5 years), completed primary education (completed 6 to 8 years), completed secondary education (completed 9 to

[2]125 women were interviewed after 2009 and were born 1980 or later. We placed these women in with the 1970–79 cohort because although they were born after 1979, there were too few to label the cohort as representative of women through 1982.

11 years), completed high school (12 years), and attended at least some college (13+ college). School enrollment is a dichotomous time-varying indicator that equals 1 when the respondent is enrolled in school and equals 0 otherwise.

Controls for Sociodemographic background

We measure sociodemographic background with three indicators. These include two indicators of social disadvantage during childhood, i.e. whether the respondent lived in a household with poor infrastructure at age 12 (operationalized as having no running water for drinking and/or no plumbing) and whether the respondent lived in a rural area at age 12 (with rural including rancherias, pueblos, ejidos, haciendas, villas, and ‘other’ rural areas). Further, we control for whether the respondent identifies as indigenous. We recognize that material indicators of poverty have likely changed across cohorts as Mexico becomes more developed and urbanized, e.g. having been raised in a house with poor household infrastructure is less common among more recent cohorts and as a result may now be more indicative of deep poverty than was the case for earlier cohorts. Notwithstanding these potential changes, we reiterate our purpose for including the sociodemographic controls which is to increase comparability in the results across cohorts, leaving open the possibility that their substantive meaning may have changed over time.

Methods

We utilize event history modeling, i.e. discrete-time regression analysis, to examine whether the relationship between educational attainment and timing to first union has changed across recent generations. In these analyses, the unit of analysis is person years and the dependent variable is whether a woman has entered a first union. Because we estimate our models with logistic regression, the outcome variable is interpreted as the log odds of entering into a first union for woman *i* at year *t*, conditional on remaining single through year *t*-1, where p_{it} is the probability of having married for woman *i* at year *t*:

$$\ln[p_{it}/(1- p_{it})]$$

In order to facilitate comparisons between women across cohorts, we follow women from age 11 (when no women are in a union) through age 30. If a woman has entered a union by age 30, in each year prior to the recorded age at first union, she is given a value of 0 on the ever-partnered variable. Once a woman enters a first union she receives a 1 on the ever-partnered variable and is censored from the analysis. Women who are never-partnered through age 30 are censored from the analysis at age 31. Another important feature of discrete-time regression analysis is the need to control for the duration dependence; otherwise the model treats the duration dependence as constant (i.e. it does not account for different women having different exposure times regarding the risk of a first union) (Box-Steffensmeir and Jones 2007). We operationalize the duration dependence with the inclusion of age dummies in each model. The inclusion of only the age dummies in a logistic regression model enables estimation of the baseline hazard. A key assumption of discrete-time logistic regression models is that the categories of each predictor entered into a given model will share proportional baseline hazards. If this assumption is violated, for example across age cohorts, than it risks biased estimates, not only for each age cohort, but also for any indicators with which they may be interacted (e.g. educational attainment). Because of our interest in estimating the effects of educational attainment on timing to first union across age cohorts, we first test the assumption of proportional hazards across cohorts. To accomplish this, we estimate Models A and B,

MODEL A: $\ln[p_{it}/(1- p_{it})] = \beta_1 DUR_{it} + \beta_2 COHORT_i$

MODEL B: $\ln[p_{it}/(1- p_{it})] = \text{MODEL A} + \beta_3(DUR_{it} \text{ X } COHORT_i)$

Where β_1 represents the duration dependence, β_2 represents categories for age cohort, and β_3 represents interactions between duration with cohort. Because Model A is nested in Model B, we are able to test the assumption of proportional hazards with a Likelihood Ratio Test. Results from this test were significant (*X2*=103.21; p<.001), indicating that the proportional hazards assumption was violated across cohorts[3]. Figure 1 illustrates the predicted baseline hazard of first union entry from ages 11 to 30 separately by age cohort (estimated from Model B). If the proportional hazards assumption was not violated, we would see parallel lines for each cohort (i.e. the individual cohort paths would not cross or diverge from one another in shape). Because the proportional hazards assumption is violated, we run all subsequent models separately for the mature, middle, and young age cohorts.

