ECONOMIC POTENTIAL AND ENTRY INTO MARRIAGE AND COHABITATION*

YU XIE, JAMES M. RAYMO, KIMBERLY GOYETTE, AND ARLAND THORNTON

This article explores the relationship between economic potential and rates of entry into marriage and cohabitation. Using data from the 1990 census and the 1980–1992 High School and Beyond (Sophomore Cohort), we developed a method for explicitly estimating five time-varying measures of earnings potential. The analyses of union formation are based on an intergenerational panel study of parents and children, to which our measures of earnings potential were appended. The results indicate that all five measures of earnings potential strongly and positively influence the likelihood of marriage for men, but not for women. Earnings potential does not affect entry into cohabiting unions for either men or women.

The formation of unions by 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 appears that the proportion who will never marry is increasing (e.g., Bumpass and Sweet 1989; Bumpass, Sweet, and Cherlin 1991; 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 of changes in marital 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 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 marital 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 focused exclusively on marriage, ignoring cohabitation. If "cohabitation is very much a family status" (Bumpass et al. 1991:926), research on family formation should also study entry into 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.

*Yu Xie, Population Studies Center, Institute for Social Research, 426 Thompson Street, University of Michigan, Ann Arbor, MI 48106; E-mail: yuxie@umich.edu. James M. Raymo, Department of Sociology, University of Wisconsin–Madison; Kimberly Goyette, Department of Sociology, Temple University; Arland Thornton, Survey Research Center, Population Studies Center, and Department of Sociology, University of Michigan. An early version of the article was presented at the 2001 annual meeting of the Population Association of America, Washington, DC, March 29–31. This research was supported by a research grant from the National Institute of Child Health and Development. The authors are grateful to Frances Goldscheider for her helpful comments on an early version of the article.

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 et al. (1995), this article 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 developed 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 the quality of colleges and fields of study for individuals who completed their postsecondary education. Our research on entries into marriage and cohabitation is based on the same intergenerational panel study of parents and children used by Thornton et al. (1995), with updated information on the respondents' life experiences through age 31. The auxiliary estimation of predicted earnings used data from the 1990 Census 5% Public Use Microsample (PUMS) and the 1980–1992 High School and Beyond (HS&B), 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, thereby 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 relevant to this hypothesis has been mixed. Some studies, based on aggregate-level data and cross-sectional survey data, have found a negative relationship between indicators of women's economic status (i.e., educational attainment, employment, earnings) and the prevalence or incidence of marriage. Research that has used more appropriate longitudinal, individual-level data, however, 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) presented a thorough review of this literature. Results from investigations of men's marital 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; Sassler and Schoen 1999; Sweeney 2002; Teachman, Polonko, and Leigh 1987).

With attitudinal data providing little support for a rejection of marriage by economically independent women, Oppenheimer 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 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 marital partners.

Cohabitation

One of the most notable trends in family behavior in the United States is the rapid increase in nonmarital cohabitation. Although cohabiting unions resemble marriage in many respects and often serve as precursors to marriage (Bumpass et al. 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 views of cohabitation in the literature are: (1) cohabitation as an alternative to marriage, (2) cohabitation as an alternative to being single, and (3) 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 childbearing) 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 (1990) showed 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 being 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 premarital cohabitation may be similar to those on marriage. On the other hand, for some cohabiting couples who are planning to marry, one reason for cohabiting before marriage may well be the lack of sufficient economic resources for marriage (e.g., Oppenheimer 1988:71). This theoretical ambiguity about the nature of cohabitation suggests the need to treat marriage and cohabitation as two distinct types of unions.

Economic Potential

As we described earlier, there is a large and well-researched literature on the influence of economic resources on marital 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 et al. 1997; Sweeney 2002), educational attainment (Clarkberg 1999; Goldscheider and Waite 1986; Goldstein and Kenney 2001; Mare and Winship 1991; Oppenheimer et al. 1995; Oppenheimer et al. 1997; Sweeney 2002; Thornton, Axinn, and Teachman 1995; Waite and Spitze 1981), work experience (Clarkberg 1999; Oppenheimer et al. 1997; Sweeney 2002), employment (Goldscheider and Waite 1986; Oppenheimer et al. 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). These empirical measures, however, 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: it is only postmarriage economic well-being that should have any *direct* relevance for marital 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 the 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 assess accurately 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 young people may still invest in the accumulation of human capital—by receiving formal education in school or undertaking training—for rapid earnings growth in the future. At the same time, individuals base their decisions to form unions 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 postmarriage economic behaviors of men and women can be substantially altered by marriage itself.

