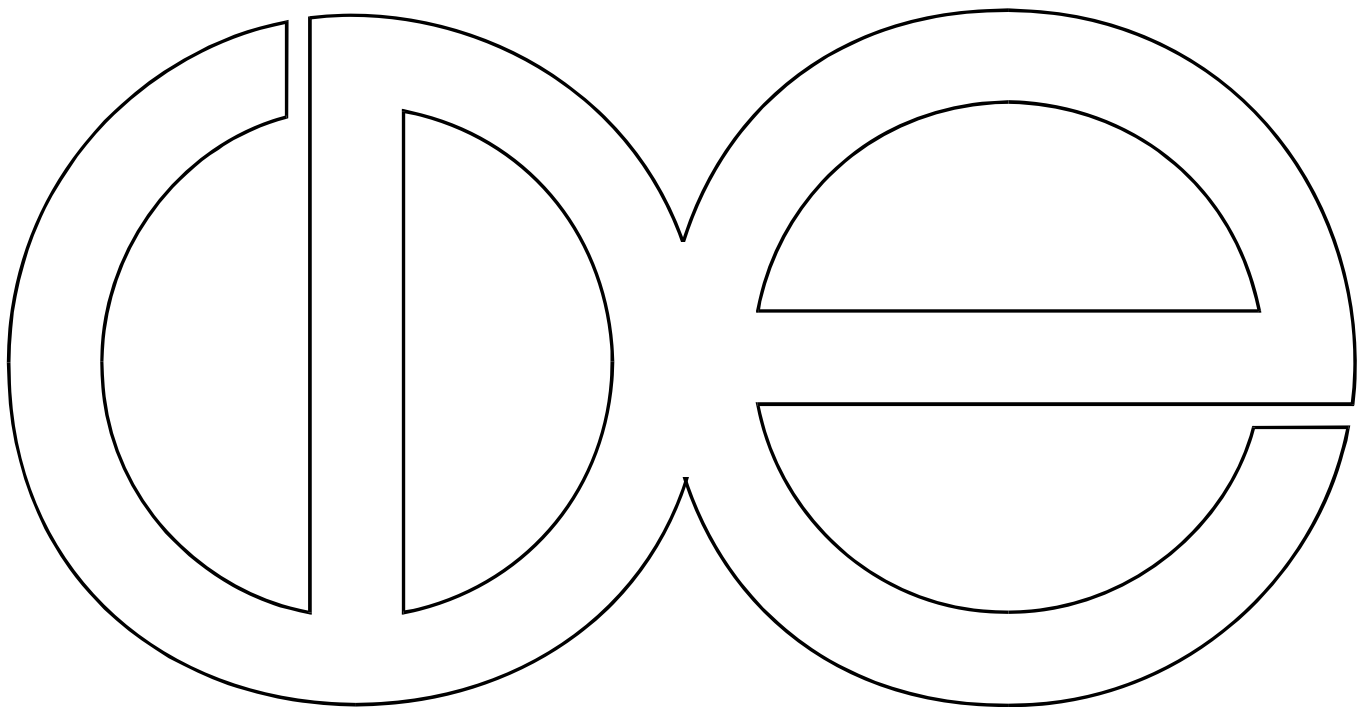


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Cohabitation and First Marriage in Japan

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Abstract

In this paper, we use nationally representative data to describe basic characteristics of cohabiting unions in Japan. We also examine the correlates of cohabitation experiences and evaluate the relationship between cohabitation and the transition to first marriage. We demonstrate that cohabitation has increased rapidly among recent cohorts of women and that cohabiting unions in Japan tend to be relatively short and almost as likely to dissolve as to result in marriage. Simple models for the timing and nature of first marriage suggest that cohabitation is only weakly associated with marriage timing. Results change dramatically when we estimate models that account for self-selection into cohabitation. Controlling for unobserved characteristics associated with both selection into cohabitation and first marriage, we find that cohabitation experience itself is strongly associated with earlier marriage. We conclude with speculation about the likelihood of further increases in cohabitation and the potential implications for marriage and fertility.

Introduction

The rapid and widespread emergence of nonmarital cohabitation is a defining feature of recent family change in industrialized countries and is closely linked with concurrent changes in marriage and childbearing. In the U.S, declining rates of marriage at young ages have been offset by increases in cohabitation (Bumpass, Sweet, and Cherlin 1991) and an increasing proportion of conceptions and childbirths occur within cohabiting unions (Kennedy and Bumpass 2007; Wu, Bumpass, and Musick 2001). Explanations for increases in the prevalence of nonmarital cohabitation have emphasized relatively universal forces of social, economic, and ideational change (Bumpass 1990; Cherlin 2004; Lesthaeghe 1995; van de Kaa 1987), but it is also clear that the prevalence and nature of cohabitation vary substantially across countries (Heuveline and Timberlake 2004; Kiernan 2001) and subgroups of the U.S. population (Kennedy and Bumpass 2007; Raley 2000).

Research on cohabitation in Japan is conspicuous for its absence. Fragmentary data suggest that the prevalence of nonmarital coresidential unions remains low (Atoh 2001) despite steady decline in marriage rates (Raymo 1998, 2003; Retherford, Ogawa, and Matsukura 2001). The trend toward later and less marriage, combined with extremely low levels of extramarital childbearing (Iwasawa 2002), is the primary reason for the decline in Japan's period fertility from 2.14 in 1973 to 1.26 in 2005 (Tsuya and Mason 1995). The implications of delayed marriage and very low fertility for population aging and population decline are the subject of great social and political concern in Japan and several high-profile policy efforts (collectively referred to as *shōshika taisaku*) have been implemented to promote family formation and increase fertility rates. In this context, it is surprising that little scholarly or political attention has been devoted to the emergence of premarital cohabitation and its potential impact on marriage

and fertility.

A better understanding of cohabitation in Japan is also a potentially rich source of theoretical insight regarding family change. On one hand, Japan is one of the few non-western countries characterized by very late marriage, lowest-low fertility, and relatively high divorce rates and may thus shed light on the institutional and normative contexts that limit the emergence of cohabiting unions. On the other hand, recently collected survey data indicate that the prevalence of cohabitation in Japan is substantially higher than previously believed. Prospective documentation of the emergence of cohabitation in Japan and its role in the family formation process may thus provide valuable insights into the generality of patterns observed in the U.S. and elsewhere.

