ECONOMIC POTENTIAL AND ENTRY INTO MARRIAGE AND COHABITATION*

Yu Xie, University of Michigan

James Raymo, University of Wisconsin-Madison

Kimberly Goyette, Temple University

Arland Thornton, University of Michigan

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* Direct all correspondence to Yu Xie (e-mail: yuxie@umich.edu), Population Studies Center, Institute for Social Research, 426 Thompson Street, University of Michigan, Ann Arbor, MI 48106. An earlier version of the paper was presented at the 2001 Population Association of America annual meeting (Washington DC, March). This research was supported by a research grant from the NICHD. The authors are grateful to Frances Goldscheider for helpful comments on an earlier version of the paper.

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ABSTRACT

This study explores the role of economic potential in determining rates of entry into marriage and cohabitation. Instead of using reported earnings, which are a poor indicator of economic potential for young people, or educational attainment and employment as crude proxies, we develop a method for explicitly estimating five time-varying measures of earnings potential: current earnings, earnings over the next five years, future earnings, past earnings, and lifetime earnings. Our research on entries into marriage and cohabitation draws on data from an intergenerational panel study of parents and children, and the auxiliary work of estimating predicted earnings utilizes data from the 1990 Census 5% Public Use Microsample (PUMS) and the 1980-1992 High School and Beyond, Sophomore Cohort. Results of this research show that all five measures of earnings potential strongly and positively influence the likelihood of marriage for men, but not for women. Another important finding is that the measures of earnings potential do not affect entry into cohabiting unions for either men or women.

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The union formation of young American men and women has undergone significant change in recent decades. The age of first marriage has risen; cohabitation has become more prevalent; and it also appears that the proportion who will never marry is increasing (e.g., Bumpass, Sweet, and Cherlin 1991; Bumpass and Sweet 1989; Cherlin 1992; Manning 1995; Schoen and Weinick 1991; Sweet and Bumpass 1987; Thornton 1988). These trends in union formation have coincided with the rapid increase in women's participation in the labor force (Bianchi and Spain 1986; Spain and Bianchi 1996), prompting the question: Is women's increasing employment responsible for the trend toward later and less marriage?

This explanation for changes in marriage behavior, commonly referred to as the "economic independence" hypothesis, is based on the assumption of gender role specialization within the family (see Oppenheimer 1997 for a recent review). Although empirical evidence in its support is rather weak, this explanation has a great deal of face validity and has become a dominant paradigm for explaining recent changes in marriage behavior (Oppenheimer 1997). Assuming that an important motivation for marriage lies in gender role specialization within the family – with the wife specializing in household work and the husband specializing in market labor – the economic independence hypothesis predicts declining rates of marriage as more women participate in the labor force.

With the exception of Clarkberg (1999), Raley (1996), and Thornton, Axinn, and Teachman (1995), previous discussions of the role of economic resources in family formation have exclusively focused on marriage, ignoring cohabitation. If "cohabitation is very much a family status" (Bumpass, Sweet, and Cherlin 1991, p.926), research on family formation should also study entry to cohabitation. Whether or not the hypothesized effects of economic resources on marriage apply to entry into cohabitation is an important question because it helps us understand the differences and similarities between cohabitation and marriage. An affirmative answer to this question would lend support to the contention that cohabitation is a form, albeit a less stable form, of marriage. A negative answer would suggest a marked differentiation, at least behaviorally, between marriage and cohabitation.

Extending the earlier work of Thornton, Axinn, and Teachman (1995), this paper contributes to the literature on union formation in two important ways. First, our study considers the role of economic potential in determining the rates of entry into both marriage and cohabitation. Second, instead of using reported earnings, which are a poor indicator of economic potential for young people, or educational attainment and employment as crude proxies, we develop a method for explicitly estimating five time-varying measures of earnings potential: current earnings, earnings over the next five years, future earnings, past earnings, and lifetime earnings. These estimations are based on information pertaining to educational attainment, work experience, and cognitive ability, as well as college quality and field of study for individuals who have attained postsecondary education. Our research on entries into marriage and cohabitation is based upon the same intergenerational panel study of parents and children used by Thornton, Axinn, and Teachman (1995), with updated information on respondents' life experiences through age 31. The auxiliary estimation of predicted earnings utilizes data from the 1990 Census 5% Public Use Microsample (PUMS) and the 1980-1992 High School and Beyond, Sophomore Cohort.

THEORETICAL ISSUES

Economic Resources and Marriage

According to the economic independence hypothesis, relative improvements in women's economic position in the labor market are expected to reduce the gains to gender role specialization within marriage, thus making marriage less attractive for both women and men (e.g., Becker 1973, 1974, 1991; Goldscheider and Waite 1986; Preston and Richards 1975). Empirical evidence assessing this hypothesis has been mixed. Some studies based on aggregate-level data and cross-sectional survey data find a negative relationship between indicators of women's economic status (i.e., educational attainment, employment, earnings) and the prevalence or incidence of marriage. However, research using more appropriate longitudinal, individual-level data has typically shown the relationship between measures of women's economic status and the likelihood of marriage to be positive or, in some cases, insignificant. Oppenheimer (1997) offers a thorough review of this literature. Results from investigations of men's marriage behavior are less dependent on the nature of the data analyzed. Consistent with theoretical expectations, analyses of both cross-sectional and longitudinal data have invariably shown that greater economic resources are associated with significantly higher rates of marriage for men (e.g., Cooney and Hogan 1991; Goldscheider and Waite 1986; Lloyd and South 1996; MacDonald and Rindfuss 1981; Mare and Winship 1991; Oppenheimer, Kalmijn, and Lim 1997; Sassler and Goldscheider 1997; Sweeney 2002; Sassler and Schoen 1999; Teachman, Polonko, and Leigh 1987).

