Essays in Economics of the Family: Incorporating Cohabitation

Arif A. Mamun

A dissertation submitted in partial fulfillment of the requirements for the degree of

Doctor of Philosophy

University of Washington

2005

Program Authorized to Offer Degree

Department of Economics

UMI Number: 3183388

University of Washington
Graduate School

This is to certify that I have examined this copy of a doctoral dissertation by

Arif A. Mamun

and have found that it is complete and satisfactory in all respects,
and that any and all revisions required by the final
examining committee have been made.

Chair of Supervisory Committee:

Shelly Lundberg

Shelly J. Lundberg

Reading Committee:

Fahad A. Khalil

Shelly Lundberg

Shelly J. Lundberg

Robert D Plotnick

Robert D. Plotnick

Claus C. Pörtner

Date: 07/14/05

University of Washington

Abstract

Essays in Economics of the Family: Incorporating Cohabitation

Arif A. Mamun

Chair of the Supervisory Committee
Professor Shelly J. Lundberg
Department of Economics

The first essay of this dissertation provides new evidence on wage premiums for men in relation to marriage and cohabitation. Using data from the National Longitudinal Survey of Youth 1979 (NLSY79), we show that even after accounting for selection there is a cohabitation wage premium, albeit smaller than the marriage premium, for white and black men but not for Hispanic men. We find empirical support for a joint human capital hypothesis which suggests that intra-household spillover effects of partner's education can explain the existence of the wage premiums.

A recent strand of literature in demography argues that young unmarried Americans value marriage so highly that it is perceived as a family status to be chosen after certain economic preconditions are fulfilled – after they have achieved the so-called "white picket fence dream" (a house, surplus income etc.). Motivated by these claims, in the second essay we use data from the NLSY79 to examine whether there is any direct relationship between the individual's housing and financial assets and his/her transition into marriage or cohabitation. For both men and women, analysis using a proportional hazard model indicates a positive association of asset ownership with transition into marriage, but not with transition into cohabitation. However, instrumental variables probit estimations, designed to account for the endogeneity of asset-accumulation, either remove the statistical significance of the association between asset ownership and family union transitions, or identify effects that are in the opposite direction to those derived from the time-to-event analysis, indicating dissuading effects of asset ownership on transition to marriage.

The existing theoretical literature on household decision-making makes no distinction between different institutional processes of household formation, namely, cohabitation and marriage. In the third essay, we develop a simple two-period model of family union that distinguishes cohabitation and marriage. The analytical results of the model suggest that compared to marital unions, cohabiting unions have higher risk of dissolution in the future, and involve less intra-household specialization. The model also indicates that improved labor market conditions for men provide stronger incentives for marriage than for cohabitation; and that cost of divorce affects married women's labor supply choice.

TABLE OF CONTENTS

LIST OF FIGURES

LIST OF TABLES

ACKNOWLEDGEMENTS

I wish to thank Professor Shelly Lundberg for the guidance, support and encouragement she provided in completing my dissertation. I would also like to express my sincere gratitude to my supervisory committee members – Professors Fahad Khalil, Robert Plotnick and Claus Pörtner, for their help and thoughtful suggestions.

I would like to acknowledge the financial support that I received over the past few years from the Center for Studies in Demography and Ecology, the William and Flora Hewlett Foundation, the Center for Research on Families, and the Grover and Creta Ensley Fellowship. Their support provided me with the opportunity to conduct research in full capacity, and also allowed me to gain valuable professional experience.

I consider myself lucky to have found a set of wonderful friends in Seattle – in the entering class of 2000, and outside. Having Kisa, Erika, Alan, Kundan, Ron, Dan, Aaron, Evren, Ian, Arabinda, and Shin has made my time in school much easier than I anticipated. Also, often times I would forget that I was away from home as I had Keya bhabi, Morshed bhai, Nandini apa, Nasreen apa, Shams bhai, Hassan bhai, Basher bhai, Nishi, Kazi, Sabah, Adnan bhai, Topu, Fazal bhai, Sohel bhai, Shakil bhai and many others – with whom I spent so many weekends and evenings. Thank you all.

I would also like to express my utmost appreciation to my family for their support and for enduring with me the time that I was away. Last, but not the least, I wish to thank my wife, Nadia Howlader – for being a friend, a helpful statistical consultant, and for bearing with me the challenging days.

DEDICATION

To my parents –

Sufia Khatun and Mohammad Mutiur Rahman

Chapter 1

Is There a Cohabitation Premium in Men's Earnings?

I. Introduction

It is now empirically well recognized among economists that marriage is associated with higher earnings for men. However, a perfunctory look at the data would reveal that over the last three decades men in the U.S. have not only been postponing marriage, but further, a growing percentage of the population are never getting married. A plausible explanation for such delay in reaping the benefits of marriage might lie in the fact that men can acquire similar benefits in another comparable form of family union, viz., cohabitation. The present study brings together new evidence in this regard as we try to identify whether or not wage benefits accrue to cohabiting men as they do to married men.

Considering the dramatic rise in cohabitation over the last four decades, the significance of understanding the implications of cohabitation cannot be overemphasized. Evidently, men benefit substantially from marital union, as indicated by a marital wage premium even after accounting for selection effects. In order to understand if the individual's incentives to cohabit emanate from similar benefits in cohabitation, it is imperative to ask whether there is any cohabitation premium in men's wages. The answer to this question would also suggest whether the underlying intra-household choice decisions in cohabitation are similar to those in marriage.

The focus of this paper is to understand the nature of the marital and cohabitation wage premiums for men by estimating wage equations using longitudinal data that allow for both differential wage growth and selection effects. Our empirical results indicate that there is a cohabitation wage premium, albeit smaller than the marital premium, for white and black men but not for Hispanic men. For white and black men the cohabitation premium persists even after accounting for selection. We consider a joint human capital hypothesis, which has not received adequate attention in the literature heretofore, as an explanation for the existence of marital as well as cohabitation wage premium. According to this hypothesis, partner's human capital contributes positively to a married or cohabiting man's effective stock of human capital and thereby increases his productivity, which translates into increased wage earnings for him. In other words, the marriage and cohabitation wage premiums in men's earnings are reflections of intra-household spillover effects of partner's human capital. Our findings provide some evidence supporting this hypothesis which offers a causal mechanism that links men's productivity and their family status.

The paper provides comparable analysis of white, black, and Hispanic men. Since the current literature is typically restricted to white men, and no comparable longitudinal analysis is available across these three major ethnic groups[1], this study is expected to bridge this gap in the literature. The results presented here would also provide additional substantiation of the generally recognized ethnic diversity in the structure of U.S. families (e.g., Brien 1997; McLanahan and Casper 1995).

In the following section we present a conceptual discussion on the underlying mechanism that can explain men's family status and wage differentials. Section III presents a brief review of the relevant empirical literature. Section IV describes the estimation procedure, and section V delineates the NLSY data and summary statistics. Empirical results are presented in Section VI. Summary and conclusions follow in section VII.

II. Some Conceptual Discussions

II.1 Wage differential for cohabiting men

Over the past several decades there has been a considerable growth in the number of people making their first family union in cohabiting relations. As data from the March 2000 Current Population Survey (CPS) show, more than 3.8 million households were classified as unmarried-partner households representing about 4 percent of all households in the U.S.

[1] Blackburn and Korenman (1994) and Cohen (1999, 2002) examine marital wage premium for non-Whites, but they do not account for individual specific heterogeneity.

(Fields and Casper, 2001). To give a perspective to these numbers we may consider that in 1977 an estimated 1.1 million households were formed by cohabiting couples comprising about 1.5 percent of all households in that year (Casper and Cohen, 2000). Moreover, the percentage of marriages preceded by cohabitation rose from about 10 percent for those marrying during the period 1965-1974 to well over 50 percent for those marrying during 1990-1994 (Bumpass and Sweet 1989; Bumpass and Lu 2000). Bramlett and Mosher (2002), using the National Survey of Family Growth 1995 data, also identify that about a third of all ever married women have cohabited at some point in their lives. Altogether these studies indicate that over the years cohabitation is evolving to be a major mode of living arrangement – either as a transitional phase prior to a more permanent arrangement in marriage, or as an alternative family arrangement.

Considering such a remarkable rise in the prevalence of cohabitation, it would be appropriate to understand the mechanism through which cohabitation might influence men's earnings. To that end, let us briefly discuss how cohabitation compares with marriage as a family arrangement. This discussion would also point out why one may expect the effect of cohabitation on men's wages to be similar to that of marriage.

The functional aspects of these two forms of co-residential relationships are very similar. For instance, the type of intra-household specialization that occurs in marriage can also be hypothesized in a cohabiting union. Also, intra-household spillover effects of a partner's human capital endowment can be similar in married and cohabiting households. As a result, we might expect cohabiting men to demonstrate labor market experiences that are similar to those of a married man and significantly different from the experiences of non-cohabiting single men. The theoretical explanations that are offered for the existence of men's marital wage premium can all be equally applied for cohabiting men. Hence, a priori there are sufficient grounds to expect that cohabiting men would earn a wage premium in the labor market.

Despite these broad similarities, there are certain differences between marriage and cohabitation. The primary difference is legal. In the U.S., the majority of jurisdictions follow the general rule that unmarried cohabitating couples do not achieve legal rights or obligations analogous to those possessed by a married couple (Seff, 1995). Given the legal differences regarding marriage and cohabitation, cohabiting couples are less responsible for supporting one another than are married couples. Also, given the relatively higher cost of dissolving a marital union than a cohabiting union, the degree of commitment to the relationship could vary widely. As a result of these differences in the level of commitment and legal responsibilities, married men may choose to be more dedicated toward employment, increasing their effort, and hence productivity at work, and consequently, earn a higher wage premium than cohabiting men.

In the present study, we consider cohabitation as a co-residential union since the NLSY79 identifies cohabiters by the presence of an unmarried partner of the opposite sex in the household. However, cohabitation is recognized to be a particularly diverse arrangement that ranges from a short-term co-residential relationship considered to be an alternative to dating, to an alternative family arrangement in itself which is economically identical to marriage (Manning and Smock, 2003; Seltzer, 2000). The heterogeneous nature of cohabitation could cause the interpretation of the effect of cohabitation on men's wages to be somewhat ambiguous. We try to capture the heterogeneous nature of cohabitation in our empirical model by considering whether the cohabitation spell ends in marriage or in separation, and examine their effect separately on men's wage. However, we should note at the outset that the possibility of selection regarding the end status of a current cohabiting relationship demands cautious interpretation of these empirical results.