Our primary goal in this paper is to advance our understanding of premarital cohabitation in Japan within the limitations of newly available, nationally-representative data. We do this by (a) presenting basic descriptive information about cohabitation unions, (b) describing the socioeconomic and attitudinal correlates of cohabitation experiences, and (c) evaluating relationships between cohabitation and the transition to first marriage. We begin by describing the prevalence of cohabitation experience, the mean duration of cohabiting unions, and the proportion of cohabiting unions that end in marriage. We then estimate multivariate models for these three outcomes as a function of established covariates of cohabitation experience in the U.S. and elsewhere. Finally, we evaluate the relationship between cohabitation experience and the timing of first marriage.

Background

Cohabitation prevalence, duration, and outcomes

In the U.S., premarital cohabitation rapidly evolved from a relatively rare, “deviant” arrangement

(Rindfuss, Choe, Bumpass, and Tsuya 2004) to the modal type of first union (Bumpass and Lu 2000). Data from the most recent (2002) round of the National Survey of Family Growth indicate that over 60% of women age 25-39 had ever cohabited and that a similar proportion of recent marriages were preceded by cohabitation (Kennedy and Bumpass 2007). Cohabiting unions in the U.S. are fragile and typically short in duration. Although most cohabitators plan to marry their partner (Brown 2003), only half of recent cohabiting unions resulted in marriage (Bumpass and Lu 2000). The median duration of cohabiting unions is less than two years (Heuveline and Timberlake 2004), with approximately half ending in less than one year and only one in ten lasting for five years or more (Bumpass and Lu 2000; Bumpass and Sweet 1989).

These basic characteristics of cohabiting unions vary substantially across countries (Heuveline and Timberlake 2004; Kiernan 2001). The prevalence of cohabitating experience is particularly variable, ranging from over 80% in France to less than 10% in Italy.¹ Compared with the U.S., cohabiting unions in many European countries are relatively long in duration and more likely to end in marriage. For example, median union duration ranges from over four years in France to just over one year in the U.S. while the proportion of cohabiting unions resulting in marriage exceeds 75% in several countries (e.g., Austria, Belgium, Finland) but is less than half in others (e.g., France and New Zealand).²

Cross-national comparisons of cohabiting unions have not included Japan because adequate data have not been available. The primary source of information on cohabitation in

¹ These figures refer to the life table proportion of women who have ever been in a cohabiting union by age 45 (Heuveline and Timberlake 2004).

² Median duration is the life table estimate of the duration at which 50% of cohabiting unions end in either marriage or breakup (Heuveline and Timberlake 2004).

Japan – the National Fertility Surveys (NFS) – has collected information on current and past cohabiting unions since 1987, but only from unmarried respondents. From these surveys, it is clear that the cross-sectional prevalence of cohabitation is very low, ranging from less than 1% of unmarried respondents in 1987 to about 2% in 2002. The proportion of unmarried men and women who report ever cohabiting is also relatively low, ranging from 3% in 1987 to about 8% in 2002 (National Institute of Population and Social Security Research 2004). The actual proportion that has ever cohabited is unquestionably higher, but without corresponding data on married respondents, it is impossible to ascertain how much higher. Nevertheless, these incomplete data do suggest that Japan is a country where cohabitation remains “marginal,” with normative sanctions and institutional penalties combining to limit cohabitation to a small minority of couples (Heuveline and Timberlake 2004:1216).

The limited prevalence of cohabiting unions in Japan is somewhat surprising given the growing gap between earlier initiation of sexual activity (Japanese Association for Sex Education 2000) and later marriage (Raymo 1998; Retherford, Ogawa, and Matsukura 2001), increasing marital instability (Raymo, Iwasawa, and Bumpass 2004), and widespread skepticism about the benefits of marriage (Tsuya, Mason, and Bumpass 2004). In addition to lingering normative disapproval (Atoh 2001; Rindfuss et al. 2004), other disincentives to the formation of nonmarital unions in Japan presumably include the high cost of establishing a temporary residence, ready access to “love hotels” (Retherford, Ogawa, and Matsukura 2001), and limited financial and emotional incentives to leave the parental home before marriage (Atoh 2001; Miyamoto, Iwakami, and Yamada 1997; Yamada 1999, 2007).

Recent survey data indicate, however, that the prevalence of cohabitation in Japan is substantially higher than previously believed. For example, one in five 24-34 year old female

respondents to the first rounds of the Japanese Gender and Generations Survey and the Survey on Population, Family, and Generations (both conducted in 2004) reported that they had lived in a cohabiting union (Iwasawa 2005; Tsuya 2006). This figure is twice as high as the proportions reported by respondents just 10 years older. Documenting this rapid increase in cohabitation and observing subsequent changes in both the prevalence and nature of cohabiting unions in Japan represents a valuable opportunity to evaluate the generality of findings from research on other countries where cohabitation emerged much earlier. Does the spread of cohabitation occur as rapidly in Japan as in other settings? Are cohabiting unions in Japan relatively short in duration, as suggested by the very low cross-sectional prevalence? Are socioeconomic and attitudinal correlates of the prevalence and nature of cohabiting unions in Japan similar to those observed in the U.S. and elsewhere? To what extent does cohabitation influence the timing of marriage and the transition to parenthood? In light of ongoing policy efforts to promote family formation in Japan, answers to these questions are also of tremendous practical importance.

Cohabitation, marriage, and fertility

Recognizing that cohabitation, marriage, and childbearing are interrelated components of an increasingly complex and heterogeneous family formation process in the U.S. and Europe, recent studies have modeled the three outcomes simultaneously (Baizán, Aassve, and Billari 2001, 2004; Brien, Lillard, and Waite 1999; Musick 2006). In Japan, however, almost all childbearing takes place within marriage (2% of births were registered to unmarried mothers in 2005). We are therefore primarily interested in assessing the extent to which cohabitation experience influences the transition to first marriage. In the only previous study to address this question, Tsuya (2006) found cohabitation experience to be associated with significantly higher rates of marriage. We extend Tsuya's (2006) research by providing a fuller elaboration of the potential mechanisms

linking cohabitation to the timing of first marriage, making fewer assumptions about the timing of cohabiting unions, and considering the implications of unobserved correlates of both cohabitation and marriage timing.

Drawing on related research in the U.S. and an understanding of the trend toward later and less marriage in Japan, we consider four mechanisms through which cohabitation may contribute to early marriage: (a) lower financial barriers to marriage, (b) lower psychological barriers to marriage, (c) a larger pool of “acceptable” partners, and (d) a higher likelihood of pregnancy. We also discuss reasons to expect that the emergence of cohabitation may contribute to further delays in marriage. Finally, we consider the importance of the role of non-random selection into cohabiting unions.