With attitudinal data providing little support for a rejection of marriage among economically independent women, Oppenheimer has proposed an alternative model of marriage timing in which the spouse search process is prolonged for women with greater economic resources (Oppenheimer 1988, 1994, 1997; Oppenheimer, Blossfeld, and Wackerow 1995; Oppenheimer and Lew 1995). In this "extended spouse search" model, greater economic resources contribute to later marriage by increasing women's incentive as well as their financial ability to conduct longer and more exacting searches in the marriage market. An important distinction between the extended spouse search model and the gender role specialization model is that the former posits a positive, rather than a negative, relationship between women's economic resources and their attractiveness as marriage partners.

Cohabitation

One of the most notable trends in family behavior in the U.S. is the rapid increase in nonmarital cohabitation. Although cohabiting unions resemble marriage in many respects and often serve as precursors to marriage (Bumpass, Sweet, and Cherlin 1991), cohabitation is ostensibly not the same as marriage. In fact, cohabitation is more an empirical operationalization than a theoretical construct, with researchers still struggling with satisfactory conceptualizations of cohabitation (see Rindfuss and Van den Heuvel 1990). Three different views of cohabitation in the literature are: (a) cohabitation as an alternative to marriage, (b) cohabitation as an alternative to being single, and (c) cohabitation as a precursor to marriage. The conceptualization of cohabitation as an alternative to marriage emphasizes the similarities between cohabiting unions and marriages (e.g., sexual intimacy, expressed commitment, shared household, and even child-bearing) and views the difference between the two as a choice of lifestyle. The conceptualization of cohabitation as an alternative to being single emphasizes the dissimilarities between cohabitation and marriage. For example, Rindfuss and Van den Heuvel (1992) show that cohabitors more closely resemble single men and women than married couples across a wide range of attitudes and family activities. The conceptualization of cohabitation as a precursor to marriage considers cohabitation as an intermediate step between being single and married but treats marriage and cohabitation as qualitatively different. This view is supported by the fact that cohabiting unions are typically short in duration and that a large proportion of cohabiting unions are followed by marriage (Bumpass and Lu 2000).

Given the ambiguity in the meaning of cohabitation, it is not surprising that the relationship between economic resources and cohabitation remains unclear (Clarkberg 1999). If cohabitation is considered an alternative to marriage, it seems reasonable to expect that economic resources positively affect entry into cohabitation, at least for men, in the same way that they influence entry into marriage. If cohabitation is viewed as an alternative to being single, then economic resources should not affect entry into cohabitation. If cohabitation is best understood as a precursor to marriage, the relationship between economic resources and entry into cohabitation is less clear. On the one hand, the effects of economic resources on pre-marital cohabitation may be similar to those on marriage. On the other hand, for some cohabiting couples planning to marry, one reason for cohabiting prior to marriage may well be the lack of sufficient economic resources for marriage (e.g., Oppenheimer 1988, p.71). This theoretical ambiguity about the nature of cohabitation suggests the need to treat marriage and cohabitation as two distinct types of union formation.

Economic Potential

As described above, there is already a large and well-researched literature on the influence of economic resources on marriage behavior. Previous studies have typically measured economic well-being using variables observed either at, or immediately preceding, marriage or cohabitation. Most prominent among such measures are current earnings (Clarkberg 1999; MacDonald and Rindfuss 1981; Mare and Winship 1991; Oppenheimer, Kalmijn, and Lim 1997;

Sweeney 2002), educational attainment (Clarkberg 1999; Goldscheider and Waite 1986; Goldstein and Kenney 2001; Mare and Winship 1991; Oppenheimer, Blossfeld, and Wackerow 1995; Oppenheimer, Kalmijn, and Lim 1997; Sweeney 2002; Thornton, Axinn, and Teachman 1995; Waite and Spitze 1981), work experience (Clarkberg 1999; Oppenheimer, Kalmijn, and Lim 1997; Sweeney 2002), employment (Goldscheider and Waite 1986; Oppenheimer, Kalmijn, and Lim 1997; Waite and Spitze 1981), and parental resources (Clarkberg 1999; Goldscheider and Waite 1986; MacDonald and Rindfuss 1981; Oppenheimer and Lew 1995; Sweeney 2002; Waite and Spitze 1981). However, these empirical measures do not closely match the intended theoretical concept of economic well-being. Theoretically, researchers are interested in measuring the concept of *perceived* long-term economic potential following marriage, as it is only post-marriage economic well-being that should have any *direct* relevance for marriage behavior. The various concurrently measured variables used in the literature should therefore be viewed as proxies of perceived long-term economic well-being.