DATA AND METHODS

The Sample

The primary data for this research came 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 were 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, postmarital 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 mothers, representing 92% of the families drawn for the sample. For this study, we restricted the sample to focal children who had not entered a marital or cohabiting relationship before age 15 and who provided valid information on all the explanatory variables, yielding a sample of 428 men and 436 women. This analytic sample suffered 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 our analytical sample closely resembles that used by Thornton et al. (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: 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). The 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

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

Type of Union Formation	Men		Women	
	Mean Monthly Probability	Total Number of Events	Mean Monthly Probability	Total Number of Events
Marriage With Cohabitation as Competing Risk	0.0031	146	0.0050	206
Cohabitation With 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	

Notes: The data are from the IPS. 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.

and cohabitation vary across birth order or metropolitan area. It has sometimes been hypothesized, and empirical data have confirmed, that local circumstances (e.g., economic climate) influence the rapidity of union formation, but these models and data do not suggest that the processes and causal mechanisms themselves interact with geographic area (Lichter, LeClere, and McLaughlin 1991; Lichter et al. 1992). Although it is possible that the processes and causal parameters underlying marriage and cohabitation among the families who participated 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 the case. Furthermore, past research has provided convincing evidence that when comparable data are available, results based on the IPS are similar to those based on national studies (Thornton, Freedman, and Axinn 2002).

Measures of Marriage and Cohabitation

Cohabitation and marital 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 focused on entries into first marriage and first cohabitation. For entry into marriage, we considered 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 treated marriage as a competing, absorbing state. We also combined 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 was at risk of entering marriage or cohabitation in a given month until an event occurred or until the respondent reached the end of the study at age 31. As is typical for event-history data, we arranged 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 number of these events. Note that the dependent measures are actually "probabilities," rather than "rates." The two terms are practically interchangeable, given the small scale of time (i.e., months) as units for the discrete-time event-history analysis. We observed, for example, that 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.

Measures of Earnings Potential

Although marital and cohabitation events are precisely measured in the IPS data, we do not have measures of current earnings during intervals between the 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 made a serious effort to develop measures of earnings potential. By *potential* we mean a latent, unobservable capacity. Because earnings potential is inherently unobservable, not only to us as researchers but also to individuals and their potential partners, it can affect entry into marriage and cohabitation only through subjective understanding. In forming such perceptions, however, individuals may be myopic and rely mainly on current and past situations. Given the uncertainty of how individuals perceive the earnings potential of possible union partners, we developed five measures to capture earnings potential in five 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 compared the explanatory power of these five measures, all of which were estimated from the 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 approximated work experience as the difference between current age and the normative age at which the respondent's highest level of education was attained. Letting j denote education (1 = less than high school, 2 = high school, 3 = some college, and 4 = college+), and k $(1, 2, \ldots, K)$ denote potential work experience, we have the following approximations:

$$\begin{aligned}
&\text{For } j = 1,\ k = \text{age} - 16,\ k = 1 \ldots 44.\\
&\text{For } j = 2,\ k = \text{age} - 18,\ k = 1 \ldots 42.\\
&\text{For } j = 3,\ k = \text{age} - 20,\ k = 1 \ldots 40.\\
&\text{For } j = 4,\ k = \text{age} - 22,\ k = 1 \ldots 38.
\end{aligned} \tag{1}$$

We allowed final year of experience to vary by educational attainment so that all workers were conveniently assumed to retire at age 60. Letting i denote sex (1 = male and 2 = female), we estimated earnings as a nonparametric function of education, sex, and work experience for the entire $i \times j \times k$ cross classification using the 5% 1990 PUMS.[1] The dependent variable in these equations is the natural logarithm of total yearly earnings in 1989. We restricted 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 $i \times j \times k$ cell. We then took the exponential function of the mean and denoted this variable by $Y_{ijk.}$. We call Y_{ijk} the "unmodified" predicted *current earnings*.

1. This nonparametric approach means that we allowed full interactions among education, sex, and experience, all of which are represented by dummy variables subject to usual normalization constraints. An early version of this article compared this nonparametric approach with Mincer's (1974) quadratic-function approach and found the nonparametric approach preferable. Note that the 5% PUMS is large, with more than 5 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 Eq. (1).

The meaning of "unmodified" is 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,

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

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

(3)

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

(4)

where k_j is the *actual* years of work experience at educational level j, constructed from the life-history calendar, and subscript 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 Eqs. (3) and (4) yields the predicted *lifetime earnings*:

(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.