Although most young Japanese men and women intend to marry at some point (National Institute of Population and Social Security Research 2004), it is clear that many perceive important barriers to marriage at young ages. For example, a large proportion cites lack of money and adequate housing as important reasons for remaining single (National Institute of Population and Social Security Research 2004). The high perceived financial barriers to marriage in Japan have been linked to the large proportion of unmarried Japanese men and especially women who remain in the parental home while postponing marriage (Raymo 2003; Yamada 1999). Psychological barriers are also important, with the trend toward later and less marriage linked to skepticism about the benefits of marriage, particularly an increasing distaste among women for the “package” of onerous roles and obligations that typically accompany marriage and motherhood (Rindfuss et al. 2004; Tsuya and Mason 1995).

Perhaps reflecting these high material and psychological barriers to marriage, it is clear that a large proportion of young Japanese men and women find it difficult to locate a “suitable”

marriage partner (NIPSSR 2004). The relatively low prevalence of romantic partnerships in Japan (Iwasawa 2004) has been linked to a decline in facilitated marriages (Retherford, Ogawa, and Matsukura 2001), including workplace marriages (Iwasawa and Mita 2005), mismatches in the marriage market (Raymo and Iwasawa 2005), and mismatches between normative expectations of a gender-asymmetric division of labor among spouses and increasing economic uncertainty at young ages (Yamada 1996, 2006).

In this context, cohabitation may facilitate early marriage to the extent that it is perceived as a more egalitarian arrangement (South and Spitze 1994) offering many of the benefits of marriage (e.g., coresidential intimacy, sexual access, economies of scale) but without either the normative assumptions of long-term emotional and financial commitment or the legal barriers to dissolution that accompany marriage. Cohabitation may also facilitate marriage if cohabiting couples are less likely to adhere to the relatively restrictive pairing norms that characterize marriage (Blackwell and Lichter 2000), and if couples are able to evaluate their compatibility more efficiently in cohabiting unions than in extended non-coresidential dating relationships. In the context of increasing marital instability (Raymo, Iwasawa, and Bumpass 2004), cohabitation may also facilitate marriage by serving as an attractive, low-commitment trial marriage for young Japanese couples who, like their counterparts in the U.S., believe it is important to assess their compatibility before marrying (Bumpass, Sweet, and Cherlin 1991). The role of cohabitation as a trial marriage may be particularly important given widespread exposure to messages about the economic, physical, and psychological difficulty of marriage and childrearing (NIPSSR 2007).

Cohabitation may also promote marriage by increasing the risk of premarital conception (Brien, Lillard, and Waite 1999; Manning and Landale 1996) and subsequent marriage prior to

childbirth (Manning 1993, 1995) or following childbirth (Manning and Smock 1995). This may be particularly relevant in Japan, where nonmarital childbearing remains extremely rare but pregnancy is an increasingly important reason for marriage. Unlike the U.S., where “shotgun marriages” have been described as a “relic of the past” (Brown 2005), *dekichattakon* (marriages preceded by pregnancy) accounted for 17% of first marriages in Japan during the 1990s (Raymo and Iwasawa 2007). We expect that cohabitation experience may hasten marriage if it increases the risk of pregnancy as a result of more frequent sexual activity (Bachrach 1987) or less vigilant contraception. Even if cohabitation is unrelated to the risk of premarital pregnancy, it is conceivable that the presence of a partner with whom they are willing to live may decrease the relative likelihood that cohabiting women choose to abort an unplanned pregnancy and hence increase the likelihood of marriage. This may be particularly true for early marriages, given evidence of recent increases in abortion at young ages (Sato and Iwasawa 2006).

These hypothesized linkages between cohabitation and earlier marriage are predicated on the assumption that a substantial proportion of cohabiting unions result in marriage. However, if cohabiting unions tend to exacerbate, rather than allay, couples’ concerns about their relationships (or about marriage more generally), the emergence of cohabitation may contribute to further delays in marriage by extending the spouse search process (Oppenheimer 1988). That is, assuming that investments of time and emotional resources are greater in cohabiting unions than in non-coresidential dating relationships, the duration between dissolution of one relationship and initiation of another relationship may be longer for those who left cohabiting unions. Alternatively, it is possible that cohabitation may contribute to later marriage by influencing attitudes in ways that may not be conducive to marriage and childbearing (Axinn and Barber 1997) or by providing couples with an attractive, long-term alternative to marriage.

Assuming that cohabiting unions do not emerge as an increasingly common context for childbearing in Japan (as they have in the U.S. and many European countries), this scenario also suggests further decline in fertility rates.

Any attempt to evaluate relationships between cohabitation experience and the transition to first marriage must address the role of non-random selection into cohabitation. Research on cohabitation in the U.S. has repeatedly shown that those who form cohabiting unions differ systematically from those who do not with respect to religiosity, family attitudes, and relationship skills (Clarkberg, Stolzenberg, and Waite 1995; Thornton, Axinn, and Hill 1992). These differences have received a great deal of attention in studies addressing relationships between cohabitation experience and subsequent marital stability. For example, Lillard, Brien, and Waite (1995) estimate simultaneous models for premarital cohabitation and marital dissolution to demonstrate that self-selection of more divorce-prone individuals into cohabitation explains the theoretically unexpected positive relationship between cohabitation and subsequent marital dissolution. Brüderl, Diekmann, and Engelhardt (1999) find similar results using German data. We expect that self-selection into cohabiting unions may be particularly relevant for understanding relationships between cohabitation and marriage timing in Japan where nonmarital coresidential unions remain a decidedly non-normative arrangement.

Data and Methods

Sample

We use data from the 1st Survey on Population, Family, and Generations (SPFG), a nationally representative survey of 2,421 women aged 20-49. The SPFG was conducted in April 2004 by the Population Problems Research Council of the Mainichi Shimbun and had a response rate of 61%. The survey contained three questions about cohabitation experience. The first asked

respondents if they had ever cohabited, with response options of “no,” “yes, in the past,” and “yes, currently cohabiting.” The second asked the length of the most recent (or ongoing) cohabiting union, with ten categorical response options ranging from “less than one month” to “five years or more.” The third asked respondents whether they married their most recent cohabiting partner or intend to marry their current partner. Using these data, we are able to calculate basic descriptive characteristics of cohabiting unions in Japan, examine the correlates of these characteristics of cohabiting unions, and evaluate the extent to which the timing of first marriage differs for those who have and have not lived in a cohabiting union.