The use of these proxy measures can be justified by the recognition that evaluation of potential mates in the marriage market is subject to a great deal of uncertainty and information asymmetry (Oppenheimer 1988). It is simply not possible for individuals to accurately assess their own future economic well-being, much less that of potential spouses. For example, current earnings at young ages are often uninformative because they can be artificially low or even zero for some individuals with high future earnings. That is, the current earnings of young people are often a poor measure (i.e., underestimation) of long-term or even short-term economic potential, because these youth may still invest in human capital accumulation--by receiving formal education in school or undertaking training--for rapid earnings growth in the future. At the same time, individuals base their union formation decisions not only on their current and past economic well-being, which is observable, but also on their expectations regarding future economic well-being, which are unobservable. This problem is further compounded by the fact that post-marriage economic behaviors of men and women can be substantially altered by marriage itself.

DATA AND METHODS

The Sample

The primary data for this research come from an intergenerational panel study of mothers and children (hereafter IPS) consisting of a probability sample of first-, second-, and fourth-born white children drawn from 1961 birth records from the Detroit metropolitan area. The mothers and their children have been interviewed periodically between 1962 and 1993, by which time the focal children had reached age 31. From eight interviews with the mothers between 1962 and 1993, the data set contains a wealth of information about the personal, social, economic, and religious circumstances of the parents as well as a history of the mothers' marriage, post-marital cohabitation, and remarriage experiences. In addition, interviews with focal children were conducted at ages 18, 23, and 31. Because these three interviews were designed to study the family formation behavior of the children, they collected extensive information about relevant experiences, including education, work, cohabitation, marriage, and childbearing from age 15 through age 31.

The original data collection in 1962 interviewed 1,113 women, representing 92 percent of the families drawn for the sample. For this study, we restrict the sample to respondents who had not entered a marriage or cohabiting relationship before age 15 and provided valid information on all of the explanatory variables, yielding a sample of 428 men and 436 women. This analytic sample suffers from only a small amount of attrition and missing data. As a result, the characteristics of the analytical sample are almost identical to those of the original sample. Note that this same sample was used by Thornton, Axinn, and Teachman (1995).

One unfortunate limitation of the IPS sample is the restriction of the population universe to white families. As a consequence, our results should not be generalized to other racial groups, since the union-formation process may be different among blacks than among whites (Bennett, Bloom, and Craig 1989; Lichter et al. 1992; Mare and Winship 1991). A second limitation of the sample is its restriction to first, second, and fourth births from the Detroit metropolitan area. The sample is, thus, a regional one that was not designed to draw inferences regarding the national population of births of all parities.

Although we do not claim that our sample is statistically representative of the entire country, there are good reasons for using the IPS for our research purpose. We are unaware of theoretical models positing that the underlying causal mechanisms of marriage and cohabitation vary across birth order or metropolitan area. While it is sometimes hypothesized, and empirical data confirm, that local circumstances (e.g., economic climate) influence the rapidity of union formation, these models and data do not suggest that the processes and causal mechanisms themselves interact with geographical area (Lichter, LeClere, and McLaughlin 1991; Lichter et al. 1992). While it is possible that the processes and causal parameters underlying marriage and cohabitation among the families participating in our study may be different than those observed in a nationally representative sample, it is doubtful that our conclusions would be qualitatively different. At least, there is no prior theoretical or empirical basis for expecting this to be so. Furthermore, past research provides convincing evidence that, when comparable data are available, results based on the IPS are very similar to those based on national studies (Thornton, Freedman, and Axinn 2002).

Measures of Marriage and Cohabitation

Cohabitation and marriage transitions were measured using a life-history calendar (Freedman et al. 1988). This procedure provides the precise timing (month and year) of entries into and exits from cohabitation and marriage between the ages of 15 and 31. Cohabitation is defined as living with a person of the opposite sex in an intimate relationship without being married.

For this study, we focus on entries into first marriage and first cohabitation. For entry into marriage, we consider two transition rates, the "partial" rate of marriage with cohabitation as a competing risk and the "total" rate of marriage ignoring cohabitation. For entry into cohabitation, we treat marriage as a competing, absorbing state. We also combine the two states of marriage and cohabitation to consider the total union transition rate, which is defined as the rate of entry into either marriage or cohabitation. A respondent is at risk of entering marriage or cohabitation in a given month until an event occurs or until the respondent reaches the end of the study at age 31. As is typical for event history data, we arrange our data into person-month records, with 47,194 observations in the male subsample and 41,332 observations in the female subsample. Table 1 reports the mean monthly probabilities for the four types of transitions as well as the numbers of these events. Note that the dependent measures are actually "probabilities" rather than "rates." The two terms are practically interchangeable given the very small scale of time (i.e., months) as units for the discrete-time event history analysis. We observe, for example, the average monthly probability of marriage with cohabitation as a competing risk is 0.0031 for men and 0.0050 for women. The higher probability for women than for men is attributable to the social norm of age hypergamy; that is, women typically marry men older than themselves. We also note that in our sample, 146 men and 206 women married without cohabiting, while, in total, 296 men and 352 women married.

Table 1 About Here

Measures of Earnings Potential

While marriage and cohabitation events are precisely measured in the IPS data, we do not have measures of current earnings during intervals between interviews. However, the measurement of current earnings is of less theoretical importance than is the measurement of long-term economic potential. In this research, we therefore make a serious effort to develop measures of earnings potential. By "potential" we mean a latent, unobservable capacity. Since earnings potential is inherently unobservable, not only to us as researchers but also to individuals themselves and to their potential partners, it can affect entry to marriage and cohabitation only through subjective understanding, i.e., perception. In forming such perceptions, however, individuals may be myopic and rely mainly on current and past situations. Given the uncertainty as to how individuals perceive the earnings potential of possible union partners, we develop five different measures to capture earnings potential in five different segments of the life course: predicted current earnings, predicted earnings over the next five years, predicted past earnings, predicted future earnings, and predicted lifetime earnings. We will compare the explanatory power of these five measures, all of which are estimated from respondents' past and current observed characteristics through a two-step statistical procedure. These measures are time-varying and ascertained at the person-month level.