Step 2. 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 predict earnings more precisely, we then modified the sex-education-experience-specific values of the measures calculated in Eqs. (2) through (5) according to individual variation in other observable characteristics: cognitive ability assessed when the respondents were aged 18, school quality (for college graduates), and college major (for college graduates). We accomplished this modification by using "shift" parameters derived from the estimation of sex- and education-specific wage functions based on data from the sophomore cohort of the HS&B. 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 the HS&B. After collapsing colleges attended by the HS&B respondents into a 17-category classification scheme, we measured college quality as the mean SAT score for entering students in these different types of schools. Similarly, we grouped college majors into 14 categories to capture between-group variation while maintaining reasonable sample sizes within the groups.[3]

3. The 17 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

We then used 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 so, we coded college majors in the IPS data using the same classification system as in the HS&B data and appended institution-specific mean SAT scores for respondents who attended college. The IPS survey did not test the respondents in any subject matter but gave a 13-item general aptitude test that asked the respondents to identify the similarity between pairs of words. Although the test scores from the HS&B data and those from the IPS data are therefore not strictly comparable, we assumed that they are highly correlated. We converted both scores to a standardized scale (with a mean of 0 and variance of 1), so that the coefficient of test scores from the HS&B data could 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 assumed 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 noninteractive model with logged earnings as the dependent variable, a common practice in research on earnings (e.g., Mincer and Polachek 1974). These modified earnings measures were then incorporated as covariates in models for the timing of first union formation.

One advantage of our approach is that we were able to estimate earnings potential 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. It is not surprising 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.

In choosing other explanatory variables, we closely followed the earlier research of Thornton et al. (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 the formation of unions: religion, parents' total years of education, mother's age at first marriage, mother's premarital pregnancy, and mother's previous marital experience. Finally, we included dummy variables representing the mother's parity at the respondent's birth because the original sample was stratified by parity.

Statistical Models

Our statistical models are similar to those used by Thornton et al. (1995), but we extended their work in three significant ways. First, we studied the likelihood of entry to marriage and cohabitation until age 31, whereas they had data only up to age 23. Second, we introduced a spline function for modeling the age pattern of entry into marriage and cohabitation. This modification was 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 important, we added the aforementioned five measures of economic potential to Thornton et al.'s baseline model.

We estimated a series of logistic regressions for the formation of four types of unions 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.

---

categories for college majors are physical science, math, biological science, engineering, preprofessional, computer science, business, social science, humanities, art and music, education, communications, agriculture, and other.

Table 2. Descriptive Statistics of Explanatory Variables, by Gender

	Men		Women	
Variable	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	
Nonfundamentalist 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 1962–1980	0.78		0.81	
Mother widowed 1962–1980	0.04		0.04	
Mother divorced and remarried 1962–1980	0.07		0.06	
Mother divorced and not remarried 1962–1980	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 the next five 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.

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 "odds ratios" because they represent ratios in odds for dummy variables or for a one-unit change in interval variables. However, as Powers and Xie (2000:51) showed, odds ratios are virtually equivalent to relative risks in terms of rates when probabilities are small, as in our case (see 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 formed. 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 1-standard-deviation increase in the earnings measures.

Consistent with our theoretical expectations and with empirical results in the literature, we found 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. These effects are above and beyond the accumulation of schooling considered by Thornton et al. (1995). As column 1 indicates, among the different measures of earnings potential, the effect of past earnings is the largest, increasing the odds of marriage by 48% per standard deviation, followed by similar effects for the other four measures (16%–21% increase). From these results, it is tempting to conclude that past earnings are more important because they are likely to be known to both the male respondents and their marital 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 histories of actual labor-force participation in constructing this measure (see Eq. (4)). It is possible that larger measurement errors for other earnings measures attenuate their estimated effects. Furthermore, it is worth noting that current earnings potential is the second-best predictor, indicating that future earnings potential either is estimated with more error or 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 1) or ignored (column 4). 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 1) but not statistically different from 0. Thus, we do not find support for Becker's (1973, 1974, 1991) theory that greater economic capacity makes marriage less attractive to women by reducing their economic gains from marriage.