Basic descriptive characteristics

We begin by constructing descriptive characteristics similar to those presented by Heuveline and Timberlake (2004). The percent ever cohabiting refers to both current and past unions, mean duration is a weighted average of the midpoints of the ten duration categories, and the proportion ending in marriage refers to completed cohabiting unions.³ We provide tabulations of each of these characteristics by respondents’ birth cohort. This basic information represents a major step in our empirical understanding of cohabitation in Japan and allows for comparison with other industrialized countries.

Correlates of cohabitation

To examine the correlates of cohabitation, we regress these three basic characteristics of cohabiting unions on a series of sociodemographic and attitudinal variables. We estimate logistic regression models to examine the correlates of cohabitation experience and the

³ For the open-ended duration category (five or more years), we arbitrarily assign a value of seven years to calculate mean union duration. Seven percent of women with cohabitation experience reported unions of five or more years.

likelihood that completed unions ended in marriage and estimate a linear regression model for the duration of cohabiting unions. Each of these models includes the following variables: a linear measure of *birth cohort*, a five-category measure of *respondent's education* (less than high school, high school, vocational school, junior college, university); a four category measure of *first job after school* (sales/clerical, professional, part-time or no job, other/missing); a measure of *father's education* (years of schooling plus a dichotomous indicator of missing values), a dichotomous indicator of *rural background* (lived in a rural area during elementary school); and a linear measure of *number of siblings* (topcoded at 4+).

We also include two attitudinal indices. The first is a *liberal attitudes index* constructed by summing responses to a series of Likert-type questions about the degree to which respondents (dis)approve of married couples having different last names, the use of surrogate mothers, homosexual couples, and putting children under three in daycare so that mothers can work. This index ranges in value from 4-16 with higher scores indicating more liberal (approving) attitudes. The second is a *marital attitudes index* constructed by summing responses to questions about the degree to which respondents (dis)agree that never marrying is a desirable life course, couples should marry if they get pregnant, it is important to have individual goals, individual pursuits should be sacrificed for family, men should work while women take care of the home, married couples should have children, couples should not divorce just because they don't get along, and love and marriage are two different things. This measure ranges from 8 to 32, with higher values indicating less conservative attitudes toward marriage.

Cohabitation and marriage timing

Our ability to estimate relationships between cohabitation experience and the transition to first marriage is limited by the fact that the SPFG did not collect information on the beginning and

end dates of cohabiting unions. Without information on the timing of transitions into and out of cohabiting unions, we are unable to model transitions to cohabitation and first marriage as competing risks or to treat cohabitation as a time-varying covariate in models of first marriage rates. Treating cohabitation as a time-invariant characteristic in (as in Tsuya 2006) is also problematic. The lack of information forces us to make untestable assumptions about the timing of cohabitation experience and we discuss these limitations further at the end of the paper.

Our preferred analytical strategy is to begin by estimating separate models for cohabitation experience and several dichotomous measures of the timing of first marriage. Recognizing that marriage is rare while enrolled in school and that educational differences in age at marriage are large (Raymo 2003), we begin by calculating the cumulative proportion of women in each educational group who have married by a given duration since completing school. We then classify respondents within each of the five educational categories as marrying very early, early, or “on time” if their marriage puts them in the first decile, first quartile, or first half of these education-specific cumulative distributions. For example, very early marriages for high school graduates are those that took place within 30 months of completing school (approximately 21 years old). The corresponding figures for other educational categories are 53 months for those who did not complete high school (approximately 20 years old), 30 months for graduates of vocational school (approximately 23 years old), 35 months for junior college graduates (approximately 23.5 years old), and 24 months for university graduates (approximately 24.5 years old). Similar measures are constructed, by educational attainment, for marriage within the first quartile and first half of marriages.

Each of these dichotomous marital outcomes is analyzed via the estimation of probit regression models that include cohabitation experience as a covariate. The analytic sample for

each model is limited to respondents for whom the duration from school completion to the survey date is at least as long as the duration used to calculate the measures of marriage timing. Simple versions of these models allow us to assess the extent to which cohabitation is associated with the timing of first marriage net of other relevant covariates available in the SPFG. These other covariates include birth cohort, father's education, first job, rural upbringing, and sibship size (all defined above).

We can express the models for cohabitation experience and first marriage timing generally as follows.

$$c_i^* = \sum_{k=0}^K \beta_k x_{ik} + \varepsilon_i^c, \quad (1)$$

where c_i^* is the latent propensity to cohabit for respondent i , x_{ik} is the value of the k^{th} covariate for respondent i , β_k is the coefficient to be estimated for the k^{th} covariate, and ε_i^c is a standard normally-distributed residual term. We observe either $c_i=1$ (i.e., respondent cohabited) if $c_i^* > 0$ or $c_i=0$ (i.e., respondent did not cohabit) if $c_i^* \leq 0$.

$$m_i^* = \sum_{k=0}^K \beta_k z_{ik} + \alpha c_i + \varepsilon_i^m, \quad (2)$$

where m_i^* is the latent propensity for respondent i to experience the marital outcome of interest, z_{ik} is the value of the k^{th} covariate for respondent i , β_k is the coefficient to be estimated for the k^{th} covariate, c_i is the dichotomous indicator of cohabitation experience, α is the associated coefficient, and ε_i^m is a standard normally-distributed residual term. Again, we observe either $m_i=1$ (i.e., respondent experienced marital event of interest) if $m_i^* > 0$ or $m_i=0$ (i.e., respondent did not experience marital event of interest) if $m_i^* \leq 0$.

A basic assumption of these models is that the error terms are uncorrelated with the independent variables. However, evidence of selective entry into cohabitation suggests that this assumption is unlikely to hold for equation (2). To the extent that unobserved attitudes, personality traits, and behaviors associated with cohabitation experience are also associated with the likelihood of experiencing the marital outcome of interest, correlation between ε_i^m and ε_i^c will result in biased estimates of α . For example, if young Japanese women who cohabit tend to be less committed to the institution of marriage, like their American counterparts (e.g., Clarkberg, Stolzenberg, and Waite 1995), we might also expect cohabitators to be less likely to marry within a given period of time since the completion of school. In this case, the estimated value of α will be downwardly biased by the negative correlation between ε_i^m and ε_i^c .

To account for this potential endogeneity, we estimate equations (1) and (2) simultaneously as a bivariate probit model. This approach is similar to that adopted in the research on interrelated family formation processes and the relationship between cohabitation and subsequent marital dissolution cited above. These models allow us to assess the extent to which estimated relationships between cohabitation and first marriage are consistent with the theoretical expectations outlined above and the extent to which they reflect self-selection into cohabitation of women who are more or less likely to experience different marital outcomes for unobserved reasons.