Step 1:

We first used the 1990 Census data to estimate sex- and education-specific earnings equations as functions of potential work experience. Following Mincer (1974), we approximate work experience as the difference between current age and the normative age at which the respondent's highest level of education is attained. Letting j denote education (1=less than high school, 2=high school, 3=some college, and 4=college+), and k (1,2, ….K) denote experience, we have the following approximations:

For j=1, k=age-16, k=1… 44

For j=2, k=age-18, k=1… 42 (1)

For j=3, k=age-20, k=1… 40

For j=4, k=age-22, k=1… 38

We allow final year of experience to vary by educational attainment so that all workers are conveniently assumed to retire at age 60. Letting i denote sex (1=male and 2=female), we estimate earnings as a nonparametric function of education, sex, and work experience for the entire i×j×k cross-classification using the 5-percent 1990 PUMS.[1] The dependent variable in these equations is the natural logarithm of total yearly earnings in 1989. We restrict the sample to full-time workers who worked year round and had positive earnings.[2] Regression analysis in this case is tantamount to computing the mean of logged earnings for each i×j×k cell. We then

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[1] This means that we allow full interactions among education, sex, and experience, all of which are represented by dummy variables subject to usual normalization constraints. An earlier version of this paper compared this nonparametric approach to Mincer's (1974) quadratic function approach and found the nonparametric approach preferable. Note that the 5-percent PUMS is very large, with more than five million cases in our analysis.

[2] Full-time work is operationalized as having worked at least 35 hours per week, and year-round work is operationalized as having worked for at least 50 weeks in 1989. We also excluded respondents who turned out to have negative years of experience according to equation (1).

take the exponential function of the mean and denote this variable by $Y_{ijk.}$. We call Y_{ijk} the "unmodified" predicted *current earnings* potential. The meaning of "unmodified" will be apparent in Step 2. For now, we are dealing with unmodified earnings only.

Predicted *earnings over the next five years* is calculated as the sum of the predicted earnings at the current level of educational attainment and work experience and the predicted earnings over the following four years. That is,

$$Y\_5_{ijk} = \sum_{x=k}^{k+4} Y_{ijx} \ . \tag{2}$$

Calculation of predicted *future earnings* is based on a convenient assumption that permanent labor force exit (i.e., retirement) occurs at age 60 for men and women of all levels of educational attainment. This variable is thus calculated as:

$$Y\_AF_{ijk} = \sum_{x=k}^{60-\theta_j} Y_{ijx} \ , \tag{3}$$

where θ_j refers to the normative ages of school completion (i.e., θ_j = 16, 18, 20, 22 respectively for j=1, 2, 3, 4). Similarly, we construct an analogous measure for total *past earnings*. This variable is calculated as the sum of cumulative earnings at all levels of educational attainment:

$$Y\_BF_{ij\bar{k}} = \sum_{j=1}^{4} \sum_{x=0}^{k_j} Y_{ijx} \ , \tag{4}$$

where k_j is the *actual* years of work experience at educational level j constructed from the life-history calendar, and subscript $\bar{k}$ refers to respondents' observed work history. Calculation of this variable proved challenging in that it required the construction of four additional variables representing cumulative past work experience at each of the four levels of educational attainment. Finally, summing (3) and (4) yields the predicted *lifetime earnings*:

$$Y\_T_{ij\bar{k}} = (Y\_BF_{ij\bar{k}} + Y\_AF_{ijk})\,. \qquad (5)$$

These five variables were then appended to the person-period data in IPS by matching on values of sex, educational attainment, and educational attainment-specific labor force experience.

<u>Step 2:</u>

The measures discussed in step 1 are crude because they do not take into account other observed attributes in the data that predict earnings. To more precisely predict earnings, we then modified the sex-education-experience-specific values of the measures calculated in equations (2)-(5) according to individual variation in other observable characteristics: cognitive ability assessed when the respondents were age 18, school quality (for college attendants and graduates), and college major (for college attendants and graduates). This modification is accomplished by employing "shift" parameters derived from the estimation of sex- and education-specific wage functions based on data from the sophomore cohort of the High School and Beyond (HS&B) study. To accomplish this, we first estimated the 1992 logged earnings of the HS&B respondents as a function of cognitive ability, college quality, and college major. We approximated cognitive ability using the total scores from math and reading tests in HS&B. After collapsing colleges attended by the HS&B respondents into a 17-category classification scheme, we measure college quality as the mean SAT score for entering students in these different school types. Similarly, we grouped college majors into 14 categories to capture between-group variation while maintaining reasonable sample sizes within groups.[3]