The results pertaining to entry into cohabitation (column 2) are simple and straightforward: none of the measures of earnings potential has any discernible 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

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

Earnings Variables	Marriage With Cohabitation as a Competing Risk (1)	Cohabitation With Marriage as a Competing Risk (2)	Total Union Formation (3)	Marriage Ignoring Cohabitation (4)
Men				
A. Current earnings	1.21*	0.98	1.10	1.20**
B. Earnings over the next five years	1.21*	0.98	1.10	1.20**
C. Future earnings	1.16*	0.97	1.07	1.15**
D. Past earnings	1.48**	1.04	1.25*	1.35**
E. Lifetime earnings	1.17*	0.98	1.08	1.17**
Women				
A. Current earnings	1.13	1.05	1.09	1.07
B. Earnings over the next five years	1.11	1.06	1.09	1.06
C. Future earnings	1.08	1.08	1.08	1.05
D. Past earnings	1.20	0.99	1.09	1.08
E. Lifetime earnings	1.09	1.08	1.08	1.06

Notes: Entries are odds ratios associated with a 1-standard-deviation increase in the 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.

$^{*}p < .05$; $^{**}p < .01$

cohabitation—economic resources hasten marriage but not cohabitation. For women, earnings potential appears to be irrelevant for the formation of both types of unions.

The results for total union formation (column 3) 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%, with only one measure (past earnings) remaining statistically significant at the .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 20s, plateauing in the mid-20s, and declining thereafter. The coefficients of most other covariates are in the expected direction. For example, consistent with Thornton et al. (1995), we found a significantly negative effect of school enrollment on marriage. In contrast to Thornton et al., however, we no longer found a significantly positive effect of educational attainment on women's likelihood of marriage.[4]

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

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

	Men		Women	
Variable	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)				
Nonfundamentalist 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)				
Yes	1.466	1.587	1.111	0.575
Mother's Previous Marital Experience (excluded = mother stably married 1962–1980)				
Mother widowed 1962–1980	1.679	1.189	0.536	–1.323
Mother divorced and remarried 1962–1980	0.616	–1.114	1.167	0.525
Mother divorced and not remarried 1962–1980	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 the next five years (× 100,000)	1.323	2.430	1.120	1.371
Model Chi-Square (*df* = 21)	215.45		184.58	
Number of Person-Months	47,194		41,332	

Notes: Coefficients are in odds-ratios scale; *z* ratios are asymptotic test statistics for the hypothesis that the odds ratios are 1.

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

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

	Men		Women	
Variable	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)				
Nonfundamentalist 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 1962–1980)				
Mother widowed 1962–1980	1.854	1.747	1.812	1.722
Mother divorced and remarried 1962–1980	2.011	3.142	2.297	3.424
Mother divorced and not remarried 1962–1980	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 the next five years (× 100,000)	0.968	–0.234	1.069	0.663
Model Chi-Square (*df* = 21)	160.69		118.92	
Number of Person-Months	47,194		41,332	

Notes: Coefficients are in odds-ratios scale; *z* ratios are asymptotic test statistics for the hypothesis that the odds ratios are 1.

expected direction. As in Thornton et al. (1995), we also found 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%. However, we did not find such an effect for men.

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 was facilitated by a rich, longitudinal data set that includes fairly accurate education and work histories and scores on an aptitude test at age 18. For the respondents who graduated from college, we were also able to use 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 showed 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 et al. 1997; Sweeney 2002).[5] One possible explanation for this finding is that the sample size of the IPS data is too small and/or measurement error for predicted earnings is 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: women who strive to maximize current earnings while forgoing future growth in earnings may be more likely to enter marriage early. As has been shown in the human capital literature (e.g., Polachek 1979, 1981), optimal premarital jobs for women who plan to specialize in household production when they are 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 a future growth in earnings. Alternatively, women who intend to specialize in market production should, like men, choose jobs in career tracks that may have lower starting wages but 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, found that employment potential (rather than actual employment) has a negative effect on marriage for white women. Using earnings potential estimated by a different method, Sweeney (1999) reported negative effects of earnings potential on marriage for an earlier cohort of women and small positive effects for a recent cohort of women.

The literature on cohabitation is much smaller and less conclusive. Clarkberg (1999) reported 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 et al.'s (1995) finding that accumulated schooling has a negative impact on entry into cohabitation. In our analysis, we found a negative effect of educational attainment on cohabitation for women.

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

Although 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 found 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 did not find similar gender asymmetry for cohabitation. One theoretical implication of these results is that marriage seems more "gendered" than does 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 HS&B) are not 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 that 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 analyses (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, 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. Although we recognize that some of the nonfindings in this article 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 emerged in this study to future research.