Although not necessary for the identification of bivariate probit models with endogenous regressors, the inclusion of instrumental variables should contribute to more robust estimates of α than those identified only by functional form (Wilde 2000). However, identification of adequate instrumental variables is difficult for both theoretical and practical reasons. Because cohabitation and marriage are both coresidential, romantic unions that are frequently modeled as

competing risks, it is difficult to identify variables that are correlated with the decision to cohabit but not with marriage timing. At a more practical level, our job is complicated by the fact that the SPFG questionnaire contained a limited number of questions. We take advantage of the fact that our measures of marriage timing are unrelated to respondents' educational attainment - a strong predictor of cohabitation experience (Table 2). The fact that the education-specific probabilities of experiencing the three marriage outcomes of interest are identical by construction provides us with one valid instrumental variable. Furthermore, recognizing that cohabitation remains a non-normative type of union in Japan, we were also able to identify a subset of the attitudinal items described above that were related to cohabitation but unrelated to the indicators of marriage timing. In particular, we use an index constructed from responses to questions about surrogate mothers, homosexual couples, and employment among mothers of preschool age children.

Results

Basic descriptive characteristics

In Table 1, we present basic descriptive characteristics of cohabiting unions reported in the SPFG. The first row shows that 15% of the sample has ever cohabited, that the mean duration of cohabiting unions is slightly less than two years, and that a little more than half (58%) of completed unions resulted in marriage.⁴ Tabulations by birth cohort show a substantial increase

⁴ The proportion of completed cohabiting unions resulting in marriage will be biased upward to the extent that women in this sample have experienced multiple cohabiting unions. We suspect that serial cohabitation is uncommon in Japan but have no way of evaluating this assumption. It is also higher than would be the case if we limited the tabulations to women currently in their first marriage. The responses of formerly married women suggest that post-divorce cohabitation

in cohabitation experience among women born after 1965. One in five women born in the 1970s (i.e., women age 24-34 at the time of the survey) reported ever living in a cohabiting union. This figure is far lower than in the U.S. (Bumpass and Lu 2000) and most European countries (Kiernan 2001) but represents a substantial change in the family formation process in Japan. The apparent decline in mean duration for cohorts born after 1970 reflects the relatively limited opportunity to experience cohabiting unions of five or more years.⁵ The sharp decline in the proportion of completed unions resulting in marriage for cohorts born after 1975 reflects two different processes. The first is a matter of timing, with some of these younger cohorts reporting that they intend to marry, but have not yet married, their former cohabiting partner. There also appears to be a real decline in the likelihood that cohabiting unions result in marriage.

In sum, the figures in Table 1 indicate that cohabitating unions in Japan are relatively uncommon, short in duration, and have a somewhat lower than average likelihood of resulting in marriage in comparison with countries included in recent cross-national studies by Heuveline and Timberlake (2004) and Kiernan (2001). At the same time, it is clear that cohabitation is an increasingly common experience in Japan. In Figure 1, we provide a striking picture of how the emergence of cohabitation, along with increases in bridal pregnancy (Raymo and Iwasawa 2007), is transforming the family formation process in Japan. Based on SPFG data, this figure presents the distribution of marriages in three different time periods (1980-89, 1990-99, and 2000-04) by premarital cohabitation experience and pregnancy status at marriage. The decline in the conventional pathway to marriage is remarkable. While the large majority (79%) of women

may be relatively common and that these unions have a relatively high likelihood of resulting in remarriage but we are not able to confirm this with the available data.

⁵ Median duration is 15 months for all cohorts born between 1965 and 1979.

marrying in the 1980s had never cohabited and were not pregnant at marriage only slightly more than half (55%) of women marrying during 2000-04 followed this family formation trajectory. Almost one-third had cohabited prior to marriage and one-third of those who had cohabited were pregnant at the time of marriage. This is very substantial change in a relatively short period of time.

Correlates of cohabitation

In Table 2, we describe relationships between the three basic characteristics of cohabiting unions and several sociodemographic and attitudinal characteristics available in the survey. After describing the means and distributions of the covariates in column 1, we present coefficients from a logistic regression model of ever cohabiting (column 2), coefficients from a linear regression model of cohabitation duration (column 3), and coefficients from a logistic regression model of marrying subsequent to cohabitation (column 4).

Looking first at the likelihood of ever cohabiting (column 2), the positive coefficient for birth cohort describes the increase in cohabitation over time already described in Table 1. Like Tsuya (2006), we find large educational differences in cohabitation experience, with the least educated most likely to cohabit and the most educated least likely to cohabit. The likelihood of cohabitation is positively related to sibship size (at $p < .10$) but is not significantly related to father's education, rural background, or first job after school. The two attitudinal indices are, not surprisingly, strongly related to cohabitation experience. Although it is not possible to distinguish the influence of attitudes on the likelihood of cohabitation from the influence of cohabitation on attitudes with cross-sectional data like these, it is clear that women with more liberal attitudes were significantly more likely to report living in a cohabiting union. For

example, women who were one standard deviation above the mean on the liberal (marital) attitudes index are 29% (17%) more likely to report having cohabited.

In column 3, the only statistically significant relationship indicates that the daughters of more highly educated men tend to have longer cohabiting unions. The results in column 4 indicate that cohabiting unions are less likely to result in marriage for women in more recent birth cohorts, those who did not complete high school, those with more educated fathers, and those with more liberal attitudes.

Relationships between cohabitation and first marriage

Tables 3-5 present the results of models for each of the three marital outcomes. In each of these tables, coefficients for the simple cohabitation model are presented in the first column, coefficients for the simple marriage model are presented in the second column, and coefficients for the bivariate probit model are presented in the third and fourth columns.

Table 3 presents results of models for very early marriage. Given our definition of very early marriage, we exclude 97 women for whom the duration between completion of school and the survey date was shorter than the education-specific durations by which 10% of the sample had married. The first column indicates that cohabitation is more common among more recent cohorts, women who report having a part-time job or no job after completing school, and women who did not report their father's educational attainment. By construction, the liberal family attitudes index is also positively associated with the probability of cohabitation. Educational attainment is a strong predictor of cohabitation experience, with junior high school graduates 10% more likely and women with at least a two-year college degree 4-5% less likely than high school graduates to report living in a cohabiting union.