We then borrowed the exponentiated coefficients from these regression equations based on the HS&B as shift parameters for the earnings potential measures in the IPS data. To do this, we code college majors in the IPS data using the same classification system as in the HS&B data and append institution-specific mean SAT scores for those respondents who attended college. The IPS survey did not test respondents in any subject matter but gave a 13-item general aptitude test asking respondents to identify the similarity between pairs of words. While the test scores from the HS&B data and those from the IPS data are therefore not strictly comparable, we make the assumption that they are highly correlated. We converted both scores to a standardized scale (with a mean of zero and variance of one) so that the coefficient of test scores from the HS&B data can be used as a shift parameter for the ability measure available in the IPS data. Our approach necessitates the assumption that the effects of cognitive ability and school characteristics are multiplicative and do not vary by age. For example, we assume that the positive effect of cognitive ability estimated using the HS&B data shifts wages upward by a proportional amount at all levels of work experience. This assumption is tantamount to a non-interactive model with logged earnings as the dependent variable, a common practice in research

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[3] The seventeen college categories are combinations of visibility (national versus regional), type (public versus private), rank (tier 1 through tier 4), and curriculum (university, liberal arts college, specialty school). The categories for college majors are physical science, math, biological science, engineering, pre-professional, computer science, business, social science, humanities, art and music, education, communications, agriculture, and other.

on earnings (e.g., Mincer and Polachek 1974). These modified earnings measures are then incorporated as covariates in models for the timing of first union formation.

One advantage of our approach is that we are able to estimate earnings potentials for all individuals in the sample, regardless of their work status and experience. At the bottom of Table 2, we present descriptive statistics for the five measures of earnings potential by gender. Not surprisingly, we observe that men's future earnings potential is much greater than women's, although gender differences in current and past earnings are small. We also present the descriptive statistics for other explanatory variables used in our multivariate analysis.

Table 2 About Here

In choosing other explanatory variables, we closely follow the earlier research of Thornton, Axinn, and Teachman (1995). Of particular interest are two separate time-varying variables measuring cumulative years of attained education and enrollment status. We also control for other factors that are known to affect union formation: religion, parents' total years of education, mother's age at first marriage, mother's premarital pregnancy, and mother's previous marital experience. Finally, we include dummy variables representing different parities, since the original sample was stratified by parity.

Statistical Models

Our statistical models are similar to those used in Thornton, Axinn, and Teachman (1995), but we extend their work in three significant ways. First, we study the likelihood of entry to marriage and cohabitation until age 31, whereas Thornton, Axinn, and Teachman had data only up to age 23. Second, we introduce a spline function for modeling the age pattern of entry into marriage and cohabitation. This modification is necessitated by the seven and a half extra years of data, which render the linearity of the age effects implausible (as will be shown later).

Finally, and most importantly, we add the aforementioned five measures of economic potential to the baseline model of Thornton, Axinn, and Teachman (1995).

We estimate a series of logistic regressions for four types of union formation from the event history data. The first dependent outcome is marriage as the destination state and cohabitation as a competing (i.e., censoring) state. Likewise, the second dependent outcome is cohabitation as the destination state and marriage as a competing state. The third dependent outcome is the union of the first two. The fourth dependent outcome treats marriage as the sole destination state and ignores cohabitation.

In a logistic model, exponentiated coefficients represent the multiplicative effects of independent variables on odds (i.e., p/(1-p)). Exponentiated coefficients are commonly labeled as "odds-ratios," as they represent ratios in odds for dummy variables or for a one-unit change in interval variables. However, as shown in Powers and Xie (2000 p.51), odds-ratios are virtually equivalent to relative risks in terms of rates when probabilities are very small, as in our case (Table 1). That is, exponentiated coefficients from our logistic models can be interpreted as multiplicative effects on the hazard rates of union formation.

RESULTS

In Table 3, we present the exponentiated coefficients for the five key earnings measures (after standardization) in five alternative model specifications (A through E), for each combination of gender and type of union formation. Because meaningful comparisons are made difficult by the fact that the different earnings measures vary greatly in scale (see Table 2), we standardized these coefficients so that they all indicate the multiplicative effects on the odds of union formation for a one standard deviation increase in the earnings measures.

Table 3 About Here

Consistent with our theoretical expectations and with empirical results in the existing literature, we find that potential earnings have a significant positive effect on entry into marriage for men. This is true whether cohabitation is treated as a competing risk (column 1) or ignored (column 4), although the effects are attenuated somewhat when cohabitation is ignored. Note that these effects are above and beyond the accumulation of schooling considered by Thornton, Axinn, and Teachman (1995). Looking at column 1, we see that, among the different measures of earnings potential, the effect of past earnings is the largest, increasing the odds of marriage by 48 percent per standard deviation, followed by similar effects for the other four measures (16-21 percent increase). From these results, it is tempting to conclude that past earnings are more important, since they are likely to be known to both the male respondents and their marriage partners and thus enable them to "afford" to marry early. However, we caution the reader that past earnings are estimated with more accuracy, given our use of actual labor force participation histories in constructing this measure (see equation 4). It is possible that larger measurement errors for other earnings measures attenuate their estimated effects. Further, it is worth noting that current earnings potential is the second best predictor, indicating that future earnings potential is either estimated with more noise or in fact does not matter more than current earnings potential. Regardless of the relative importance of alternative measures, we are confident in drawing the conclusion that economic capacity clearly accelerates the process of marriage for men.

In contrast, these same earnings measures have no statistically significant effects on women's likelihood of marriage. This is true whether cohabitation is treated as a competing risk (column 5) or ignored (column 8). These results demonstrate the asymmetric role of economic potential in marriage formation between men and women. However, it is also noteworthy that there is no evidence in our data that economic potential has any negative effect on women's marriage. In fact, all the coefficients are estimated to be positive (i.e., exponentiated coefficients are greater than one) but not statistically different from zero. Thus, we do not find support for Becker's theory that greater economic capacity makes marriage less attractive to women by reducing their economic gain from marriage.