II.2 Explaining Family Union Premium in Men's Wage: A Theoretical Framework

Practically all cross-section estimates of human capital wage equations find that married men earn more than otherwise comparable single men. To explain the marital wage premium, a number of hypotheses have been put forward in the literature, and these may be broadly identified into two categories: selection hypothesis, and causal explanations. The selection hypothesis indicates that men with higher unobservable skills that are valued in the labor market select into marriage (Becker 1991, Nakosteen and Zimmer 1987, Korenman and Neumark, 1991). Causal explanations for marital wage premium include the following. First, the specialization hypothesis which suggests that marriage per se makes a man more productive by allowing him to specialize in non-household work, and pay differentials ensue from the differentials in productivity (Becker 1985, 1991). Second, employers discriminate against single men, and the marital wage premium reflects employer favoritism (Hill 1979, Bartlett and Callahan 1984). Third, the marriage premium is a reflection of the differences in the workers' taste and compensating wage differentials. According to this argument, married men have greater demand for costly family goods (such as children), and to meet the larger financial burden married workers choose jobs that offer greater wage and less attractive non-pecuniary compensation (Reed and Harford, 1989). And fourth, the joint human capital hypothesis, which suggests that the wife's human capital contributes to the husband's productivity and consequently married men have higher earnings (Benham, 1974).

Substantial research has been conducted to identify the selection bias in the estimated marital wage premium. The existing literature has also extensively dealt with the first three causal explanations, although the empirical tests do not provide compelling evidence for accepting or rejecting any of these three hypotheses. The literature, however, does not adequately explore the potential for the joint human capital hypothesis. In the present paper, we consider the joint-human capital hypothesis as an explanation for marital and cohabitation wage premiums.

We follow Benham (1974) in considering the effective stock of human capital for a man in a family union (H_t^*) to be a positive function of his own stock of human capital as well as his partner's stock of human capital:

$$H_t^* = H(H_t^O, H_t^P), \text{ where } \partial H_t^* / \partial H_t^O > 0 \text{ and } \partial H_t^* / \partial H_t^P > 0 .$$

Since a man's market (and non-market) productivity is a function of his effective stock of human capital, increments to the capital stock of the partner will be reflected in his productivity in the market, which would translate into increased wages for him. The presence of a better educated spouse or partner can contribute to men's work performance directly or indirectly, e.g., by assisting in the central task related to his job, by influencing decisions regarding job changes and transfers, by investing in his human capital, by assisting in his peripheral tasks such as social relations and networks at workplace etc. (Grossbard-Shechtman, 1993). Any of such contributions can substantially affect his earnings. Thus, married or cohabiting men would benefit from the presence of a better educated spouse or partner, and obtain a wage premium compared to their unmarried, non-cohabiting counterparts.

However, a positive relationship between a partner's human capital and men's earnings can arguably reflect selective-mating. The hypothesis of selective-mating asserts that the more productive men marry or cohabit with more highly educated women. In other words, men with higher observable and unobservable skills that are valued in the labor market might be more likely to partner with better educated women. Benham (1974) as well as Jepsen (2005) identified a positive association between wife's schooling and the husband's earnings, but due to the cross-sectional nature of their data, they were unable to conclusively remove the possibility that such positive association might reflect selective mating. Using longitudinal data one can address the marriage market sorting issue, at least to the extent that such sorting occurs over men's time-invariant unobservable characteristics.

III. Family Union Status and Wage Differentials: A Brief Literature Review

Partial summaries of cross-section studies on men's marital wage premium are available in Korenman and Neumark (1991), Nakosteen and Zimmer (1987), and Kenny (1983).[2] However, as has already been noted, cross-sectional studies are commonly criticized for their inability to address selection. Here we would present a brief review of the studies that directly address the selection issue in men's marital wage premium, and also consider research that looks into the effect of cohabitation on men's earnings.

III.1 Studies of wage differential for married men

Korenman and Neumark (1991), using data for white males from the 1976-1980 National Longitudinal Survey of Young Men (NLSYM), compare cross-sectional and fixed-effects estimates of the wage equation and show that selection on the basis of fixed

[2] A recent review of the literature on benefits of marriage is provided by Ribar (2004).

unobservable characteristics account for less than twenty percent of the observed wage premium. They also find that the marriage premium seems to arise slowly, resulting more from faster wage growth for married men as compared to never-married men than from an intercept shift associated with any particular marital status. They claim that their evidence is consistent with the productivity explanation for marital wage premium.

Cornwell and Rupert (1997), in contrast, conclude that selection underlies much of the marriage premium. They find that accounting for fixed individual characteristics substantially attenuates the marriage premium, and that any return to marriage is an intercept shift rather than a return to marriage over time. They consider this as evidence against the hypothesis that marriage enhances productivity through specialization.

Loh (1996) also suggests that labor productivity differences between married and never-married men are unlikely to be the cause of the marital wage premium. He postulates that men whose wives are active in the labor market must have relatively less scope for intra-household specialization, and hence, should receive a lower marriage premium. However, his cross-section estimates show that the size of the marriage wage premium did not vary with the nature of the wife's labor supply decision. These results can be challenged on grounds of biases originating from individual heterogeneity as well as endogeneity of wife's labor supply choice with respect to the husband's wage earnings. A more interesting result from his paper is that the wife's education underlies the positive wage premium earned by married men, although the estimated coefficients do not remain statistically significant in sibling fixed-effects estimation.

Gray (1997) uses data from the NLSYM 1966 and the National Longitudinal Survey of Youth 1979 to examine changes in the marriage premium between men in these two cohorts. Like Blackburn and Korenman (1994), Gray reports a significant decrease in the marriage premium over time. His results indicate that, for the earlier cohort, the marriage premium resulted from the productivity enhancing effects of marriage, but for the later cohort, the marriage premium reflects only selection.

Hersch and Stratton (2000) further examines the specialization hypothesis as they observe in the National Survey of Families and Households data that married men spent virtually the same amount of time on home production as did single men. Estimates from a fixed-effects wage equation show that the male marriage wage premium is not substantially affected by controls for home production activities. The authors conclude that the marriage premium of white men is not explained by either selection into marriage or specialization within the household. However, their results do not have direct bearing on whether or not marital wage premium reflect changes in the amount of time in the labor market by married men.

Ginther and Zavodny (2001) use 'shotgun weddings' to address the issue of selection, and find that selection accounts for less than ten percent of the marriage premium for white men. Their findings further suggest that men with conventional marriages have a larger time-invariant return to marriage, as measured by the intercept shift, while men with shotgun marriages earn a larger marriage premium for each additional year in marriage. The use of shotgun wedding as a natural experiment, however, may not be entirely convincing since pregnancies may not be unanticipated, and even if they are, the subsequent marriages could still be selective (Ribar, 2004).

Antonovic and Town (2004) use data from the Minnesota Twin Registry to study the impact of marriage on men's wages. Their within-twin-pairs estimate of the marital wage premium is as high as 26 percent, and is, in fact, higher than the estimated cross-section premium. While evidently the marital wage premium in their data is not a reflection of selection-bias, the inclusion of potentially endogenous variables, such as years of schooling, could produce biased estimates. The authors also do not provide any further analysis for such a surprisingly high marital wage premium.

III.2 Studies of wage differential for cohabiting men

Research on wage differentials for cohabiting men is considerably more limited. Both Daniel (1992) and Loh (1996) hypothesize that if the marital wage premium is due to intra-household specialization, then cohabiting men should also experience a wage differential. Daniel (1992), Loh (1996) and Cohen (1999, 2002) all report that currently cohabiting men receive significantly higher wages than non-cohabiting never married men. Cohen (1999) uses data from the 1995-97 Current Population Surveys (CPS) to report cross-section estimates showing that there are significant cohabitation benefits in wage earnings for white, black, and Hispanic men. Since none of these studies take into account the prospect of selection into cohabitation, one cannot infer on the extent to which the cohabitation premium in wages reflects selection into such a relationship. Further, these studies do not incorporate the duration of cohabitation as an explanatory variable, and hence are unable to identify the mechanism through which such relationships affect men's earnings.

Richardson (2000), using longitudinal data from the Swedish Level of Living Survey, identifies that even after accounting for selection through fixed effects estimation, Swedish men employed during the period 1968 to 1991 earned significant marriage and cohabitation premiums. The study further finds that the marriage and cohabitation premium for Swedish men has declined over time. The decline in marital premium has primarily been due to the diminishing productivity differences between married and single men.

Stratton (2002) examines the nature of the wage differentials arising from marriage and cohabitation for white men in the National Survey of Families and Households interviewed during 1987-88 and 1992-94. The paper treats marriage and cohabitation comparably, and finds that controlling for individual specific selection effects, the marital wage premium persists but the cohabitation premium disappears.[3]

Bardasi and Taylor (2005) investigate the relationship between family status and men's wages using data from the British Household Panel Survey. Their results from fixed effects as well as instrumental variables estimation suggest that while there is substantial marriage premium in men's earnings, cohabitation has no significant effect on men's earnings. The marriage premium from instrumental variables estimation is considerably higher than the least squares estimates. The authors claim their results as evidence in support of specialization hypothesis. However, they do not provide any further explanation on the absence of similar intra-household specialization in cohabiting families.

In summary, our review of the literature indicates that the marital wage premium is a debated issue – from the size of the premium to the underlying causes that can explain the

[3] While the use of fixed effects estimation techniques are appropriate in this paper, the inclusion of education in the fixed effect specification seems questionable since the author excludes the full-time enrolled respondents from the sample, and thereby makes the variables indicating levels of education time-invariant.

existence of such a premium. For cohabitation, cross-sectional evidence suggests the existence of a cohabitation premium in men's earnings. Furthermore, a handful of studies that have taken into account the potential selection bias do not provide clear evidence on whether cohabitation premium simply reflect selection. The review suggests that further research is needed to identify the effects of cohabitation on men's wage rate. The need for such explorations seems all the more relevant as the meaning of cohabitation is shifting from a stage prior to marriage to an alternative family arrangement in itself (Seltzer, 2000), thereby indicating a modification in the nature of selection of individuals into marriage and cohabitation.