Column 2 shows that the probability of very early marriage is lower among more recent birth cohorts and daughters of more highly educated men but higher among women who grew up in rural areas and have more siblings. Relative to women whose first post-school job was in a semi-skilled occupation (clerical, sales, service, manual labor), very early marriage is 6-7% more likely among those who reported that their first post-school job was either part-time/not working or other/missing. The coefficient of primary interest in this model indicates that the experience of cohabitation is associated with a significantly higher likelihood of early marriage. More specifically, women who cohabited are 5% more likely to marry very early relative to those who did not cohabit.

The strength of this relationship increases dramatically when we allow for correlation between the errors of the two equations. The large negative correlation between the error terms for the two equations ($\rho = -.59$) indicates the presence of unobserved characteristics that are positively associated with the propensity to cohabit but negatively associated with the propensity to marry very early. More specifically, we see that cohabitation experience is now associated with a 42% higher likelihood of being among the first 10% of one's educational group to marry. Once we account for the fact that cohabitation is selective of women who are less likely to marry very young, the estimated relationship between cohabitation and early marriage is consistent with a scenario in which cohabitation hastens marriage. As discussed above, cohabitation might promote earlier marriage by increasing in the likelihood of premarital pregnancy or by lowering the barriers to marriage in a setting where the perceived benefits of marriage at young ages are relatively limited.

Results are similar when we consider the less extreme outcome of being among the first 25% to marry. These results, presented in Table 4, are based on the subsample of 1,858 women

for whom the duration between completion of school and the survey date was shorter than the education-specific durations within which 25% of the sample had married. Significant correlates of cohabitation (in the first column) are the same as in Table 2. In the simple marriage model, results are similar to those for very early marriage except that the coefficients for rural background and number of siblings are no longer significant. The most important difference is that the estimated coefficient for cohabitation experience in this model is small in magnitude and statistically insignificant. Net of the other covariates in this model, cohabitation experience is unrelated to the likelihood of marrying early.

As in Table 2, the bivariate probit model shows that the correlation between the error terms of the two equations is large and negative ($\rho = -.51$) and that the downward bias introduced by self-selection into cohabitation remains important for the earliest 25% to marry. Once we control for unobserved factors related to the relatively low propensity of early marriage among women who have cohabited, the magnitude and significance of the cohabitation coefficient increases substantially. Cohabitation experience is associated with a 39% increase in the probability of being within the first 25% to marry. Estimates of other coefficients are largely unchanged in the bivariate probit model.

Results for the third set of models, for being among the first 50% to marry, are very similar. These models, presented in Table 5, are based on the subsample of 1,688 women for whom the duration between completion of school and the survey date was shorter than the education-specific durations by which 50% of the education-specific samples had married. Results for the simple cohabitation and marriage models are similar to those for earlier marriages. One exception is that limited labor force attachment following school is no longer associated with relatively early marriage. Importantly, the coefficient for cohabitation is now essentially

zero. On average, cohabitation is unrelated to the probability of being among the first half of women to marry.

However, results again change dramatically when we estimate the two equations simultaneously. The estimated correlation between the residual terms of the two equations ($\rho = -.60$) indicates a strong concentration of unobserved characteristics associated with later marriage among those women who report having been in a cohabiting relationship. Once we account for the presence of these unobserved characteristics, the estimated coefficient for cohabitation experience becomes large, positive, and statistically significant. As in the previous models, we see that cohabitation itself contributes to earlier marriage.

Discussion

In contrast to the rapidly growing body of research on cohabitation in the U.S. (Seltzer 2000; Smock 2000) and Europe (Kiernan 2002), nonmarital unions have received extremely little attention in studies of family formation in Japan. This likely reflects the relatively low prevalence of cohabitation as well as the lack of suitable data with which to construct even the most basic demographic measures of cohabitation. This is an important limitation given major policy emphasis on the promotion of marriage and fertility in Japan as well as the potential relevance of the Japanese experience for evaluating the generality of findings from research on cohabitation in low-fertility western societies.

In this paper, we used data from one of the only existing sources of nationally representative data on cohabitation to begin filling this large gap in our understanding of the family formation process in Japan. These data allowed us to provide a first look at the characteristics of cohabitating unions in Japan, trends and correlates of these characteristics, and linkages between cohabitation and the transition to first marriage. We demonstrated that

cohabitation has increased rapidly among recent cohorts of women and that cohabitating unions in Japan tend to be relatively short and almost as likely to dissolve as to end in marriage. We have also shown that educational attainment and attitudes are strongly associated with cohabitation experience. Based on simple models, we concluded that cohabitation is only weakly related to the likelihood of early marriage. These simple relationships mask important heterogeneity, however. Controlling for the spurious influence of a strong negative correlation between the propensity to cohabit and the propensity to marry early (for reasons not observable in the data), cohabitation experience itself was associated with a substantially higher likelihood of relatively early marriage. In sum, our results suggest that cohabitation is a rather selective experience in contemporary Japan and that an understanding of the role of cohabitation in the family formation process must account for this selectivity.

Our results provide an empirical basis for interesting speculation. If cohabitation in Japan, like cohabitation in the U.S. and elsewhere, rapidly emerges as a more common, less selective type of union, what are the implications for marriage and fertility? Naïve extrapolation of our findings suggests that decline in the selectivity of cohabitation may contribute to an increase in earlier marriage. This would presumably have positive implications for fertility as well, both through the promotion of marriage and the increased likelihood of premarital pregnancy. As a relatively low-risk option that involves exchanging the comforts of the parental home for coresidence with an intimate partner, cohabitation may contribute to a reversal of the long-term trend toward later marriage by reducing perceived barriers to marriage. This kind of speculation is obviously predicated on assumptions that estimated coefficients for cohabitation in our bivariate probit models are robust to alternative model specifications and remain relatively

stable over time. Both of these assumptions should be carefully evaluated in subsequent research based on richer data.

It is certainly possible that a decline in the selectivity of cohabiting unions may be accompanied by change in the relationship between cohabitation and marriage, with cohabitation emerging as an alternative to marriage that may or may not involve childbearing. It is also possible that changes in the relationship between premarital pregnancy and marriage, combined with further growth in the prevalence of cohabiting unions, could lead to either an increase in abortions or nonmarital births rather than increased marriage rates. Finally, the emergence of cohabitation may also contribute to aggregate-level changes in family attitudes in ways that may not be conducive to marriage and childbearing (Axinn and Thornton 1993).