The results pertaining to entry into cohabitation (columns 2 and 6) are simple and straightforward: none of the measures of earnings potential has any discernable effect, for either men or women. Recall that earnings potential has large and positive effects on the likelihood of marriage for men but not for women. The results for the transition to cohabitation suggest that, for men, the causal mechanisms leading to marriage are different from those leading to cohabitation--economic resources hasten marriage but not cohabitation. For women, earnings potential appears to be irrelevant for both types of union formation.

The results for total union formation (columns 3 and 7) are predictable: they lie between the results for marriage and the results for cohabitation. For women, the effects of earnings potential remain insignificant. For men, collapsing marriage and cohabitation into a single destination state dilutes the significant effects of earnings potential on marriage by more than 50 percent, with only one measure (past earnings) remaining statistically significant at the 0.05 level of confidence.

In Table 4, we present the estimated coefficients of Model B for the likelihood of entering marriage with cohabitation as a competing risk. As in Table 3, the coefficients are presented as odds-ratios. The coefficients of the age spline show the expected inverted-U shape: increasing rapidly between ages 15 and 19, slowing down in the early twenties, plateauing in the mid-twenties, and declining thereafter. The coefficients of most other covariates are in the expected direction. For example, consistent with Thornton, Axinn, and Teachman (1995), we find a significantly negative effect of school enrollment on marriage. In contrast to Thornton, Axinn, and Teachman (1995), however, we no longer observe a significantly positive effect of educational attainment on women's likelihood of marriage.[4]

Table 4 About Here

Similarly, in Table 5, we present the estimated coefficients of Model B with cohabitation as the dependent variable and marriage as a competing risk. The age pattern of cohabitation as represented by the spline function is similar to that for marriage. With the notable exception of educational attainment, other estimated coefficients are also in the expected direction. As in Thornton, Axinn, and Teachman (1995), we also find a significantly negative effect of educational attainment on women's entry into cohabitation, with each additional year of education reducing the likelihood by about 21 percent. However, we do not find such an effect for men.

Table 5 About Here

DISCUSSION AND CONCLUSION

In this study, we developed an innovative method for measuring earnings potential and used this information as a predictor of the likelihood of entering marriage or cohabitation. Our approach

[4] This finding is not due to the inclusion of earnings potential in the model. When we exclude the earnings measure, the educational attainment coefficient remains essentially the same. For men, our estimated effect (a 16 percent increase in the likelihood of marriage per additional year of education) is also much smaller than that reported by Thornton, Axinn, and Teachman (1995), who reported a 45 percent increase for an additional year of education.

is facilitated by a rich, longitudinal data set which includes fairly accurate education and work histories, and scores on an aptitude test at age 18. For those respondents attending college, we were also able to utilize information about college quality and college major. We calculated five measures of earnings potential: current earnings, earnings over the next five years, total future earnings, past earnings, and lifetime earnings. We show that all five measures of earnings potential strongly and positively influence the likelihood of marriage for men, but not for women. Another important finding is that the measures of earnings potential do not affect entry into cohabiting unions for either men or women.

The rationale for devising these measures of earnings potential is that observed earnings are a poor indicator of young people's economic potential. Indeed, our results seem to contradict the prevailing view in the literature that women's current earnings/income positively affect their likelihood of marriage (e.g., Clarkberg 1999; MacDonald and Rindfuss 1981; Oppenheimer, Kalmijn, and Lim 1997; Sweeney 2002).[5] One possible explanation for this is that the sample size of the IPS data is too small, and/or measurement error for predicted earnings too great, for us to detect the relatively smaller effects of economic potential on marriage for women. It is also possible that women may be heterogeneous, with the effects of economic potential being positive for some but negative for others, resulting in overall insignificant effects. In addition, we speculate that the observed relationship in the past literature between women's earnings and their likelihood of marriage may be confounded by a selectivity bias: those women who strive to maximize current earnings while foregoing future earnings growth may be more likely to enter

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[5] Our results are consistent with Smock and Manning's (1997) finding that men's, but not women's, economic resources, speed up transition from cohabitation to marriage.

marriage early. As shown in the human capital literature (e.g., Polachek 1979; 1981), optimal pre-marital jobs for women who plan to specialize in household production once married are those that offer relatively high starting wages and allow for easy reentry following temporary disruption, but as a consequence offer little prospect for future earnings growth. Alternatively, women who intend to specialize in market production should, like men, choose jobs in career tracks that may have lower starting wages but also offer long-term potential for earnings growth. Career-oriented women may therefore be observed to have low current earnings at young ages even though their economic potential (both current and future) is high. These women are also likely to postpone marriage. If true, the combination of these two scenarios would produce biased results in which current earnings are observed to accelerate women's marriage. There is some support in the literature for this conjecture. Mare and Winship (1991), for example, find that employment potential (rather than actual employment) has a negative effect on marriage for white women. Also using earnings potential estimated by a different method, Sweeney (1999) reports negative effects of earnings potential on marriage for an earlier cohort of women and very small positive effects for a recent cohort of women.