REFERENCES

Becker, G.S. 1973. "A Theory of Marriage: Part I." *Journal of Political Economy* 81:813–46.

———. 1974. "A Theory of Marriage: Part II." *Journal of Political Economy* 82:S11–S26.

———. 1991. *A Treatise on the Family*. Cambridge, MA: Harvard University Press.

Bennett, N.G., D.E. Bloom, and P.H. Craig. 1989. "The Divergence of Black and White Marriage Patterns." *American Journal of Sociology* 95:692–722.

Bianchi, S.M. and D. Spain. 1986. *American Women in Transition*. New York: Russell Sage Foundation.

Bumpass, L. and H.-H. Lu. 2000. "Trends in Cohabitation and Implications for Children's Family Contexts in the United States." *Population Studies* 54:29–41.

Bumpass, L.L. and J.A. Sweet. 1989. "National Estimates of Cohabitation." *Demography* 26:615–25.

Bumpass, L.L., J.A. Sweet, and A. Cherlin. 1991. "The Role of Cohabitation in Declining Rates of Marriage." *Journal of Marriage and the Family* 53:913–27.

Cherlin, A.J. 1992. *Marriage, Divorce, Remarriage*. Cambridge, MA: Harvard University Press.

Clarkberg, M. 1999. "The Price of Partnering: The Role of Economic Well-being in Young Adults' First Union Experiences." *Social Forces* 77:945–68.

Cooney, T.M. and D.P. Hogan. 1991. "Family-Building Patterns of Professional Women: A Comparison of Lawyers, Physicians, and Postsecondary Teachers." *Journal of Marriage and the Family* 51:749–58.

Freedman, D., A. Thornton, D. Camburn, D. Alwin, and L. Young-DeMarco. 1988. "The Life History Calendar: A Technique for Collecting Retrospective Data." *Sociological Methodology* 18:37–68.

Goldscheider, F.K. and L.J. Waite. 1986. "Sex Differences in the Entry Into Marriage." *American Journal of Sociology* 92:91–109.

Goldstein, J.R. and C.T. Kenney. 2001. "Marriage Delayed or Marriage Forgone? New Cohort Forecasts of First Marriage for U.S. Women." *American Sociological Review* 66:506–19.

Lichter, D.T., F.B. LeClere, and D.K. McLaughlin. 1991. "Local Marriage Markets and the Marital Behavior of Black and White Women." *American Journal of Sociology* 96:843–67.

Lichter, D.T., D.K. McLaughlin, G. Kephart, and D.J. Landry. 1992. "Race and the Retreat From Marriage: A Shortage of Marriageable Men?" *American Sociological Review* 57:781–99.

Lloyd, K.M. and S.J. South. 1996. "Contextual Influences on Young Men's Transition to First Marriage." *Social Forces* 74:1097–119.

MacDonald, M.M. and R.R. Rindfuss. 1981. "Earnings, Relative Income, and Family Formation." *Demography* 18:123–36.

Manning, W.D. 1995. "Cohabitation, Marriage, and Entry Into Motherhood." *Journal of Marriage and the Family* 57:191–200.

Mare, R.D. and C. Winship. 1991. "Socioeconomic Change and the Decline of Marriage for Blacks and Whites." Pp. 175–202 in *The Urban Underclass,* edited by C. Jencks and P.E. Peterson. Washington, DC: The Urban Institute.

Mincer, J. 1974. *Schooling, Experience, and Earnings*. New York: Columbia University Press.

Mincer, J. and S. Polachek. 1974. "Family Investments in Human Capital: Earnings of Women." *Journal of Political Economy* 82:S76–S108.

Oppenheimer, V.K. 1988. "A Theory of Marriage Timing." *American Journal of Sociology* 94: 563–91.

———. 1994. "Women's Rising Employment and the Future of the Family in Industrial Societies." *Population and Development Review* 20:293–342.

———. 1997. "Women's Employment and the Gains to Marriage: The Specialization and Trading Model of Marriage." *Annual Review of Sociology* 23:431–53.

Oppenheimer, V.K., H.-P. Blossfeld, and A. Wackerow. 1995. "United States of America." Pp. 150–73 in *The New Role of Women: Family Formation in Modern Societies*. Boulder: Westview Press.

Oppenheimer, V.K., M. Kalmijn, and N. Lim. 1997. "Men's Career Development and Marriage Timing During a Period of Rising Inequality." *Demography* 34:311–30.