IV. Empirical Methodology

A standard cross-section log wage regression augmented by controls for family status is the starting point for analyzing the effects of family status on wage. However, given that unobservable characteristics that enhance labor market achievements may also augment a man's prospects of finding a partner, either in marriage or in cohabitation, an improved model would be of the following form:

$$\ln(W_{it}) = \alpha + \beta \cdot X_{it} + \gamma \cdot FST_{it} + A_i + \varepsilon_{it} \qquad (1)$$

where W_{it} is the wage of individual i in year t, X_{it} is a vector of observable characteristics, FST_{it} is the family status of individual i in year t, and A_i is an unobserved characteristic of individual i, which is assumed to be time-invariant. The selection of men with wage-enhancing attributes into family union suggests that $Cov(FST_{it}, A_i) > 0$. Since A_i is unobservable, for ordinary least squares estimation, A_i would be part of the error term, and hence the estimated γ would be biased upwards. Using the panel structure of the data, we can employ a "within" or fixed effects estimation technique to remove the selection bias.[4] The estimation model in this case would be:

[4] One may also use random-effects model. However, test statistics from a Hausman test rejected random effects in favor of fixed effects in our data.

$$\ln(W_{it}) - \ln(\overline{W_i}) = \alpha + \beta \cdot (X_{it} - \overline{X_i}) + \gamma \cdot (FST_{it} - \overline{FST_i}) + \nu_{it} \quad (2)$$

where for any variable Z, $\overline{Z}_i$ denotes the mean of Z for individual i across the years t. We use the GLS estimator in Stata to estimate the fixed effects model.

As pointed out in Korenman and Neumark (1991), the benefits of a co-residential (marital or cohabiting) union may not accrue quickly and may depend upon the duration of the union. Consequently, we include measures of duration in different family status to examine whether the effect of marriage and cohabitation changes over time, and whether such duration effects differ across married and cohabiting men. These duration measures not only include duration in current status, but also duration in previous marriage and duration in cohabitation with current wife prior to marriage. The latter two measures of duration are intended to control for the effect of a man's life-course experience in co-residential unions on his earnings. Also, considering the plausible non-linear effects of relationship duration, we incorporate a quadratic term of each duration measure included in the specification.

In order to test the joint human capital hypothesis as a causal mechanism that can explain the family union premium in men's wages, one specification includes measures of partner or wife's level of schooling, considering schooling to be a measure of her stock of human capital. The schooling measures are included in the specification as interaction terms with marriage and cohabitation indicators. Thus the coefficients on these interaction terms would indicate whether the marital premium varies by the level of partner's schooling. However, we recognize that due to the possibility of selective-mating, partner's education is potentially endogenous in our specification. With the identifying assumption that such selective mating occurs over men's time-invariant characteristics, fixed effects estimates of the associated coefficients will be consistent.

We also explore the view that the impact on earnings of partner's human capital may change over the life cycle if communicating knowledge within the household improves over time, or if partner's education increases a man's earnings by reducing the depreciation rate of his stock of knowledge. Thus, in an extension of the previous specification, we interact partner or wife's level of schooling with the number of years the couple spent in marriage or in cohabitation.

Our specifications consciously exclude a number of variables such as tenure in current job, coverage by collective bargaining, occupation and industry, and presence of children in the household. Although these are commonly included in the related literature as explanatory variables, we feel that all of them are potentially endogenous to family status and wage.[5] Our estimates (not reported, but available upon request) from specifications that include tenure, tenure squared, dummy for union coverage, industry and occupation dummy variables, and number of children in the household show that the results remain qualitatively similar to those presented here.

V. NLSY79 Data and Summary Statistics

The National Longitudinal Survey of Youth 1979 (NLSY79) is a nationally representative sample of young men and women who were 14-22 years old when they were first interviewed (CHRR, 2001). The respondents were interviewed annually until 1994, and biennially since then. In this paper, we use data from the 12th through the 19th round (1990 – 2000) of the survey to construct a panel of repeated observations on individuals. Data from the later years of the survey are considered for two reasons: first, data on the duration of cohabiting relationships is available only for the survey rounds implemented since 1990; second, and more importantly, we would like to conduct our analyses based on the post-schooling labor market experience of the men in the sample and data from the later rounds can be considered to be generally more appropriate in this regard (the youngest men were 25 years old in 1990). In fact, we exclude all individuals who were ever enrolled in school during 1990 to 2000, so that our analyses would not be confounded by the individual's decision to work for a temporary period only to return to school later on. Thus, the sample is restricted to men who completed schooling by 1990, and for whom all the required variables are available.

[5] For example, Bratsberg and Terrell (1998) provide a discussion of the endogeneity of experience and tenure in a wage equation.

The analyses in this paper employ three stratified samples of men from the NLSY. The stratification is based on the broad definition of ethnicity: white, black, and Hispanic [6]. The white sample includes only those Non-black/Non-Hispanic men (identified by the survey screener) who self-identified themselves as ethnically European. Thus, we excluded the Native Americans, Asians and others from the Non-black/Non-Hispanic sample. All analyses are conducted separately for each of these ethnic groups. A formal assessment of the stratification was conducted using the Chow test which rejects, at less than 1 percent significance level, pooling of the data for the specifications reported in the paper. After meeting all data requirements, we have 1465 white, 1130 black and 682 Hispanic men in the sample.

The baseline (1990) summary statistics for all men in the sample are presented in Table 1 by their family union status. The mean values of the different variables in the pooled sample provide us with a succinct way of getting oriented with the sample, since the general features observed in the overall sample are observed in the ethnically stratified samples as well.[7] The first row of Table 1 presents mean hourly wages in dollars for men in different family contexts. The dependent variable in the regressions that follow is the natural logarithm of the hourly wage. In NLSY79 wages are available as hourly wages for hourly workers, and are constructed (in NLSY) from weekly or annual earnings divided by the appropriate hours for those who report non-hourly earnings. Married men appear to have much higher wages than men in the three other family status groups, viz., never married, currently cohabiting, and divorced or separated.

The figures in Table 1 also indicate that non-wage characteristics differ according to a man's family status. For example, never married and cohabiting men in the sample are younger than the married men. Married men worked about three hours more per week than never married men, about an hour more per week compared to cohabiting men, and about four weeks more per year than never married or cohabiting men. Also, compared to the

[6] The NLSY variable "Racial/Ethnic Cohort from Screener" (R0214700), "Racial/Ethnic Origin with Which R Identifies Most Closely" (R0010200) and "1st Or Only Racial/Ethnic Origin" (R0009600) are used to identify the individual's race.

[7] The interested reader may examine the Tables A-1, A-2, and A-3 in Appendix 1 to find the sample characteristics by race.

other groups, married men have enhanced job stability as indicated by the substantially higher mean tenure at current job. Between cohabiting and married men, the fraction of the sample having at least a high school graduate education is larger in the married group. On average married men are more likely to have wives who completed at least high school than the partners of cohabiting men. The figures in Table 1 also indicate that married men in this sample have spent, on average, about two years living together with their spouse prior to marriage, while the average length of current marriage for these men is little less than six years. The average length of cohabiting relationships for currently cohabiting men is about three years, which is substantially less than the average years in marriage for married men. The last two rows in Table 1 reflect that while a significant proportion of men who were never married in 1990 were married by 2000, the proportion of men in cohabiting unions remained unexpectedly stable over this period.

VI. Empirical Findings

Since selection has been the center of debate regarding wage differentials by marital status, in presenting our empirical findings we focus mostly on the fixed effects estimates of the different wage equation specifications. Parallel cross-sectional estimates are also produced to provide convenient comparison. All the specifications discussed include age, age squared, dummy variable for residence in Standard Metropolitan Statistical Area (SMSA), unemployment rate for labor market of current residence, three dummy variables for region of residence, and seven year dummy variables. The empirical findings for the three ethnic groups are reported separately in Tables 2, 3 and 4.

VI.1. Results for white men

The first column in Table 2 presents the fixed effects estimates of a wage equation specification which contains two marital status dummy variables (viz., married, and divorced or separated) for white men and the common set of covariates indicated above, but does not control for cohabitation status. These results, therefore, correspond to the longitudinal estimates of the marriage premium available in the literature. The estimates show that controlling for the other covariates, white married men earn about 5 percent more than the unmarried men. A comparison with the cross-section estimates of the same specification in column 6 shows that a large proportion of the cross-section marital wage premium is associated with individual specific selection effects, the estimated cross-section premium being more than 20 percent. The estimated coefficient for the divorced or separated dummy is not statistically significant. These fixed effects estimates are very similar to those in Korenman and Neumark (1991) who report a 6 percent marriage premium and no significant effect of divorce/separation, when any control for duration of relationships is not included. Longitudinal estimates in Stratton (2002) with a similar specification, however, indicate that there is no significant effect of marriage on men's earnings.

Columns 2 and 7 of Table 2 present the estimates for the specification which directly includes an indicator variable for cohabitation as an additional family union status. We find that both cross-section and longitudinal estimates of the marital wage premium are a little larger than the corresponding estimates with cohabiting men in the reference group. If we consider cohabitation to be somewhat closer to marriage as a household structure than non-cohabiting singlehood, we might expect the marital wage premium to rise as we change the reference group from never married men to non-cohabiting never married men. These estimates suggest that inclusion of cohabitation as a control might reduce the dramatic reduction in men's marital wage premium reported in Gray (1997) who uses longitudinal estimation techniques on NLSY79 data for the period 1989-93, and shows that there is no statistically significant marital wage premium.[8] This is also consistent with the results in Cohen (2002).

More interestingly, cross-section estimates show that cohabiting white men earn 6.5 percent more than the non-cohabiting never married men. The size of the cohabitation wage premium is about a third of the marital wage premium identified in the same equation. These cross-section estimates are very similar to those in Cohen (1999). The fixed effects estimate

---

[8] Gray (1997) uses specifications that included actual experience, union, child, occupation and industry dummy variables. As has already been indicated, even with a very comparable specification, our results do not change in any significant way.

of the cohabitation premium is smaller (3.2 percent) than the cross-section estimate, but it is still statistically significant. This essentially indicates that for white men only about half of the cross-sectional cohabitation premium reflects selection into cohabitation. Our fixed effects estimate of the cohabitation premium contradicts Stratton's (2002) findings that show the entire cohabitation premium to reflect individual selection effects.