To examine the influence of educational attainment on the odds of first union entry, we estimate the following models for each cohort:

MODEL 1: $\ln[p_{it}/(1- p_{it})] = \beta_1 DUR_{it} + \beta_2 EDUCATION_{it} + \beta_3 CONTROLS_i$

MODEL 2: $\ln[p_{it}/(1- p_{it})] = \text{MODEL 1} + \beta_4 ENROLLED$

Where in Model 1, in addition to the duration dependence, estimates for time-varying educational attainment (β_2) and the time-invariant controls (β_3) for poor household infrastructure at age 12, rural household at age 12, and indigenous identity are also included. This model will allow us to estimate the effect of educational attainment on timing to first union while adjusting for social background factors that bias access to education. In Model 2, school enrollment (β_4) is added to equation, allowing estimation of educational attainment net of the delaying effects of school enrollment.

Results

Table 1 presents means for key variables in the analysis, both for the total sample (N=7,276), and by age cohort (Mature = 1,478; Middle = 3,471; Young = 2,327). The majority of women in each cohort have entered their first union- even in the youngest cohort 86% of women have done so. Every variable has significantly different values across cohorts. In the case of education, over 75% of women in the mature cohort are in the lowest schooling category (0–5 years) and only 3% have 13+ years of education. For the youngest cohort (1970–79), only 16% have 0–5 years education and 10% have 13+ years. In addition, fewer women report living in rural households or in households with poor infrastructure at age 12 in the young cohort than in the middle and mature cohorts.

[3]We also calculated the AIC (Akaike's information criterion) for both models and Model B had a smaller value than Model A (40700 vs 40727), indicating better model fit.

Table 2 illustrates survival estimates for the ages at which 25, 50, and 75 percent of women within each cohort enter their first union. Comparing across cohorts, the time span over which specific proportions of women marry has increased significantly, despite early union formation remaining a dominant pattern. For example, 25% of women in the mature cohort entered a first union by age 17, while 75% of them did so by age 24, resulting in a 7-year range between the first and third quartiles. This inter-quartile range increases steadily across cohorts. For instance, among women in the young cohort, 25% formed their first union by age 18, while 75% of women in this most recent cohort married by age 27, increasing the inter-quartile range to 9 years. The largest cohort differences are observed in the third quartile, indicating increasing variation in age at first union entry over multiple generations. These patterns suggest that in the younger generations, different pathways to union formation are coexisting alongside more traditional models.

Table 3 presents discrete-time logistic regression of first union formation on duration dependence, educational attainment, social background factors, and school enrollment (introduced in Model 2), separately by cohort.

Mature Cohort

For women in the mature cohort (1930–49), Model 1a demonstrates a curvilinear relationship between educational attainment and timing to first union. Before accounting for the delaying effects of school enrollment, having 6 to 8 years of education decreases the odds of union entry by 25% (1-exp(−.29)=.25) and for women with 9 to 11 years of education by 37% (1-exp(−.47) =.37). Women in the two highest education categories show no significant differences in their odds of first union formation as compared to women with less than primary school. Only women with intermediate levels of education demonstrate lower odds of union formation. Enrollment explains some of this pattern. Among women with 9 to 11 years of education, their lower odds become non-significant when we account for school enrollment (Model 2a), indicating that for these women the delaying effect of education on timing to first union operates solely through enrollment in school. The same is not true for women with between 6–8 years of education. Even after controlling for school enrollment, having 6 to 8 years of education is associated with 19% (1- exp(−.22)=.19) lower odds of entering a first union compared to having 0 to 5 years of education. School enrollment itself is highly significant; during years when women are enrolled in school, their odds of entering a first union decrease by 85% (1-exp(−1.91)=.85). It is notable that even after accounting for educational attainment and school enrollment, women who identify as indigenous have 26% (exp(.23)=1.26) greater odds of forming a first union than their nonindigenous counterparts. Further, having lived in a rural household at age 12 increases the odds of forming a first union by 17% (exp(.16)=1.17).