While we are speculating about the potential implications of further increases in the prevalence of cohabitation in Japan, it is useful to also speculate about the likelihood that cohabitation will indeed emerge as a more common component of the family formation process. In addition to the empirical regularity with which cohabitation has spread in other late-marriage, low-fertility societies, there are several reasons to expect that cohabitation may become more common in Japan. The first is the powerful role of social networks and diffusion processes. In light of evidence that half of adult Japanese know someone who has cohabited and that disapproval of cohabitation may be weaker than previously thought (Rindfuss et al. 2004), we suspect that normative barriers to cohabitation are waning. The SPFG survey provides more direct evidence of the potential for further increases in cohabitation. When asked if they would like to cohabit in the future, only a minority of never married respondents said no. One-third said yes while another one-fourth was unsure.

There are many limitations to this study that we hope can be addressed in subsequent research based on different data. The most important is the limited information collected on cohabitation in the SPFG. As a result, we were unable to model cohabitation and marriage as competing risks or to treat cohabitation as a time-varying covariate in models of first marriage timing. Our solution of estimating bivariate probit models for cohabitation experience and dichotomous indicators of marriage timing simplifies the data problems but does not resolve them. For example, among respondents who did not marry by the cutoff dates used to define our marriage timing variables, it is certainly possible that entry into cohabitation occurred after that date. In such cases, cohabitation experience obviously cannot influence marriage timing but the decision to cohabit may be influenced by delayed marriage. For the sake of simplicity, we have discussed the “effect” of cohabitation on marriage timing, but the estimated coefficients for cohabitation should clearly be viewed as associations rather than causal effects.

Without information on the beginning and ending dates of all cohabiting unions, it is also impossible to determine the temporal ordering of pregnancy and union entry. Additionally, it is difficult in some cases to determine whether the cohabiting union reported directly preceded the marriage we observe. Furthermore, because we have information only on respondents’ current marriage, we cannot examine the relationship between cohabitation experience and first marriage timing for women who have separated from their first spouse. Finally, the absence of information on aborted pregnancies prevents an examination of the extent to which women in cohabiting unions are more likely than their unpartnered counterparts to legitimate premarital pregnancies. Given that at least one-fifth of Japanese women will live in a cohabiting union, collecting data that allow researchers to address these questions should be a priority.

In this paper, we focused only on the relationship between cohabitation and the timing and nature of first marriage. This decision was motivated by the very strong relationship between marriage and childbearing in Japan. However, evidence of increasing delay in transition to parenthood within marriage (Sasai 2004) suggests a potentially important role for cohabitation. Although we have no reason to expect that cohabitation will emerge as an important context for childbearing in Japan, our results do suggest that cohabitation likely hastens first marital birth by increasing the likelihood of marriage in response to pregnancy. This relationship may be muted in simple models of marital fertility that do not take account of selectivity into cohabiting unions and the potential association of these unobserved factors with the risk of first marital birth among women who do not marry while pregnant. Another important subject for future study is to extend research from the U.S. and elsewhere on relationships between cohabitation experience and other marital outcomes such as divorce. Divorce in Japan has increased substantially over time (Raymo, Iwasawa, Bumpass 2004) and the role of cohabitation (if any) in this trend remains unexplored. We are not aware of any source of data appropriate for this kind of analysis, however.

We conclude by emphasizing that the results in this paper, along with evidence from related research on divorce (Raymo, Iwasawa, and Bumpass 2004) and bridal pregnancy (Raymo and Iwasawa 2007), highlight that conditions are not only “ripe for substantial change in family behavior in Japan” (Rindfuss et al. 2004:854) but also that substantial change is already underway. By taking advantage of the distinctive social, political, and economic context in which these changes are taking place, research on ongoing family change in Japan may provide a rich source of insight into the generality of patterns observed in the U.S. and other Western societies and the ways in which local contexts shape the pace and nature of family change.

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Table 1: Cohabitation experience, duration, and outcomes, by birth cohort

	Prevalence of cohabitation experience	Mean duration of cohabiting unions (months)	Proportion of completed cohabiting unions resulting in marriage
Total	15.01	21.46	57.67
<i>Birth cohort</i>			
1954-59	9.82	21.7	64.29
1960-64	10.4	20.7	69.57
1965-69	17.26	26.27	61.64
1970-74	21.33	19.7	60.71
1975-79	20.73	20.69	39.68
1980-84	9.63	15.93	44.44

Table 2: Multivariate analyses of cohabitation experience, duration, and outcomes

<i>Variable</i>	Pct./Mean	(s.d.)	Prevalence of cohabitation experience ^a	Mean duration of cohabiting unions (months) ^b	Proportion of completed cohabiting unions resulting in marriage ^a
<i>Birth cohort</i>	1967.66	(8.09)	0.023 **	-0.284	-0.041 *
<i>Educational Attainment</i>					
Junior high school	0.03		1.084 **	4.255	-1.116 *
High school (omitted)	0.41		0.000	0.000	0.000
Vocational school	0.17		-0.099	2.646	-0.567
Junior college	0.23		-0.608 **	-2.543	-0.402
University	0.16		-0.680 **	-0.675	-0.058
<i>First job after completing education</i>					
Part-time/no job	0.13		0.227	-1.689	0.206
Other, missing	0.08		-0.066	6.044	0.284
Professional	0.21		-0.014	-0.109	0.419
Sales/clerical (omitted)	0.58		0.000	0.000	0.000
<i>Father's educational attainment</i>	11.59	(2.74)	0.034	1.013 *	-0.181 **
<i>Father's education missing</i>	0.06		-0.118	-0.899	0.212
<i>Rural residence in childhood</i>	0.28		0.097	-1.883	-0.070
<i>Number of siblings</i>	1.52	(0.88)	0.130 #	0.976	0.203
<i>Liberal attitudes index</i>	9.84	(2.59)	0.098 **	-0.259	-0.166 **
<i>Marriage attitudes index</i>	20.39	(3.26)	0.048 *	-0.102	0.021
Constant			-4.084 **	29.718 **	2.012
n	2398		2323	344	319
(Pseudo) R2			0.05	0.03	0.11

**p < .01, *p < .05, #p < .10

Notes: a) estimated log-odds ratios from logistic regression model, b) estimated coefficients from ordinary least squares regression model

Table 3: Probit coefficients and marginal effects for models of cohabitation and very early marriage