We recognize that fixed effects estimation is only able to account for selection through time-invariant unobservable characteristics. If selection into marriage or cohabitation depends on wage growth (men with high wage growth are more likely to be in married or cohabiting), then changes in wages and family status will be interdependent, and even the fixed effects coefficients will be biased upwards. To identify whether such selection through wage-growth are important for our estimated marriage and cohabitation premiums, we examine the pre-marriage (pre-cohabitation) wage growth for men who married (cohabited) during the sample period versus other non-married (non-cohabiting) men. More specifically, we estimate two separate equations of the following form for ever-married and ever-cohabited men during their pre-union period:

$$\ln(W_{it+1}) - \ln(W_{it}) = \alpha + \beta \cdot (X_{it+1} - X_{it}) + \delta F_i + (\eta_{it+1} - \eta_{it}) \qquad (3)$$

where W_{it} is the wage of individual i in year t, X_{it} is a vector of observable characteristics[9], η_{it} is the error term. F_i is 1 if the man married (cohabited) over the sample period , zero otherwise. The sample is restricted to pre-marriage (pre-cohabitation) period of men who were single at period t. The estimated coefficient $(\hat{\delta})$ is -0.004 (with a t-statistic of 0.27) for white men who married later on, and -0.009 (with a t-statistic of 1.18) for white men who cohabited later during the sample period. These results suggest that white men who married (cohabited) during the sample period did not exhibit higher pre-marriage (pre-cohabitation) wage growth than otherwise similar white men. Therefore, the selection into marriage and cohabitation can safely be considered to be independent of wage growth.

In the next specification in Table 2, measures of duration in different family status are included. The coefficients on the various duration measures indicate the speed of wage

[9] Variables included in this pooled cross-section analysis are: age, tenure, local unemployment rate, dummy for SMSA, and regional dummy variables (3).

growth in different types of family status as compared to non-cohabiting never-married men. Fixed effects and least squares estimates for this specification are produced, respectively, in columns 3 and 8 of Table 2. Comparison of the fixed effects estimates with and without the duration controls reveal, not too surprisingly, that for white men there is no statistically significant marital *intercept* shift, and the effect of marriage is enhanced only gradually, through the steepening of the wage profile over the length of the marriage as wages grow 1.6 percent for each additional year of marriage. The marriage duration effect on wage declines over the years, as reflected in the statistically significant coefficient on marital duration squared, but only at a very minimal rate. Cross-sectional estimates reveal a similar result, although the intercept shift is statistically significant and quite large (reflecting a selection bias). While the sign of the estimated coefficients on the two other quadratic duration terms measuring years cohabited with current wife prior to marriage and years in earlier marriages are in the expected direction, none of them are statistically significant.

The cross-section estimates of the coefficients on the cohabitation dummy as well as the linear and quadratic measures of duration in cohabitation have the expected signs, but are not statistically significant. The parallel fixed effects point estimates are somewhat smaller than the cross-section estimates, indicating a positive selection bias, but these again are not statistically significant. These results suggest that there is not a significant wage growth associated with duration of cohabitation, perhaps due to the relatively transient nature of such relationships.

It would be appropriate to note that even though we have some evidence of a cohabitation premium in men's earnings, we have recognized earlier that cohabitation is heterogeneous in nature. In an attempt to distinguish between different types of cohabiting relationships, we identify three categories of cohabitation in our data: cohabitation ending in marriage, cohabitation ending in separation, and continued cohabitation when the respondent was last interviewed. We replace the cohabitation dummy variable in our specifications with the indicators for these three types of cohabitation. The results for white men, shown in column 1 of Table 5, indicate that the cohabitation premium accrues only to those who get married later on, but not to men in the other two categories of cohabitation. These results do not change significantly even when the specification includes years in different family status as shown in column 2 of Table 5. The impact of cohabitation that ends in marriage remains constant over time, since we find no statistically significant effect of the duration in cohabitation on white men's earnings. Although these results are quite intriguing, they need to be considered with caution. The three cohabitation indicators in this specification contain information about how a current spell of cohabitation ends in the future. These regressors may not be strictly exogenous in the sense that future values of the dependent variable (i.e., future wages) can influence the current value of these indicator variables. In other words, how a cohabitation spell ends in the future might be endogenous to the individual's future wages. Consequently, the dummy variables indicating the three types of cohabitation are potentially endogenous to wages, and therefore, the estimated coefficients may fail to be consistent.

The results presented so far indicate that both marriage and cohabitation have a positive effect on men's wage earnings. These findings generally support the hypothesis that family union raises men's labor market productivity. In order to identify a causal mechanism that connects a man's family status and his productivity, we proposed a joint human capital hypothesis. Hence, our next specification includes a set of dummy variables indicating the educational attainment of the married or cohabiting partner as a measure of her human capital stock. The relevant fixed effects and least squares results are presented, respectively, in columns 4 and 9 in Table 2. Note that the married and cohabitation dummies now measure the returns to men whose (marital or cohabiting) partners are not high school graduates.

Least squares estimates (in column 9) show that the higher the level of education of the wife, the better-off married men are. Fixed effects estimates are much smaller than least squares estimates, reflecting selection and positive sorting in the marriage market. However, even after addressing the selection prospects, men who are married to high school graduate wives have a statistically significant wage premium. For men with a wife who has more education than a high school degree, although there is no significant intercept shift in his wage profile from wife's education, the impact changes over the duration of the marriage as indicated by the coefficients on the interaction of marital duration and wife's education (column 5 in Table 2). A similar effect of partner's education is identified among cohabiting men. After controlling for selection and sorting into cohabitation, men with a partner who completed high school earn about 10 percent more than non-cohabiting never-married men. For men with a less than high school educated partner the positive effect accrues over time as indicated by the coefficient on cohabitation duration variable in column 5. However, the effect of a cohabitating partner's education over time decreases for those who have a partner with at least high school education.

Thus, accounting for selection and assortative mating as well as duration in current family status, we have some evidence of intra-household spillover effects of a female partner's education.[10] We should note that the last two specifications also included a man's own education interacted with their family status. Clearly, controlling for the effect of man's own education on the likelihood of being married or cohabiting does not remove the household spillover effects of a female partner's education.

As has been noted earlier, the identifying assumption for consistently estimating the fixed-effects coefficients on partner's education is that any selective mating occurs over men's time-invariant characteristics. However, partner's education can potentially be associated with men's time-varying characteristics which are not included in the specification. More specifically, men with a higher educated partner may be more likely to increase work hours, or to receive on-the-job-training, or are less likely to have quick job-turnovers. Any of this set of changes can have a positive effect on men's wages. Therefore, if such association exists, the estimated positive spillover effects of partner's education may reflect the potential positive influence of these time-varying factors. In order to identify whether the estimated effects of partner's education represents the underlying association with hours worked, incidence of training, and tenure, we look into their relationship by estimating three separate fixed-effects equations for men's weekly hours worked, on-the-job-training incidence, and tenure with current employer. Estimates for white men in Table 6 suggest that the nature of the association between partner's education and any of these three measures could not explain the positive spillover effect on men's wages. For cohabiting men,

[10] Basu, Narayan and Ravallion (2002) find similar spillover benefits of literacy in the household in Bangladesh, while Neuman and Ziderman (1992) identify cross-productivity effects of women's education in higher status occupations in Israel. Tiefenthaler (1997) also finds some evidence of spillover effects of spousal education in Brazil.

partner's education does not significantly influence their tenure in current job, or weekly hours worked, and the negative effect a college graduate partner has on a man's propensity to receive on the job training only indicates that the positive spillover effect of partner's education on men's wage is not a reflection of training. For married men, wife's education has no significant influence on tenure and propensity to receive training. Men who have a wife with more than high school education work fewer hours per week, so that the effect of a better educated wife on their earnings may not originate from increased working hours. Thus, these auxiliary estimates help us to maintain that the spillover effect of partner's education on white men's wages does not embody the relationship of partner's education with these time-varying characteristics.

VI.2. Results for black men

The cross-sectional and longitudinal estimates of the first specification (columns 1 and 6 in Table 3) which does not include a cohabitation dummy, show that for black men, individual specific selection explains the entire marital wage premium. This is an important difference from the results for white men. As we include a cohabitation dummy in the specification, pooled cross-section estimates (column 7 in Table 3) show that married men earn about 24 percent more, while cohabiting men earn about 5 percent more than non-cohabiting never married men. The fixed effects coefficient on the married dummy is larger than the initial specification but is still not significant, and thus again indicates that the cross-sectional intercept-shift associated with marriage for black men reflects only selection. However, even after accounting for selection, cohabiting men earn a 3.7 percent wage premium.

We examine the case of selection through wage growth for black men as well, and estimate equation (3) for black men. The estimated coefficient $(\hat{\delta})$ is -0.014 (with a t-statistic of 1.11) for black men who married later on, and -0.009 (with a t-statistic of 1.03) for black men who cohabited later during the sample period. Therefore, the selection into marriage and cohabitation for black men can safely be considered to be independent of wage growth.

As we add duration in each family status in the next specification in Table 3, none of the family status dummy variables, except the cross-sectional coefficient on the married dummy, has a significant effect. The fixed effects results for cohabiting black men are similar to those for cohabiting white men – neither the cohabitation dummy nor the duration in the relationship are statistically significant. The longitudinal estimates also show that duration in current and previous marriages as well as years divorced/separated have a statistically significant effect on men's wages. Thus, for black men, there is no significant intercept shift associated with any particular marital status. Rather, the effect of marriage accrues at a diminishing rate over time.