Middle Cohort

For women in the middle cohort (1950–1969), all educational attainment categories beyond 0 to 5 years are associated with lower odds of union formation before controlling for school enrollment (Model 1b). For 12 or 13+ years of education, however, the delaying effect of education operates solely through enrollment. Once enrollment is accounted for in Model 2b, women who have completed upper secondary school or who have some college education are not significantly different from women with the lowest levels of education in their odds of union formation. The same is not true for women with intermediate levels of education. Having 6 to 8 years or 9 to 11 years of education still lowers the odds of first union formation net of school enrollment by 17% (1-exp(−.18)= .17) and 23% (1- exp(−.26)=.23), respectively. Similar to the pattern in the mature cohort, we observe a curvilinear relationship, with women in the lowest and highest educational levels displaying higher odds of union formation, net of school enrollment. School enrollment itself lowers the odds of union formation by 66% (1-exp(−1.19)=.66). In terms of social background characteristics, unlike women in the mature cohort, there are no significant effects after accounting for education and school enrollment.

Young Cohort

For women in the young cohort (1970–79), a different pattern emerges with respect to education and first union timing. For the first time, women with 6 to 8 years of education are the only education category not significantly different from women with less than a primary education in their odds of forming a first union (Model 1c). Women in all the other education categories are significantly less likely to form a union compared to the least educated. Model 2c demonstrates, however, that except for those with at least some college, the delaying effect of education operates entirely through enrollment. School enrollment reduces the odds of entering a first union by 69% (1-exp(−1.17)=.69). After controlling for it, women with education levels including 6 to 8, 9 to 11 and 12 years, are not significantly different from those with 0–5 years of education in their odds of union formation. The only education category that remains significant is 13+ years of education which decreases the odds of union formation by 25%. As was the case with women in the middle cohort, social background characteristics do not significantly impact the odds of union entry for women in the young cohort net of controls for education and enrollment.

Overall, the results suggest that the curvilinear pattern noted by Parrado and Zenteno (2002) is present for the mature and middle cohorts, but not the young cohort. For the youngest, instead of the least and most educated women having the highest odds of marriage, women with 13+ years of education have lower odds of marriage compared to their counterparts with 0 to 5 years of education.

Discussion

This paper presented an updated analysis of the relationship between educational attainment and the timing to first union in the context of Mexico, a country whose nuptial regime has been characterized by early and nearly universal marriage for women. In contrast to this picture of relative stability, our findings suggest a nuptial system characterized by a mixture of stability and change, with increasing social differentiation in first union formation between the least educated and the educational elite.

Moving beyond measures of central tendency, the survival estimates of the ages at which different proportions of women marry across cohorts suggests increasing variation over time. While early age at marriage remains a dominant pattern, we also find that the time span over which specific proportions of women marry has increased. The largest differences across cohorts are observed in the third quartile, indicating an increasing age at first union for certain groups of women in Mexico. These results suggest the emergence of increasingly heterogeneous pathways with respect to timing to first union among women in Mexico (Solís and Puga 2008; Quilodran 2008).

To better understand the increased social differentiation in timing to first union we focused on the relationship between educational attainment and timing to first union across cohorts. Our analyses find strong support for the role incompatibility hypothesis, i.e. that for young women enrollment in school delays entry into the marriage market largely because of the incompatibility between student and spousal roles (Lindstrom and Paz 2001). The delay appears slightly smaller in the two more recent cohorts compared to the first, but it is positive and highly significant in all three.