Oppenheimer, V.K. and V. Lew. 1995. "American Marriage Formation in the 1980s: How Important Was Women's Economic Independence?" Pp. 105–38 in *Gender and Family Change in Industrialized Countries,* edited by K.O. Mason and A.-M. Jensen. Oxford, England: Clarendon Press.

Polachek, S.W. 1979. "Occupational Segregation Among Women: Theory, Evidence, and a Prognosis." Pp. 137–57 in *Women in the Labor Market,* edited by C.B. Lloyd, E.S. Andrews, and C.L. Gilroy. New York: Columbia University Press.

———. 1981. "Occupational Self-Selection: A Human Capital Approach to Sex Differences in Occupational Structure." *Review of Economics and Statistics* 63:60–69.

Powers, D.A. and Y. Xie. 2000. *Statistical Methods for Categorical Data Analysis*. San Diego: Academic Press.

Preston, S.H. and A.T. Richards. 1975. "The Influence of Women's Work Opportunities on Marriage Rates." *Demography* 12:209–22.

Raley, R.K. 1996. "A Shortage of Marriageable Men? A Note on the Role of Cohabitation in Black-White Differences in Marriage Rates." *American Sociological Review* 61:973–83.

Rindfuss, R.R. and A. Van den Heuvel. 1990. "Cohabitation: A Precursor to Marriage or an Alternative to Being Single?" *Population and Development Review* 16:703–26.

Sassler, S. and F.K. Goldscheider. 1997. "Revisiting Jane Austen's Theory of Marriage Timing: Union Formation Among Men in the Late 20th Century." Paper presented at the annual meeting of the Population Association of America, Washington, DC, March 27–29.

Sassler, S. and R. Schoen. 1999. "The Effect of Attitudes and Economic Activity on Marriage." *Journal of Marriage and the Family* 61:147–59.

Schoen, R. and R.M. Weinick. 1991. "Cohabitation, Marriage, and the Division of Labor." Papers on Population WP 91-10. Baltimore, MD: Hopkins Population Center, Johns Hopkins University.

Smock, P.J. and W.D. Manning. 1997. "Cohabiting Partners' Economic Circumstances and Marriage." *Demography* 34:331–42.

Spain, D. and S.M. Bianchi. 1996. *Balancing Act: Motherhood, Marriage, and Employment Among American Women*. New York: Russell Sage Foundation.

Sweeney, M.M. 1999. "Reconsidering the Importance of Women's Economic Prospects for Marriage: The Roles of Expected Earnings Trajectories and Uncertainty." Paper presented at the annual meeting of the Population Association of America, New York, March 25–27.

———. 2002. "Two Decades of Family Change: The Shifting Economic Foundations of Marriage." *American Sociological Review* 67:132–47.

Sweet, J.A. and L.L. Bumpass. 1987. *American Families and Households*. New York: Russell Sage Foundation.

Teachman, J.D., K.A. Polonko, and G.K. Leigh. 1987. "Marital Timing: Race and Sex Comparisons." *Social Forces* 66:239–68.

Thornton, A. 1988. "Cohabitation and Marriage in the 1980s." *Demography* 25:497–508.

Thornton, A., W.G. Axinn, and J.D. Teachman. 1995. "The Influence of School Enrollment and Accumulation on Cohabitation and Marriage in Early Adulthood." *American Sociological Review* 60:762–74.

Thornton, A., R. Freedman, and W.G. Axinn. 2002. "Intergenerational Panel Study of Parents and Children." Pp. 315–45 in *Looking at Lives: American Longitudinal Studies of the Twentieth Century,* edited by E. Phelps, F.F. Furstenberg, Jr., and A. Colby. New York: Russell Sage Foundation.

Waite, L.J. and G.D. Spitze. 1981. "Young Women's Transition to Marriage." *Demography* 18: 681–94.

ERRATUM

Because of a file-translation error, three equations were omitted from an article appearing in this issue. The error occurred on page 357 in "Economic Potential and Entry Into Marriage and Cohabitation," by Yu Xie, James M. Raymo, Kimberly Goyette, and Arland Thornton. We regret that the error was noticed only after the issue was printed and bound. The omitted equations and the associated text are below.

The calculation of predicted *future earnings* is based on a convenient assumption that permanent exit from the labor force (i.e., full 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., $\theta_j = 16, 18, 20, 22$, respectively, for $j = 1, 2, 3, 4$). Similarly, we constructed an analogous measure for total *past earnings*. This variable was 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 Eqs. (3) and (4) yields the predicted *lifetime earnings*:

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

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.