Once again, in order to distinguish between different types of cohabiters among black men, we include indicators of three types of cohabitation in the specification and the results are given in columns 3 and 4 of Table 5. Unlike white men, even for cohabiting black men who later marry their partner, no significant intercept-shift occurs in their wage profile as they cohabit. However, wages for Blacks grow by about 1.4 percent during each year of cohabitation. The differences in the results for white and Blacks are noteworthy given the lower rates of marriage, and the higher incidence of prolonged cohabiting relationship among Blacks in the U.S. (Tucker and Mitchell-Kernan, 1995). In our data as well, the duration in cohabitation for black cohabiting men is more than a year higher compared to that of white cohabiting men. Therefore, the wage growth that occurs for the cohabiting Blacks reflects the accrual of benefits over prolonged cohabiting relationships, while no significant wage growth arises from the relatively short-lived cohabitation among the Whites.

In the next specification, partner's education dummies are interacted with marriage and cohabitation indicators. We find that the pooled cross-section results are quite similar to those for Whites. black men with better-educated partners (in marriage or cohabitation) earn significantly higher wages than non-cohabiting never married men. However, the fixed effects point estimates of almost all the relevant coefficients are negative, but none are statistically significant. Thus, the cross-sectional benefit for black men having a better-educated partner mostly reflects selection and positive sorting in marriage and cohabitation. As can be seen from the results in column 5 of Table 3, only black men whose wife is a college graduate receive any kind of significant benefit, as in this case the effect of wife's education becomes stronger over time. While this latter finding is not driven by work hour increases or on the job training, we may not rule out the argument that black men with a college graduate wife are more likely to stay in the same job, and hence receive enhanced wages (as shown by estimates in Table 6).

VI.3. Results for Hispanic men

The Hispanic men have so far been overlooked in the longitudinal analyses of human capital wage equations. The results available here for Hispanic men (Table 4) are expected to reduce the void in the literature on family status and wage differentials.

The dummy variable specifications, with and without the cohabitation dummy, show that in the pooled cross-section married men earn about 25 percent more than never married men, while divorced/separated men earn about 11 percent more than never married men. As fixed effects estimates produce no significant coefficients on the married and divorce/separated dummy variables, all relevant cross-section intercept-shifts in Hispanic men's wage profile, thus, reflect selection. Also, cohabiting Hispanic men are no better off than non-cohabiting never married Hispanic men as shown by least squares as well as fixed effects estimates (column 2 and 7 in Table 4).

Results for the duration specification show that controlling for selection, the number of years spent in any family status has no significant effect on Hispanic men's wages. Estimates from the specification that includes partner's educational attainment dummies show results similar to those for black men. Cross-sectional estimates indicate that Hispanic men with better educated partners have improved experiences in the labor market. However, fixed effects estimates remove significance from these estimates, thus indicating selection into marriage and cohabitation.

VII. Conclusion

We recognize that with the nature of family arrangements in the U.S. changing rapidly, men's choices in the family sphere and the labor market are ever increasingly intertwined. The challenge of untangling this simultaneity is not easy to overcome. Such challenges notwithstanding, the present study is expected to further our understanding of the motivation for men's choices regarding cohabitation and marriage, and the implications thereof. The paper provides new evidence on the male wage premiums in relation to marriage and cohabitation, and has three principal findings to report.

First, hourly wage premium paid to married men are large in cross-sectional estimates for men in all ethnic groups. However, once we address the selection issue through fixed effects estimation, the marital premium persist for white and black men. Also, the marital wage premium arises slowly and is apparently the outcome of wage growth over the length of the marriage.

Second, cross-section estimates indicate the presence of a cohabitation premium, albeit smaller than the marital premium, for black and white men. Fixed effects estimates show that for white men about one-half, and for black men less than one-third of the cross-sectional cohabitation premium is associated with the individual's fixed unobservable characteristics that are positively correlated with both cohabitation and wages. More precisely, even after accounting for individual specific unobservable characteristics, white and black men earn a cohabitation wage premium of 3.2 percent and 3.7 percent, respectively. Hispanic cohabiting men are no better off than non-cohabiting men as reflected by both cross-section and fixed effects estimates. Comparison of estimates from specifications with and without measures of cohabitation-duration indicates that while cohabitation is associated with improved wage earnings, such improvements are not manifested in terms of wage growth over the length of cohabitation. We also have some weak evidence indicating that the cohabitation premium appears to accrue to white men who eventually marry their partner, while for black men the cohabitation premium is an outcome of wage growth over the length of the cohabiting relationship.

Finally, the most intriguing finding in this paper is the positive contribution of wife's or partner's education to men's hourly wage. Cross-sectional estimates show that for all ethnic groups, men with a better-educated wife/partner earn significantly more than non-cohabiting never married men. Fixed effects estimates, however, reveal that for black and Hispanic men all of the cross-sectional effects are reflections of selection on the basis of fixed unobservable characteristics of married and cohabiting men. For white men, even fixed effects estimates show a positive effect of marital or cohabiting partner's education on men's wage rate.

Our findings are consistent with the hypothesis that family union enhances men's labor market productivity. More importantly, we argue that in a joint human capital framework, intra-household spillover effects of partner's education would provide a causal mechanism linking family union status and men's productivity. With rising female labor force participation in the recent decades, the joint human capital hypothesis appears to be more appealing than a more traditional specialization hypothesis to explain wage premiums for married or cohabiting men. Our results provide some empirical support for the joint human capital hypothesis.

Table 1. Summary Characteristics of Sample by Family Union Status NLSY All Men, 1990

Variables	Never Married	Cohabiting [a]	Married	Divorced or Separated
Hourly wage (in dollars) [b]	9.68	9.43	11.47	9.14
Hours worked per week	42.58	44.25	45.39	44.43
Weeks worked per calendar year	44.40	44.62	48.53	44.04
Age (in years)	28.44	28.56	29.31	29.46
Yrs in marriage, total	-	1.58	5.96	5.58
Yrs in current marriage	-	-	5.61	-
Yrs div. or sep.	-	0.63	-	1.82
Yrs cohabiting	0.57	3.09	-	0.36
Yrs cohabited with wife before marriage	-	-	2.10	-
Has schooling level				
Less than high-school (<12)	0.19	0.25	0.15	0.22
High-school grad (=12)	0.46	0.52	0.49	0.58
Some college (>12 & <16)	0.17	0.15	0.18	0.15
College grad (>=16)	0.18	0.09	0.18	0.05
Partner's schooling level				
Less than high-school (<12)	-	0.21	0.12	-
High-school grad (=12)	-	0.50	0.47	-
Some college (>12 & <16)	-	0.17	0.21	-
College grad (>=16)	-	0.10	0.19	-
Missing	-	0.03	0.02	-
Lives in the				
North east	0.20	0.18	0.17	0.16
North central	0.22	0.21	0.25	0.23
South	0.39	0.38	0.35	0.42
West	0.19	0.23	0.22	0.19
Lives in SMSA	0.79	0.81	0.76	0.80
Local unemployment rate	2.32	2.31	2.38	2.33
Years in current job	2.56	2.63	4.08	2.73
Is covered by union	0.17	0.18	0.22	0.21
black	0.46	0.47	0.21	0.36
Hispanic	0.18	0.20	0.24	0.20
Sample Size (1990)	1081	234	1340	311
Proportion of Sample in different status:				
1990	0.40	0.09	0.49	0.11
2000	0.22	0.07	0.59	0.19

Note: a. About 71 percent of currently cohabiting men are never married.
b. For our sample we dropped the cases where hourly wages fell outside the $1 - $100 range.

Table 2. NLSY Wage Regression: White Males 1990-2000
Dependent Variable: Ln(Hourly Wage)

	Longitudinal					Cross-Sectional				
	Dummy var. spec (w/o cohabit)	Dummy var. spec	Duration Spec	Partner's Edu spec	Partner's Edu* Dur spec	Dummy var. spec (w/o cohabit)	Dummy var. spec	Duration Spec	Partner's Edu spec	Partner's Edu* Dur spec
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Married	0.052	0.061	0.034	-0.019	0.002	0.214	0.223	0.125	-0.043	0.001
	(2.73)***	(3.08)***	(1.53)	(0.32)	(0.03)	(8.57)***	(8.59)***	(3.79)***	(0.54)	(0.01)
Divorced/separated [c]	-0.005	-0.001	0.011	0.060	0.082	0.053	0.052	0.070	0.057	0.087
	(0.22)	(0.06)	(0.39)	(0.98)	(1.30)	(1.65)*	(1.63)	(1.48)	(0.67)	(0.95)
Cohabit		0.032	0.013	-0.017	-0.102		0.065	0.045	0.045	-0.076
		(1.72)*	(0.47)	(0.27)	(1.39)		(2.24)**	(1.16)	(0.53)	(0.84)
Yrs in current marriage			0.016	0.016	0.011			0.020	0.019	0.014
			(3.77)***	(3.82)***	(2.02)**			(3.38)***	(3.22)***	(1.91)*
Yrs current marr – sq			-0.001	-0.001	-0.001			-0.001	-0.001	-0.001
			(3.46)***	(3.54)***	(3.00)***			(2.05)**	(1.80)*	(1.87)*
Yrs div/sep			0.004	0.004	0.002			0.010	0.012	0.009
			(0.49)	(0.48)	(0.26)			(0.82)	(0.97)	(0.76)
Yrs div/sep, sq			-0.000	-0.000	-0.000			-0.001	-0.001	-0.001
			(0.44)	(0.39)	(0.16)			(1.13)	(1.32)	(1.17)
Yrs cohabit			0.017	0.017	0.050			0.011	0.011	0.047
			(1.41)	(1.38)	(2.73)***			(0.66)	(0.64)	(2.29)**
Yrs cohabit, sq			-0.001	-0.002	-0.002			-0.001	-0.001	-0.001
			(1.52)	(1.67)*	(2.14)**			(0.49)	(0.44)	(1.26)
Yrs in previous marriage			0.005	0.007	0.003			-0.010	-0.010	-0.016
			(0.67)	(1.05)	(0.39)			(1.14)	(1.13)	(1.54)
Yrs prev marriage, sq			-0.000	-0.000	-0.000			0.001	0.001	0.001
			(0.49)	(0.94)	(0.66)			(1.58)	(1.61)	(1.86)*
Yrs cohab w/ wife			0.001	0.002	0.001			0.009	0.010	0.010
			(0.07)	(0.25)	(0.12)			(0.88)	(0.99)	(0.96)
Yrs cohab w/ wife, sq			0.000	0.000	0.000			-0.001	-0.001	-0.001
			(0.39)	(0.27)	(0.35)			(0.63)	(0.61)	(0.56)
Married * Wife HS grad				0.069	0.068				0.143	0.078
				(1.97)**	(1.59)				(3.07)***	(1.23)
Married * Wife Somecoll				0.022	-0.007				0.141	0.085
				(0.58)	(0.14)				(2.75)***	(1.16)