In contrast to the role incompatibility hypothesis which displays fairly uniform effects over time, the other two theories appear to vary in their applicability across cohorts. Past analyses have found support for the specialization and trading model, with highly educated women in Mexico being similar to the least educated in their timing to marriage (Parrado and Zenteno 2002). According to the logic of this model, highly educated women represent "a selected elite in Mexico and their potential contribution to the household economy makes them attractive partners" (Parrado and Zenteno 2002: 768). Our analysis supports this finding for women born prior to 1970. We find a curvilinear pattern similar to the one documented by Parrado and Zenteno (2002) for women born between 1930 and 1969, precisely the same cohorts of women included in their analysis. For women who were born between 1930–1969, once we account for the delaying effect of enrollment, those with the lowest and highest levels of education are characterized by the earliest transitions to a first union relative to the other education categories. In contrast, women with either a high school education or some college show no significant delay in these cohorts. We interpret these patterns in the same way as past researchers (e.g. Parrado and Zenteno 2002; Fussell and Palloni 2004), that in face of uncertainty, the lowest and highest educational groups transition to a first union early.

There is one noteworthy difference in the shape of the curvilinear pattern across the mature and middle cohorts. For women born between 1930–49, only those women with 6–8 years of completed education demonstrate a lower risk of entering a first union relative to the least educated. In the middle cohort (1950–69), the lower risk of first union formation extends to women with between 6–11 years of education, with the largest reduction in risk for women with 9–11 years of schooling relative to those with 0–5 years. We interpret this shift in light of the expansion of education in Mexico over the later decades of the 20th century. In the 1930–49 birth cohort, completion of secondary school (9 years) was relatively rare for women and represented a high level of schooling (Torche 2010). Following the logic of the specialization and trading model, these women would have represented attractive partners, increasing their likelihood of forming a union relative to the least educated. But by the middle cohort (1950–69), completion of secondary school was more normative and thus women with between 9 to 11 years of schooling no longer represented an educational elite with early entry into a first union (Torche 2010), a pattern reserved for those with 12 or more years of education. Instead, women with 9 to 11 years of education in the middle cohort now held an intermediate level of educational attainment and, according to the specialization and trading model, faced the most uncertainty in their economic prospects. As a result, they demonstrate lower odds of union formation compared to the least educated, just as their counterparts with 6 to 8 years of education in the mature cohort.

For women born between 1970 and 1979, we find that the pattern between education and first union formation has changed. Once we control for enrollment, none of the educational groups is significantly different from the least educated in terms of timing to first union except for the most highly educated. In contrast to their peers born in earlier cohorts, highly educated women in Mexico are now significantly different from the least educated in the extent to which they delay entry into a first union. Essentially, having a college education now decreases the odds of union formation relative to having less than primary schooling whereas in the past the highly educated were similar to the least educated in their timing to first union.

These results suggest that after accounting for school enrollment, only the most highly educated women are actually postponing first union formation relative to least educated. Our findings for the most recent cohort differ from the curvilinear pattern evident in our sample for women born between 1930 and 1969 and also documented in other work (Parrado and Zenteno 2002). We offer several possible non-competing explanations for this change. First, these results could support the female independence model, which suggests that at the upper ends of the educational distribution, as women's educational opportunities increase, marriage will become deferred and, to some extent, forgone, as women become less dependent on men and their economic gains with marriage decrease. Another possibility is that the postponement of marriage until older ages among the college educated may reflect changes in the meaning of marriage, such that for women who can afford to support themselves, marriage has become more about individual fulfillment and less about economic necessity (Cherlin 2010). Relatedly, assortative mating on the basis of education, i.e. that college educated women tend to marry college educated men, may limit the availability of suitable partners until later in adulthood (Cherlin 2010). Finally, it is possible that economic uncertainty within Mexico in recent years may have increased the need for gender specialization within households (Freije, López-Acevedo, and Rodríguez-Oreggia 2011 ref), making highly educated women less desirable as partners (Torr 2011).