The current literature on cohabitation is much smaller and less conclusive. Clarkberg (1999) reports positive effects of economic variables on entry into both cohabitation and marriage, concluding that "cohabitation is like marriage in that it selects higher-income individuals out of singlehood" (p.962). However, Clarkberg's conclusion seems to contradict Thornton, Axinn, and Teachman's (1995) finding that accumulated schooling negatively impacts entry into cohabitation. In our analysis, we find a negative effect of educational attainment on cohabitation for women.

While our analysis cannot distinguish between Becker's gender role specialization model and Oppenheimer's search-theoretic model, our results are not consistent with either. Still, we find a gender asymmetry as predicted by the role specialization model in that the effects of earnings potential on marriage are close to zero for women but strongly positive for men. We do not find a similar asymmetry for cohabitation. One theoretical implication of these results is that marriage seems more "gendered" than cohabitation.

One potential criticism of our study is that our estimated measures of earnings potential are contaminated with too much error to be predictive of behavior. It is possible, for example, that the parameter estimates derived from the national data sources (PUMS and High School and Beyond) may not be directly applicable to our regional sample of a particular cohort born in Detroit. For this criticism to hold, we would need to assume that Detroit significantly differs from the nation, or this cohort significantly differs from other cohorts, not just in levels of earnings but also in the returns to the determinants of earnings. It is unfortunate that we do not have time-varying measures of current earnings to cross-validate our predicted current earnings. In additional analysis (not reported here), we experimented with a variable that measures respondents' current work status. We did not find the work status variable to contribute additional explanatory power to our statistical models and thus decided not to include it in our final analysis.

Although we know that our estimated earnings potentials are contaminated by measurement error, it is important to note that we have found large and significant effects of earnings potential on entry into marriage among men. That is, our estimated earnings potential is shown to have face validity in yielding a theoretically expected finding. While we recognize that some of the non-findings in this paper may be attributable to measurement errors or the small sample size, it is safe to reach the following conclusion: at a minimum, our analysis has demonstrated that women's likelihood of marriage is not increased by economic potential to the same extent as men's, and that entry into cohabitation is not increased by economic potential to the same extent as entry into marriage. We leave the further exploration and validation of the findings and ideas that have emerged in this study to future research.

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Table 1: Mean Monthly Probabilities of Entry to Marriage and Cohabitation by Gender.

	Men		**Women**	
Marriage w/ Cohabitation as Competing Risk	0.0031	(146)	0.0050	(206)
Cohabitation w/ Marriage as Competing Risk	0.0044	(207)	0.0045	(185)
Total Union Formation	0.0075	(353)	0.0095	(391)
Marriage Ignoring Cohabitation	0.0051	(296)	0.0071	(352)
Sample Size at Initial Exposure to Risk	428		436	

Note: Data are from IPS. The main entries are the mean monthly probability of an event occurring within a month. The number of uncensored person-months at risk of either cohabitation or marriage is 47,194 in the male subsample and 41,332 in the female subsample. The total numbers of events experienced by members in the sample are in parentheses.

Table 2: Descriptive Statistics of Explanatory Variables by Gender.

	Men		Women	
	Mean	SD	Mean	SD
Age				
15-19	0.36		0.41	
20-23	0.34		0.34	
24-28	0.22		0.19	
29-32	0.08		0.07	
Years of Education after Age 16	3.19	1.86	3.13	1.78
Enrollment Status				
Not enrolled	0.40		0.33	
Enrolled full-time	0.53		0.58	
Enrolled part-time	0.07		0.09	
Religion				
Fundamentalist Protestant	0.06		0.11	
Non-fundamentalist Protestant	0.22		0.25	
Catholic	0.58		0.55	
Jewish	0.04		0.03	
Other	0.01		0.02	
None	0.09		0.04	
Parents' Total Years of Education	25.16	3.92	24.87	3.77
Mother's Age at First Marriage	20.82	2.94	20.78	3.16
Mother's Premarital Pregnancy				
No	0.86		0.82	
Yes	0.14		0.18	
Mother's Previous Marital Experience				
Mother stably married 1962-80	0.78		0.81	
Mother widowed 1962-80	0.04		0.04	
Mother divorced and remarried 1962-80	0.07		0.06	
Mother divorced and not remarried 1962-80	0.11		0.08	
Mother's Parity				
First child	0.38		0.34	
Second child	0.27		0.36	
Fourth child	0.34		0.30	
Estimated Earnings (in 1989 dollars)				
Current Earnings	20,533	11,907	19,863	16,506
Earnings over Next 5 Years	124,134	67,332	114,432	93,927
Future Earnings	1,700,126	666,855	1,198,893	768,134
Past Earnings	84,055	90,963	73,811	100,157
Lifetime Earnings	1,784,180	710,883	1,272,704	832,172

Note: See Table 1 for an explanation of the data.