Table 2 (cont'd).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Married * Wife Collgrad				-0.001	-0.064				0.180	0.140
				(0.02)	(1.37)				(3.38)***	(1.93)*
Coh * Partner HSgrad				0.116	0.212				0.100	0.229
				(2.21)**	(3.08)***				(1.37)	(2.45)**
Coh * Partner Somecoll				0.024	0.073				0.100	0.230
				(0.38)	(0.87)				(1.09)	(1.90)*
Coh * Partner Collgrad				0.006	0.093				0.210	0.317
				(0.08)	(1.02)				(2.03)**	(2.43)**
MarrDur * Wife HSgrad					0.001					0.007
					(0.22)					(1.37)
MarrDur * Wife Somecoll					0.005					0.006
					(1.32)					(1.00)
MarrDur * Wife Collgrad					0.010					0.004
					(2.65)***					(0.65)
CohDur * Part HSgrad					-0.035					-0.034
					(2.38)**					(2.72)***
CohDur * Part Somecoll					-0.021					-0.037
					(1.01)					(1.18)
CohDur * Part Collgrad					-0.060					-0.026
					(2.11)**					(0.69)
Observations	9969	9969	9969	9969	9969	9969	9969	9969	9969	9969
Number of id	1465	1465	1465	1465	1465					
R-squared	0.25	0.25	0.25	0.25	0.25	0.29	0.29	0.29	0.30	0.30

Note:

a. Also included in the longitudinal specifications are: age, age-squared, dummy for SMSA, local unemployment rate, and regional dummy (3), and year dummy variables (7). In addition, specifications in columns 4, 5, 9, and 10 include own education interacted with family status. The full set of estimates is available upon request.

b. The pooled cross-section regressions include 3 schooling dummy variables along with the controls listed above for the longitudinal specifications; and the standard errors for this set of estimates are robust.

c. Divorce and separated category includes a small number of widowers.

d. Absolute value of t-statistic in parentheses (* significant at 10%; ** significant at 5%; *** significant at 1% level, in a two-tailed test).

Table 3. NLSY Wage Regression: Black Males 1990-2000

Dependent Variable: Ln(Hourly Wage)

	Longitudinal					Cross-Sectional				
	Dummy var. spec (w/o cohabit)	Dummy var. spec	Duration Spec	Partner's Edu spec	Partner's Edu* Dur spec	Dummy var. spec (w/o cohabit)	Dummy var. spec	Duration Spec	Partner's Edu spec	Partner's Edu* Dur spec
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Married	0.019	0.032	0.004	-0.041	-0.032	0.235	0.245	0.122	0.105	0.126
	(0.84)	(1.33)	(0.14)	(0.64)	(0.49)	(10.98)***	(11.09)***	(3.59)***	(1.46)	(1.53)
Divorced/separated [c]	0.019	0.026	0.007	-0.110	-0.094	0.085	0.087	0.060	0.051	0.068
	(0.66)	(0.92)	(0.20)	(1.67)*	(1.39)	(2.94)***	(3.02)***	(1.22)	(0.70)	(0.89)
Cohabit		0.037	0.020	0.082	0.072		0.054	0.043	-0.007	-0.070
		(2.12)**	(0.78)	(1.59)	(1.18)		(2.08)**	(1.32)	(0.10)	(0.90)
Yrs in current marriage			0.015	0.015	0.011			0.027	0.025	0.022
			(2.55)**	(2.49)**	(1.60)			(3.64)***	(3.43)***	(2.40)**
Yrs current marr – sq			-0.001	-0.001	-0.001			-0.001	-0.001	-0.001
			(2.24)**	(2.27)**	(2.26)**			(2.44)**	(2.19)**	(2.06)**
Yrs div/sep			-0.017	-0.018	-0.018			-0.024	-0.024	-0.024
			(1.75)*	(1.79)*	(1.82)*			(1.58)	(1.59)	(1.57)
Yrs div/sep, sq			0.002	0.002	0.002			0.002	0.002	0.002
			(2.07)**	(2.05)**	(2.08)**			(1.78)*	(1.85)*	(1.82)*
Yrs cohabit			0.008	0.005	0.007			0.003	0.001	0.014
			(0.85)	(0.60)	(0.59)			(0.28)	(0.08)	(0.94)
Yrs cohabit, sq			-0.000	-0.000	-0.000			-0.000	0.000	0.000
			(0.38)	(0.20)	(0.19)			(0.13)	(0.33)	(0.02)
Yrs in previous marriage			0.022	0.021	0.016			0.010	0.009	0.006
			(2.59)***	(2.47)**	(1.73)*			(0.98)	(0.93)	(0.48)
Yrs prev marriage, sq			-0.001	-0.001	-0.001			-0.000	-0.000	-0.000
			(2.25)**	(2.24)**	(1.81)*			(0.42)	(0.38)	(0.10)
Yrs cohab w/ wife			0.017	0.016	0.017			0.018	0.017	0.017
			(1.61)	(1.57)	(1.63)			(1.50)	(1.39)	(1.38)
Yrs cohab w/ wife, sq			-0.002	-0.002	-0.002			-0.002	-0.002	-0.002
			(1.77)*	(1.68)*	(1.75)*			(2.32)**	(2.09)**	(2.13)**
Married * Wife HS grad				-0.028	-0.030				0.036	0.039
				(0.67)	(0.62)				(0.69)	(0.55)
Married * Wife Somecoll				-0.019	-0.050				0.108	0.078
				(0.40)	(0.87)				(1.87)*	(0.97)

Table 3 (cont'd).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Married * Wife Collgrad				-0.026	-0.075				0.187	0.095
				(0.55)	(1.31)				(3.08)***	(1.11)
Coh * Partner HSgrad				-0.011	0.017				0.106	0.180
				(0.25)	(0.29)				(1.65)*	(2.47)**
Coh * Partner Somecoll				0.032	0.002				0.180	0.257
				(0.59)	(0.02)				(2.21)**	(2.93)***
Coh * Partner Collgrad				-0.008	-0.081				0.210	0.223
				(0.13)	(1.02)				(2.16)**	(1.71)*
MarrDur * Wife HSgrad					0.001					0.000
					(0.34)					(0.05)
MarrDur * Wife Somecoll					0.006					0.004
					(1.25)					(0.59)
MarrDur * Wife Collgrad					0.010					0.013
					(1.80)*					(1.47)
CohDur * Part HSgrad					-0.008					-0.014
					(0.82)					(1.13)
CohDur * Part Somecoll					0.006					-0.016
					(0.49)					(0.99)
CohDur * Part Collgrad					0.021					-0.002
					(1.34)					(0.10)
Observations	6969	6969	6969	6969	6969	6969	6969	6969	6969	6969
Number of id	1130	1130	1130	1130	1130					
R-squared	0.19	0.20	0.20	0.20	0.20	0.27	0.27	0.28	0.29	0.29

Note:

a, b, c: See notes in Table 2.

d. Absolute value of t-statistic in parentheses (* significant at 10%; ** significant at 5%; *** significant at 1% level, in a two-tailed test).

Table 4. NLSY Wage Regression: Hispanic Males 1990-2000

Dependent Variable: Ln(Hourly Wage)

	Longitudinal					Cross-Sectional				
	Dummy var. spec(w/o cohabit)	Dummy var. spec	Duration Spec	Partner's Edu spec	Partner's Edu* Dur spec	Dummy var. spec (w/o cohabit)	Dummy var. spec	Duration Spec	Partner's Edu spec	Partner's Edu* Dur spec
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Married	0.019	0.019	0.008	0.006	0.077	0.252	0.252	0.040	-0.037	0.005
	(0.59)	(0.56)	(0.21)	(0.10)	(1.24)	(7.88)***	(7.77)***	(0.75)	(0.45)	(0.06)
Divorced/separated [c]	0.026	0.025	0.049	-0.009	0.060	0.110	0.110	-0.013	-0.024	0.003
	(0.63)	(0.63)	(0.87)	(0.13)	(0.81)	(2.67)***	(2.67)***	(0.17)	(0.26)	(0.03)
Cohabit		-0.002	0.020	0.025	-0.033		0.002	-0.005	-0.083	-0.181
		(0.06)	(0.58)	(0.38)	(0.44)		(0.04)	(0.08)	(1.04)	(2.06)**
Yrs in current marriage			0.009	0.008	-0.003			0.033	0.033	0.027
			(1.43)	(1.31)	(0.50)			(3.50)***	(3.58)***	(2.94)***
Yrs current marr – sq			-0.000	-0.000	-0.000			-0.001	-0.001	-0.001
			(0.71)	(0.56)	(0.19)			(1.79)*	(1.53)	(1.34)
Yrs div/sep			-0.013	-0.012	-0.017			-0.013	-0.014	-0.016
			(1.09)	(1.02)	(1.38)			(0.70)	(0.77)	(0.89)
Yrs div/sep, sq			0.000	0.000	0.001			0.001	0.001	0.001
			(0.43)	(0.35)	(0.63)			(0.59)	(0.72)	(0.81)
Yrs cohabit			-0.003	-0.002	0.005			0.004	0.002	0.017
			(0.49)	(0.41)	(0.71)			(0.51)	(0.32)	(1.80)*
Yrs cohabit, sq			0.000	0.000	0.000			-0.000	-0.000	-0.000
			(0.03)	(0.04)	(0.95)			(1.06)	(1.08)	(0.15)
Yrs in previous marriage			0.008	0.008	-0.004			0.023	0.023	0.018
			(0.81)	(0.72)	(0.37)			(2.21)**	(2.23)**	(1.62)
Yrs prev marriage, sq			-0.000	-0.000	0.000			-0.000	-0.000	-0.000
			(0.58)	(0.54)	(0.16)			(0.76)	(0.62)	(0.30)
Yrs cohab w/ wife			-0.011	-0.012	-0.012			0.009	0.005	0.003
			(0.79)	(0.87)	(0.87)			(0.47)	(0.26)	(0.17)
Yrs cohab w/ wife, sq			0.001	0.001	0.001			-0.001	-0.000	0.000
			(0.96)	(1.01)	(0.97)			(0.37)	(0.05)	(0.02)
Married * Wife HS grad				-0.034	-0.112				0.150	0.073
				(0.84)	(2.18)**				(3.51)***	(1.03)
Married * Wife Somecoll				0.006	-0.090				0.225	0.259
				(0.12)	(1.47)				(3.99)***	(2.88)***