With respect to the absence of significant differences in the risk of first union formation among women with anything less than some college education, the findings support other research that argues that additional years of schooling no longer translate into significant gains in income until after one moves beyond upper secondary school (13+ years). According to Giorguli (2010), growth during the 1980s and 1990s of Mexico's large informal sector has distorted the traditional link between education and income, so young adults are now shut out of former sector jobs until finishing upper secondary and beyond. As a result, the incentive to remain in school is potentially low among much of the population. In this context, economic uncertainty likely results in adolescents, and girls in particular, accelerating their transitions to adulthood, including exits from school and early entry into marriage at all levels of education prior to college (Echarri Cánovas and Pérez Amador, 2007; Giorguli Saucedo, 2010).

Taken together, the pattern that in the youngest cohort only college-educated women delay first union formation relative to the least educated net of enrollment lends some support to Giorguli's (2010) argument that persistent inequality in Mexico has resulted in "two distinct transition patterns to adulthood, one applying to the great majority and the other to a small elite group" (24). In the case of entry into marriage, we find evidence for the increasing social differentiation of the educational elite whose pathways to adulthood vary markedly from the rest of the population.

There are several limitations to our analysis that deserve mention. First, the primary interest of this paper is female educational attainment and how it relates to timing to first union formation. We believe that this focus is warranted given the dramatic changes in education for women in Mexico over the last several decades and the anomalous curvilinear relationship documented between education and union formation in Mexico in the existing literature. That said, individual-level labor force participation as well as macro-level economic conditions are clearly a key part of any hypotheses about the changing prospects of marriage for limiting economic uncertainty. Unfortunately, we were unable to construct a suitable measure of labor-force participation due to issues with missing data and non-optimal skip patterns. Another piece of the puzzle missing from our analysis is men. We defend our exclusive focus on women because of our interest in evaluating the changing relevance of the female independence model in Mexico. We acknowledge that men are an important part of the story and plan to explore the possibility of changes in their patterns of first union formation in future analyses (Parrado and Zenteno 2002). Another limitation is our inability to distinguish between whether a union began as a marriage or cohabitation. While some research suggests that in much of Latin America the two are indistinguishable (Martin 2002), more recent research has suggested that the increasing number of cohabitations are one of the reasons that the average age at marriage in Latin America has not increased in tandem with higher education levels (Esteve et al. 2013). Cohabiting unions are clearly a distinct piece of this story and distinguishing how they may or may not differ from marriage remains an important future task.

Conclusion

Our evidence suggests that as women's opportunities for college education increase in Mexico, marriage will be increasingly deferred for an elite group of young women. We have interpreted this trend as being indicative of women becoming less dependent on men and their economic gains with marriage decrease although alternative possibilities exist. Regardless of the source, whether this trend continues to exacerbate levels of inequality in Mexico is not clear. One possibility is that as opportunities for higher education continue to expand, cultural pressures on women in Mexico to marry and have children early will become less salient for all young women, regardless of education level, instead giving way to the ideals associated with the female independence model. Currently, however, insufficiencies in the expansion of education and limited opportunities in the labor market will likely defer this process and exacerbate social differentiation between groups. Which of these scenarios will eventually prevail is left to be determined by future cohorts of young women as they transition into adulthood in Mexico.

Acknowledgments

The first author acknowledges support by an F32 NICHD Ruth L. Kirschstein National Research Service Award (F32 HD072616), as well as an NICHD center grant to the Population Research Center at The University of Texas at Austin (R24 HD042849). The first, second, and third authors also acknowledge the support of the Institute for Population Research at the Ohio State University, supported by a grant from the Eunice Kennedy Shriver National Institute for Child Health and Development (R24 HD058484).