<i>Variable</i>	Separate models				Bivariate Probit			
	Cohabitation ^a		Very early marriage ^b		Cohabitation ^a		Very early marriage ^b	
	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect
Birth cohort	0.021 **	0.004	-0.200 **	-0.003	0.022 **	0.005	-0.023 **	-0.005
<i>Educational Attainment</i>								
Junior high school	0.408 *	0.103			0.300	0.072		
High school (omitted)	0.000				0.000			
Vocational school	-0.054	-0.011			-0.084	-0.017		
Junior college	-0.243 *	-0.047			-0.243 *	-0.046		
University	-0.226 #	-0.042			-0.209 #	-0.039		
<i>First job after completing education</i>								
Part-time/no job	0.237 *	0.054	0.327 *	0.067	0.264 *	0.060	0.243 *	0.053
Other, missing	0.034	0.007	0.290 *	0.060	0.036	0.008	0.260 #	0.058
Professional	0.009	0.002	0.083	0.015	0.027	0.006	0.093	0.019
Sales/clerical (omitted)	0.000		0.000		0.000		0.000	
<i>Father's educational attainment</i>	0.015	0.003	-0.025 #	-0.004	0.013	0.003	-0.023	-0.005
<i>Father's education missing</i> ^c	-0.443 *	-0.071	0.337 #	0.072	-0.473 *	-0.074	0.369 *	0.087
<i>Rural residence in childhood</i> ^c	0.129	0.027	0.170 *	0.025	0.141 #	0.030	0.103	0.021
<i>Number of siblings</i>	0.067	0.014	0.111 *	0.020	0.074 #	0.015	0.091 *	0.018
<i>Liberal attitudes index</i> ^d	0.066 **	0.014			0.063 **	0.013		
<i>Ever cohabited</i>			0.239 *	0.047			1.397 *	0.422
Constant	-1.965 **		-1.354 **		-1.982 **		-1.335 **	
rho					-0.594			
p for chisquare test that rho = 0					0.171			
n	1978		1978		1978			
Log-likelihood	-750		-657		-1406			

**p < .01, *p < .05, #p < .10

Notes: a) 0=did not cohabit, 1=cohabited, b) 0=not among first 10% to marry, 1=among first 10% to marry, c) reference category is no, d) Constructed from responses to questions about surrogate mothers, gay marriages, and mothers' employment when children are not yet in school.

Table 4: Probit coefficients and marginal effects for models of cohabitation and early marriage

<i>Variable</i>	Separate models				Bivariate Probit			
	Cohabitation ^a		Early marriage ^b		Cohabitation ^a		Early marriage ^b	
	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect
Birth cohort	0.028 **	0.006	-0.030 **	-0.010	0.030 **	0.006	-0.034 **	-0.012
<i>Educational Attainment</i>								
Junior high school	0.501 *	0.131			0.444 *	0.113		
High school (omitted)	0.000				0.000			
Vocational school	-0.040	-0.008			-0.054	-0.011		
Junior college	-0.256 *	-0.048			-0.234 *	-0.044		
University	-0.228 #	-0.042			-0.196	-0.037		
<i>First job after completing education</i>								
Part-time/no job	0.299 *	0.069	0.177 #	0.060	0.330 **	0.077	0.085	0.029
Other, missing	0.026	0.005	0.276 *	0.096	0.059	0.012	0.254 *	0.090
Professional	0.011	0.002	0.013	0.004	0.020	0.004	0.021	0.007
Sales/clerical (omitted)	0.000		0.000		0.000		0.000	
<i>Father's educational attainment</i>								
Father's education missing ^a	-0.416 #	-0.067	0.128	0.043	-0.424 #	-0.068	0.165	0.057
<i>Rural residence in childhood^a</i>								
Number of siblings	0.154 #	0.033	0.086	0.028	0.145 #	0.031	0.053	0.018
Liberal attitudes index ^b	0.052	0.011	0.032	0.010	0.058	0.012	0.019	0.006
<i>Ever cohabited</i>								
Constant	-2.028 **		0.085	0.028	-2.051 **		1.030 #	0.385
rho					-0.507			
p for chisquare test that rho = 0					0.153			
n	1858	1858			1858			
Log-likelihood	-698	-1048			-1745			

**p < .01, *p < .05, #p < .10

Notes: a) 0=did not cohabit, 1=cohabited, b) 0=not among first 25% to marry, 1=among first 25% to marry, c) reference category is no, d) Constructed from responses to questions about surrogate mothers, gay marriages, and mothers' employment when children are not yet in school.

Table 5: Probit coefficients and marginal effects for models of cohabitation and "on time" marriage

<i>Variable</i>	Separate models				Bivariate Probit			
	Cohabitation ^a		"On time" marriage ^b		Cohabitation ^a		"On time" marriage ^b	
	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect	Coefficient	Marginal effect
Birth cohort	0.036 **	0.007	-0.040 **	-0.016	0.038 **	0.007	-0.046 **	-0.018
<i>Educational Attainment</i>								
Junior high school	0.650 **	0.177			0.595 *	0.158		
High school (omitted)	0.000				0.000			
Vocational school	-0.061	-0.012			-0.109	-0.021		
Junior college	-0.271 *	-0.050			-0.263 *	-0.048		
University	-0.206	-0.037			-0.220	-0.040		
<i>First job after completing education</i>								
Part-time/no job	0.405 **	0.097	0.114	0.045	0.431 **	0.104	-0.033	-0.013
Other, missing	0.047	0.010	0.169	0.066	0.062	0.013	0.141	0.055
Professional	0.069	0.014	0.107	0.042	0.091	0.019	0.010	0.039
Sales/clerical (omitted)	0.000		0.000		0.000		0.000	
<i>Father's educational attainment</i>								
<i>Father's education missing</i> ^a	-0.402	-0.063	-0.254 **	-0.101	-0.365	-0.059	-0.180	-0.072
<i>Rural residence in childhood</i> ^a	0.167 #	0.034	0.077	0.031	0.171 *	0.035	0.029	0.012
<i>Number of siblings</i>	0.048	0.010	0.046	0.018	0.055	0.011	0.028	0.011
<i>Liberal attitudes index</i> ^b	0.060 *	0.012			0.049 #	0.010		
<i>Ever cohabited</i>			0.002	0.001			1.077 *	0.363
Constant	-1.961 **		0.412 **		-1.944 **		0.405 **	
rho					-0.604			
p for chisquare test that rho = 0					0.072			
n	1688		1688		1688			
Log-likelihood	-619		-1111		-1728			

**p < .01, *p < .05, #p < .10

Notes: a) 0=did not cohabit, 1=cohabited, b) 0=not among first 50% to marry, 1=among first 50% to marry, c) reference category is no, d) Constructed from responses to questions about surrogate mothers and mothers' employment when children are not yet in school.

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