Table 4 (cont'd).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Married * Wife Collgrad				-0.057	-0.202				0.391	0.219
				(0.91)	(2.70)***				(5.33)***	(2.03)**
Coh * Partner HSgrad				-0.080	-0.015				0.155	0.265
				(1.35)	(0.18)				(1.88)*	(2.64)***
Coh * Partner Somecoll				-0.129	-0.081				0.317	0.513
				(1.58)	(0.77)				(2.62)***	(3.11)***
Coh * Partner Collgrad				-0.182	-0.210				0.372	0.373
				(1.51)	(1.35)				(1.84)*	(1.70)*
MarrDur * Wife HSgrad					0.010					0.008
					(2.57)**					(1.35)
MarrDur * Wife Somecoll					0.013					-0.005
					(2.69)***					(0.56)
MarrDur * Wife Collgrad					0.020					0.020
					(3.37)***					(2.17)**
CohDur * Part HSgrad					-0.017					-0.020
					(1.71)*					(1.92)*
CohDur * Part Somecoll					-0.014					-0.029
					(1.06)					(1.68)*
CohDur * Part Collgrad					-0.001					0.004
					(0.04)					(0.14)
Observations	4405	4405	4405	4405	4405	4405	4405	4405	4405	4405
Number of id	682	682	682	682	682					
R-squared	0.19	0.19	0.19	0.19	0.20	0.26	0.26	0.27	0.29	0.30

Note:
a, b: See notes in Table 2.
c. There were no widowers in the Hispanic sample.
d. Absolute value of t-statistic in parentheses (* significant at 10%; ** significant at 5%; *** significant at 1% level, in a two-tailed test).

Table 5. Heterogeneous Nature of Cohabitation

Fixed Effects Estimates of NLSY Wage Regression

	White Men		Black Men		Hispanic Men	
	Dummy var. spec	Duration Spec	Dummy var. spec	Duration Spec	Dummy var. spec	Duration Spec
	(1)	(2)	(3)	(4)	(5)	(6)
Married	0.075	0.048	0.044	0.026	0.019	0.010
	(4.04)***	(2.30)**	(1.95)*	(1.03)	(0.59)	(0.28)
Divorced or separated [c]	0.029	0.039	0.016	0.002	0.021	0.022
	(1.13)	(1.41)	(0.50)	(0.05)	(0.49)	(0.46)
Cohabit (ends in Marriage)	0.091	0.087	0.051	0.027	0.002	0.021
	(2.99)***	(2.56)**	(1.21)	(0.61)	(0.04)	(0.31)
Cohabit (ends in Separation)	0.003	-0.007	0.035	0.003	0.062	0.066
	(0.11)	(0.19)	(1.30)	(0.09)	(1.27)	(1.29)
Cohabit (truncated)	0.042	0.057	-0.021	-0.026	-0.088	-0.059
	(0.91)	(1.17)	(0.51)	(0.63)	(1.34)	(0.87)
Yrs in current marriage		0.017		0.013		0.008
		(4.04)***		(2.33)**		(1.40)
Yrs in current marriage, squared		-0.001		-0.001		-0.000
		(3.65)***		(2.04)**		(0.61)
Yrs div. or sep		0.003		-0.012		-0.009
		(0.47)		(1.10)		(0.73)
Yrs div or sep, squared		-0.000		0.001		0.000
		(0.39)		(1.57)		(0.25)
Yrs cohabit		0.010		0.014		-0.003
		(0.90)		(1.83)*		(0.59)
Yrs cohabit, squared		-0.001		-0.001		0.000
		(1.11)		(0.98)		(0.08)
Yrs in previous marriage		0.004		0.022		0.014
		(0.62)		(3.01)***		(1.63)
Yrs in previous marriage, squared		-0.000		-0.001		-0.000
		(0.38)		(2.42)**		(1.00)
Yrs cohabited w/ wife		0.002		0.015		-0.014
		(0.27)		(1.47)		(1.03)
Yrs cohabited w/ wife, squared		0.000		-0.002		0.002
		(0.27)		(1.66)*		(1.15)
Observations	9930	9930	6938	6938	4382	4382
Number of id	1465	1465	1130	1130	682	682
R-squared	0.25	0.25	0.20	0.20	0.19	0.19

Note: a. Absolute value of t-statistic in parentheses (* significant at 10%; ** significant at 5%; *** significant at 1% level, in a two-tailed test).

Table 6. Tenure, Weekly Hours Worked, and Training

Fixed Effects Estimates

	White			Black			Hispanic		
	Tenure	Weekly Hours Worked	Training	Tenure	Weekly Hours Worked	Training	Tenure	Weekly Hours Worked	Training
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Wife Less HS	-0.191	1.169	-0.027	0.486	0.623	0.000	-0.095	-0.298	-0.000
	(0.39)	(0.61)	(0.51)	(1.12)	(0.30)	(0.01)	(0.21)	(0.16)	(0.01)
Wife HS grad	0.313	-1.623	0.033	0.437	1.105	-0.026	0.372	-0.056	0.017
	(1.09)	(1.45)	(1.06)	(1.47)	(0.78)	(0.83)	(1.14)	(0.04)	(0.56)
Wife Somecoll	0.506	-3.087	0.047	1.032	0.889	-0.028	0.547	1.024	0.061
	(1.62)	(2.53)*	(1.41)	(3.13)**	(0.57)	(0.78)	(1.38)	(0.64)	(1.61)
Wife Collgrad	0.265	-3.510	0.027	0.841	0.601	-0.066	0.113	0.175	-0.070
	(0.83)	(2.82)**	(0.79)	(2.53)*	(0.38)	(1.84)	(0.23)	(0.09)	(1.47)
Partner Less HS	0.542	-2.774	0.044	0.401	-0.705	-0.000	0.026	5.002	-0.031
	(1.15)	(1.50)	(0.86)	(1.20)	(0.44)	(0.01)	(0.06)	(2.54)*	(0.70)
Partner HS grad	-0.047	-0.281	-0.062	-0.310	0.413	-0.000	0.010	0.528	0.026
	(0.11)	(0.17)	(1.34)	(1.02)	(0.29)	(0.01)	(0.02)	(0.27)	(0.59)
Partner Somecoll	-0.508	0.122	-0.066	0.001	-2.998	0.095	-1.657	-2.991	0.003
	(0.95)	(0.06)	(1.14)	(0.00)	(1.67)	(2.34)*	(2.55)*	(1.12)	(0.05)
Partner Collgrad	-0.624	-0.170	-0.186	-0.797	1.113	0.058	-1.309	-5.852	-0.014
	(1.06)	(0.07)	(2.91)**	(1.85)	(0.55)	(1.26)	(1.35)	(1.50)	(0.15)
Observations	10025	9928	10025	7018	6767	7018	4440	4344	4440
Number of id	1465	1464	1465	1130	1126	1130	682	682	682
R-squared	0.19	0.02	0.07	0.16	0.03	0.06	0.14	0.04	0.05

Note: a. Each of the estimated equation also included age, age-squared, dummy for SMSA, local unemployment rate, 3 regional dummy variables, and 7 year dummy variables.
b. Absolute value of t statistics in parentheses; * significant at 5%; ** significant at 1%.

Chapter 2

The White Picket Fence Dream:

Effects of Assets on the Choice of Family Union

I. Introduction

There have been dramatic changes in the family formation behavior of young American men and women over the last four decades. During this period, the prevalence of cohabitation has been increasing sharply while age at first marriage has also been rising, and the percentage of marriages preceded by cohabitation has been growing substantially (Fields and Casper, 2001; Casper and Cohen, 2000; Bumpass and Sweet 1989). These demographic changes have prompted serious concerns from researchers as well as policy makers about a retreat from the traditional pattern of family formation. However, both quantitative and qualitative research (e.g., Edin, 2000; Tucker, 2000; Thornton and Young-DeMarco, 2001; Gibson, Edin and McLanahan, 2003, Cherlin, 2004) indicate that among the unmarried population there is not a large-scale lack of respect for marriage and the traditional ways of family. Young Americans do place a high value on marriage and consider it as part of their future. In fact, Gibson, Edin and McLanahan (2003) argue that at least for some young unwed parents, high marital expectations may be precluding them from marrying. In a qualitative analysis of 75 unmarried young couples in the Fragile Families study, they find that marriage signals the "arrival" of the couple, both financially and emotionally. Because marriage is valued so highly, it is perceived as a family status to be chosen after certain economic and relational preconditions are fulfilled – after they have achieved the so-called "white picket fence dream".[11] Similar observations are also reported in Edin (2000), drawing

[11] For recent cohorts of men and women in the U.S. (such as the NLSY79 cohort), such postponement of marriage apparently conforms with the "Easterlin hypothesis", which relates men's

on qualitative interviews with 292 low-income single mothers in three U.S. cities. While these are interesting observations, there has not been any attempt to substantiate the qualitative evidence using quantitative methods in a large scale dataset. If people postpone their marriage until they can achieve the different level of living standard that they associate with marriage – a house, surplus income etc. – one may expect to detect a direct relationship between the individual's housing and financial assets, and his/her transition into marriage. In this paper, we examine this intriguing issue using data from the National Longitudinal Survey of Youth 1979. In addition to considering transition into marriage, we also analyze the effect of asset ownership on transition into cohabitation.

Since we are interested in identifying whether asset ownership status explains the time until the formation of a family union, we utilize the proportional hazard model – a natural estimation technique for analyzing time-to-event data. The time-to-event analysis would ensure that asset ownership status prior to a family union transition is sequentially exogenous to such a transition decision. However, the individual's intention to form a family next period or period after may influence his/her asset accumulation behavior in the current period. As a result, unobserved individual heterogeneity as well as the shocks that affect the choice of family status could be correlated with asset ownership status, and the proportional hazard model estimates would be potentially inconsistent. To address the potential endogeneity of assets we implement instrumental variables estimation, where the set of excluded instruments contains the interaction of monthly averages of the 30-year fixed-rate mortgages, the 1-year adjustable-rate mortgages, the federal funds rate, and NASDAQ stock price index, with individual's age, ethnicity and region of residence.