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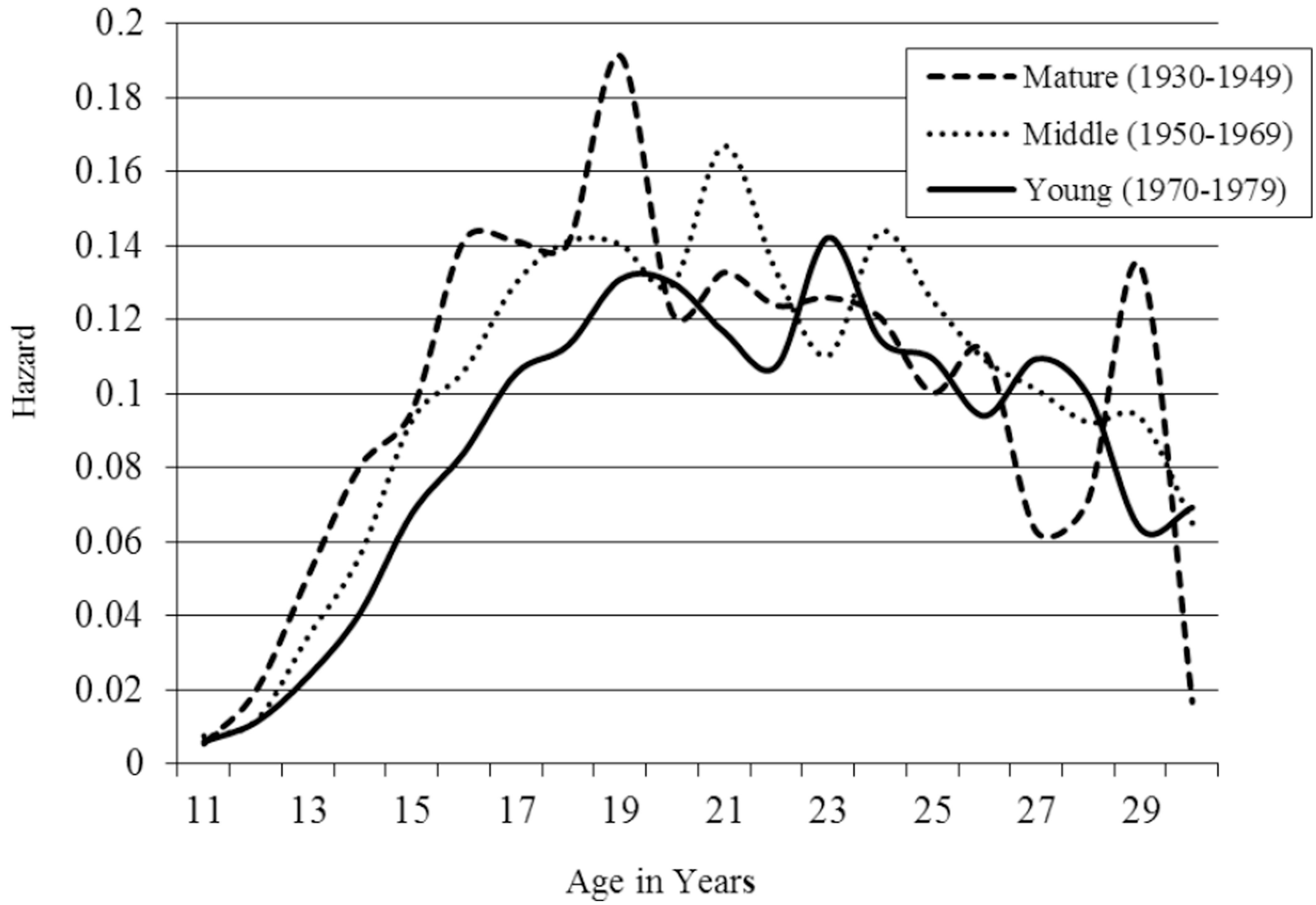