Table 3: Estimated Effects of Different Measures of Earnings Potential on Four Hazard Rates of Union Formation, by Gender

	Men				Women			
Earnings Variables	Marriage w/ Cohab. as Competing Risk (1)	Cohab. w/ Marriage as Competing Risk (2)	Total Union Formation (3)	Marriage Ignoring Cohab. (4)	Marriage w/ Cohab. as Competing Risk (5)	Cohab. w/ Marriage as Competing Risk (6)	Total Union Formation (7)	Marriage Ignoring Cohab. (8)
A Current Earnings	1.21 *	0.98	1.10	1.20 **	1.13	1.05	1.09	1.07
B Earnings over Next 5 Years	1.21 *	0.98	1.10	1.20 **	1.11	1.06	1.09	1.06
C Future Earnings	1.16 *	0.97	1.07	1.15 **	1.08	1.08	1.08	1.05
D Past Earnings	1.48 **	1.04	1.25 *	1.35 **	1.20	0.99	1.09	1.08
E Lifetime Earnings	1.17 *	0.98	1.08	1.17 **	1.09	1.08	1.08	1.06

*p<.05, **p<.01

Note: Entries are odds ratios associated with one standard deviation in relevant earnings measure, estimated from discrete-time event history models with logit specification. Five measures of earnings potential are alternately included in Models A through E, which all control for the following variables: age (spline), school enrollment, educational attainment, religious affiliation, parents' educational attainment, mother's premarital pregnancy status, mother's age at marriage, mother's marital history, and mother's parity.

Table 4: Estimated Logit Coefficients of Model B Predicting the Likelihood of Marriage with Cohabitation as a Competing Risk.

	Men		Women	
	Coefficient	z-ratio	Coefficient	z-ratio
Constant (×1,000)	0.002	-3.073	1.623	-5.883
Age (spline function)				
15-19	1.171	1.776	1.075	3.715
20-23	1.018	2.166	1.007	1.156
24-28	1.000	-0.054	0.993	-1.432
29-32	0.973	-2.154	0.972	-1.816
Years of Education after Age 16	1.164	2.165	1.068	0.886
Enrollment Status (excluded = not enrolled)				
Enrolled full-time	0.309	-3.979	0.402	-4.224
Enrolled part-time	1.111	0.437	0.634	-1.901
Religion (excluded = Fundamentalist Protestant)				
Non-fundamentalist Protestant	0.611	-1.311	1.117	0.457
Catholic	0.821	-0.590	0.798	-0.997
Jewish	0.355	-1.503	0.798	-0.410
Other	0.485	-0.911	0.535	-1.008
None	0.117	-2.751	0.662	-0.884
Parents' Total Years of Education	0.956	-1.791	0.948	-2.709
Mother's Age at First Marriage	1.001	0.038	0.970	-1.290
Mother's Premarital Pregnancy (excluded=no)	1.466	1.587	1.111	0.575
Yes				
Mother's Previous Marital Experience (excluded = mother stably married 1962-80)				
Mother widowed 1962-80	1.679	1.189	0.536	-1.323
Mother divorced and remarried 1962-80	0.616	-1.114	1.167	0.525
Mother divorced and not remarried 1962-80	0.876	-0.398	0.691	-1.162
Mother's Parity (excluded=first child)				
Second child	1.381	1.542	0.765	-1.525
Fourth child	0.959	-0.189	1.050	0.272
Estimated Earnings (in 1989 dollars)				
Earnings over Next 5 Years (×100,000)	1.323	2.430	1.120	1.371
Model Chi-Square (DF = 21)	215.45		184.58	
No. of Person-Months	47,194		41,332	

Note: Coefficients are in odds-ratios scale. Z-ratios are asymptotic test statistics for the hypothesis that the odds-ratios are one.

Table 5: Estimated Logit Coefficients of Model B Predicting the Likelihood of Cohabitation with Marriage as a Competing Risk.

	Men		Women	
	Coefficient	z-ratio	Coefficient	z-ratio
Constant (×1,000)	5.692	-5.393	0.125	-9.284
Age (spline function)				
15-19	1.049	3.157	1.070	4.604
20-23	1.007	1.166	1.011	1.661
24-28	0.999	-0.254	0.996	-0.655
29-32	0.985	-1.214	1.013	1.057
Years of Education after Age 16	0.991	-0.143	0.808	-2.479
Enrollment Status (excluded = not enrolled)				
Enrolled full-time	0.369	-4.567	0.634	-2.001
Enrolled part-time	0.521	-2.315	0.646	-1.640
Religion (excluded = Fundamentalist Protestant)				
Non-fundamentalist Protestant	0.495	-2.372	1.103	0.340
Catholic	0.551	-2.311	1.094	0.341
Jewish	0.552	-1.180	1.832	1.307
Other	0.635	-0.596	1.060	0.092
None	1.049	0.157	1.662	1.314
Parents' Total Years of Education	0.979	-0.924	1.052	2.151
Mother's Age at First Marriage	0.944	-1.897	0.981	-0.692
Mother's Premarital Pregnancy (excluded=no)				
Yes	1.243	1.157	1.727	2.975
Mother's Previous Marital Experience (excluded = mother stably married 1962-80)				
Mother widowed 1962-80	1.854	1.747	1.812	1.722
Mother divorced and remarried 1962-80	2.011	3.142	2.297	3.424
Mother divorced and not remarried 1962-80	1.207	0.842	1.627	2.076
Mother's Parity (excluded=first child)				
Second child	1.071	0.375	1.106	0.554
Fourth child	0.961	-0.220	0.983	-0.087
Estimated Earnings (in 1989 dollars)				
Earnings over Next 5 Years (×100,000)	0.968	-0.234	1.069	0.663
Model Chi-Square (DF = 21)	160.69		118.92	
No. of Person-Months	47,194		41,332	

Note: Coefficients are in odds-ratios scale. Z-ratios are asymptotic test statistics for the hypothesis that the odds-ratios are one.