Figure 1.
Predicted Baseline Hazards for First Union Formation, by Cohort

Table 1

Means of Key Variables, by Cohort

	Total Sample	Mature (1930–1949)	Middle (1950–1969)	Young (1970–1979)	Significant Cohort Differences*
Age	48.37	68.35	48.96	34.79	*a*, *b*, *c*
Ever Married	0.92	0.95	0.93	0.86	*a*, *b*, *c*
Years of Education					
0 to 5 years	0.39	0.76	0.40	0.16	*a*, *b*, *c*
6 to 8 years	0.23	0.16	0.26	0.24	*a*, *b*
9 to 11 years	0.22	0.05	0.20	0.38	*a*, *b*, *c*
12 years	0.07	0.02	0.07	0.12	*a*, *b*, *c*
13+ years	0.08	0.03	0.08	0.10	*a*, *b*, *c*
Poor HH Infrastructure Age 12	0.49	0.71	0.51	0.32	*a*, *b*, *c*
Rural HH Age 12	0.33	0.44	0.33	0.26	*a*, *b*, *c*
Identifies as Indigenous	0.13	0.16	0.13	0.11	*a*, *b*, *c*
N	7276	1478	3471	2327	

Note:

[a] = significant difference at $p<.05$ or less between mature and middle cohorts;

[b] = significant difference at $p<.05$ or less between mature and young cohorts;

[c] = significant difference at $p<.05$ or less between middle and young cohorts

Table 2

Percentiles and Interquartile Range of Age at First Union Formation, by Cohort

	25th Percentile	50th Percentile	75th Percentile	Years from 25th–75th Percentile
Mature (1930–1949)	17	19	24	7
Middle (1950–1969)	17	20	25	8
Young (1970–1979)	18	21	27	9

Note: Log-rank test for trend of survivor functions indicates that cohort differences in survival time are statistically significant ($X2$=63.53; p <.001)

Table 3

Discrete-Time Logistic Regression of First Union Formation, by Cohort

	Mature (1930–1949)		Middle (1950–1969)		Young (1970–1979)	
	Model 1a	**Model 2a**	**Model 1b**	**Model 2b**	**Model 1c**	**Model 2c**
	b (s.e.)	**b (s.e.)**	**b (s.e.)**	**b (s.e.)**	**b (s.e.)**	**b (s.e.)**
Years of Education (Time-Varying)						
0 to 5 years	ref	ref	ref	ref	ref	ref
6 to 8 years	−0.29*** (0.09)	−0.22* (0.09)	−0.27*** (0.05)	−0.18*** (0.05)	−0.07 (0.08)	0.03 (0.08)
9 to 11 years	−0.47** (0.14)	−0.21 (0.15)	−0.44*** (0.06)	−0.26*** (0.06)	−0.28*** (0.08)	−0.12 (0.08)
12 years	−0.32 (0.25)	−0.10 (0.25)	−0.36*** (0.09)	−0.17 (0.09)	−0.26* (0.10)	−0.08 (0.10)
13+ years	−0.18 (0.22)	0.08 (0.22)	−0.27** (0.09)	0.04 (0.09)	−0.60*** (0.11)	−0.29* (0.12)
Enrolled in School (Time-Varying)	---	−1.91*** (0.37)	---	−1.09*** (0.09)	---	−1.17*** (0.10)
Poor HH Infrastructure Age 12	0.07 (0.08)	0.06 (0.08)	0.13** (0.05)	0.08 (0.05)	0.10 (0.06)	0.03 (0.06)
Rural HH Age 12	0.17** (0.07)	0.16* (0.07)	0.05 (0.05)	0.03 (0.05)	0.09 (0.06)	0.05 (0.06)
Identifies as Indigenous	0.24** (0.08)	0.23** (0.08)	−0.02 (0.06)	−0.02 (0.06)	−0.02 (0.08)	−0.02 (0.08)
Intercept	−5.35*** (0.36)	−5.15*** (0.36)	−4.91*** (0.20)	−4.41*** (0.20)	−5.15*** (0.27)	−4.28*** (0.28)
N (Person-Years)	14,377	14,377	35,761	35,761	26,179	26,179
df	26	27	26	27	26	27
Likelihood Ratio $X2$	649.75***	698.08***	1546.95***	1734.11***	970.02***	1133.32***

Note:

*** p<0.001,

** p<0.01,

* $p<0.05$;

All models include age dummy variables to control for duration dependence (ommitted from above results for brevity)