In our attempt to understand the role of assets in family union choice, we consider cohabitation and marriage comparably. Thus, in addition to bridging the void in the literature on identifying the effects of assets on family union, this paper would enable us to better understand the differences and similarities between marriage and cohabitation. Furthermore, apart from providing insights into the mechanism underlying the individual's

---

marriage and fertility decisions with his earnings relative to his 'desired style of life', the latter being determined by his adolescent experience (Easterlin, 1973). This hypothesis is not pursued further in the present study.

decision to form families, the qualitative evidence on assets as determinants of family union transitions are also suggestive of the potential influence that asset building policies might have on family union behavior.[12] Particularly, public policy interventions for increasing home ownership, and enhancing savings among the low-income group through individual development accounts, would be expected to affect the family union choices of the participants. The results available from the current study will enable us to make more knowledgeable estimates about whether asset-building policies could potentially influence family union behavior.

The following section provides a conceptual discussion of the determinants of family union decisions, and a brief review of the relevant empirical literature. Section III describes the estimation procedures, and section IV delineates the NLSY data and summary statistics. Empirical results are presented in Section V. Summary and conclusions follow in section VI.

II. Economic Resources and Family Union Formation: Theoretical Perspectives and Empirical Research

II. 1. Economic resources and marriage

Economic analysis of household formation, built on the foundation of Becker's (1973, 1974, 1991) seminal work, emphasizes the effects of economic resources on the likelihood of marriage. From a microeconomic perspective, the effect of a change in the individual's economic prospects, such as a rise in the wage rate, on the timing of marriage can be analyzed as the effect of a change in the wage rate on the allocation of time among family, schooling, and market work. From this point of view, a rise in the wage rate tends to increase market work, and in turn has a negative substitution effect on the demand for family or schooling (Becker, 1973). However, rising wages also has an income effect which makes marriage and family more affordable and possibly increases the rate of return to schooling. Hence, the marriage effect of young men's or women's improved labor market conditions is an empirical question.

[12] Stern (2001), Sherraden (2001), and Seidman (2001) discuss asset-building policies in the US.

However, if gains from marriage are considered to ensue only from intra-household specialization (as in Becker, 1973), improved labor market conditions for men would unambiguously increase the likelihood of marriage, while similar improvement for women would lower the likelihood of marriage. Particularly for women, if gains from marriage are considered to come not only from specialization but also from joint consumption economies within the household, a priori, the effect of improved women's labor market conditions on the likelihood of marriage will, once again, be ambiguous (Lam, 1988).

One of the most widely noticed discussions of a causal link between economic resources and the postponement of marriage among the young Americans, particularly among the low-income black population, has been provided by Wilson (1987, 1996). Working his way through the complexity of jointly determined variables, he focused on decreased employment opportunities for men as the exogenous "prime mover" in reducing the marriage rates, and also affecting other related outcomes. One way in which the Wilson hypothesis can be put in the context of the economic theory of marriage is to consider marriage as the matching outcome from a search process in the marriage market where women only consider marriage with men who have demonstrated a minimum ability to perform in the labor market (Wood, 1995).

The nature of the effects of changing economic factors on the timing of marriage is not always clear a priori. But the theoretical discussions do point to three economic factors as being potentially important: employment and earnings of men, employment and earnings of women, and welfare benefits (Ellwood and Crane, 1990). In addition to these three factors, other measures of economic resources that feature prominently in the empirical literature include men's and women's educational attainment, work experience, and parental resources. Before we discuss the empirical studies analyzing the family union transitions, it would be appropriate to note that a persistent limitation of many of these studies is that they are unable to address the issue of individual heterogeneity which can directly influence the likelihood of a transition into marriage or cohabitation, while at the same time being correlated with the individual's economic potential. As a result, most of the findings are indicative of an associative relationship between economic resources and family union transitions, rather than a causal one. Only a handful of more careful studies have tried to overcome this challenge either using innovative estimation methods or utilizing long-term shocks to the local economy.

Empirical results indicate that men's labor market opportunities are associated with significantly higher rates of marriage, although their quantitative effects may be small (see, for example, Xie et. al., 2003; Oppenheimer, Kalmijn and Lim, 1997; Willis and Michael, 1994; Schultz, 1994; Mare and Winship, 1991; Ellwood and Crane, 1990; MacDonald and Rindfuss, 1981). A small number of studies have been attentive in addressing the issue of endogeneity of men's labor market opportunities with respect their choice of family status. Olsen and Farkas (1990) utilize a waiting-time regression model with individual fixed effects to evaluate the effect of a government program that guaranteed employment opportunities to disadvantaged youth, on family union and fertility. Using data from the Youth Incentive Entitlement Pilot Project, they find that employment opportunity encourages the formation of consensual unions.

Black, McKinnish and Sanders (2003) study the shocks to the coal and steel industries to measure the effect of long-term changes in demand for low-skilled workers on welfare expenditures and family structures. Using county level panel data for the period 1969-93, they find that expenditures on AFDC responded substantially to the long-term cyclical movements in the steel and coal industries. More pertinent to our current focus is the finding that the expansion of high-wage jobs for low-skilled men increased marriage rates, and reduced the incidence of female-headed households, thereby reducing the number of families at risk for receiving welfare benefits.

The findings in these more careful studies are supportive of the Wilson hypothesis that men's labor market opportunities are fundamental in determining the marriage rates, particularly in the low-income group. Wood (1995), on the other hand, uses SMSA level aggregated data from US Censuses in 1970 and 1980. His fixed-effects estimates suggest that up to a tenth of the decline in black marriage rates during the 1970s can be explained by declines in the current economic status of young black men. However, when he instruments for the endogeneity of marriage and 'marriageability', with changes in the number of jobs in the SMSA in ten major industry categories as the set of excluded instruments, the estimated effects are substantially reduced. Wood (1995) concludes that shrinking pool of high earning young black men explains little of the decline in black marriage, thereby contradicts the Wilson hypothesis.

Studies focusing on women's labor market prospects have found mixed empirical evidence. Some of these studies have found that better economic prospects for women are associated with declines in marriage (e.g., Aassve, 2003; Blau, Kahn and Waldfogel, 2000; McLanahan and Casper, 1995; Willis and Michael, 1994; Schultz, 1994; Mare and Winship, 1991), while others find that the estimated relationship between indicators of women's economic status and incidence of marriage is either positive or insignificant (e.g., Xie et. al. 2003; Oppenheimer and Lew, 1995; Mare and Winship, 1991).

One aspect of family union transitions that has been ignored until recently is the transition from cohabitation into marriage. Recent empirical endeavors in this regard find mixed evidence. Some studies have shown positive association of men's earnings and the transition from cohabitation to marriage (Carlson, McLanahan and England, 2004; Brown, 2000; Sanchez, Manning and Smock, 1998; Smock and Manning, 1997), while others have reported a significant negative association between higher men's earnings and cohabiters' decision to marry (Sassler and McNally, 2003, Wu and Pollard, 2000). Previous research mostly tended to indicate that there is no significant effect of women's economic opportunities on transition from cohabitation to marriage (Sassler and McNally, 2003; Sassler and Schoen, 1999; Clarkberg, 1999), although a recent study showed that women's education encourage transition to marriage among young unwed mothers (Carlson, McLanahan and England, 2004).

II.2. Economic resources and cohabitation

With a handful of recent exceptions, the existing quantitative literature on the role of economic resources in family formation has focused exclusively on marriage, ignoring cohabitation. The studies that consider cohabitation include Xie et. al. (2003), Clarkberg (1999), Smock and Manning (1997), Raley (1996), and Thornton, Axinn and Teachman (1995). Two studies have indicated that improvement in men's economic opportunities encourage the formation of cohabiting unions (Clarkberg, 1999; Smock and Manning, 1997), another study found that men's school enrollment deters entrance into cohabitation (Thornton, Axinn and Teachman, 1995), while still others have reported no significant effect of several measures of men's economic potential (Xie et. al., 2003) on the rate of transition into cohabitation. With regard to women's economic potential, previous studies indicate that enhancing women's economic potential discourages the formation of cohabiting unions (Thornton, Axinn and Teachman, 1995), while others found that women's economic status has no significant effect on transition into cohabitation (Xie et. al., 2003). Kravdal (1999) relates the growth in cohabitation in Norway with the issues of 'affordability' of marriage. He finds that women's cumulated income has a positive association with the likelihood of cohabitation among women with children, but not for women without children. The author interpreted these results as weak evidence on marriage requiring a stronger economic underpinning than cohabitation. However, since sufficient attention has not been paid to the issue of selection, one has to be very cautious in interpreting such findings any more than reflections of positive correlations.

II.3. Asset ownership and family union formation

While income is certainly critical, wealth and assets are also important complementary measures of an individual's command over economic resources. The individual's assets give us an estimate of their economic readiness to marry in relation to their ideational value of marriage. Individuals with higher exogenous endowments are more likely to marry because they have more to share and can provide greater access to credit and insurance (Lam, 1988). On the other hand, higher exogenous endowments can potentially reduce the marital surplus by way of requiring less specialization in the household, and thereby dissuade the individual from marriage. Therefore, a priori, the effect of assets on the choice of family union is ambiguous.

More realistically, assets are not exogenous and they reflect accumulated past income, and savings behavior. The economic model of the determinants of marriage considers the concept of *potential* wage rates, instead of *actual* or *realized* wage rates. *Realized* wages is as much the result of the marriage decision as its cause. Since assets (net of inheritance) are accumulated savings from *realized* earnings, they are likely to be directly affected by the marriage decisions. And even inheritance can be influenced by the individual's choice regarding family union. There is substantial empirical research showing how family composition affects household wealth and savings (e.g., Aizcorbe, Kennickell and Moore 2003; Lupton and Smith 2003; Wolff, 2001; Mulder and Wagner, 2001; Browning and Lusardi 1996; Lusardi, Gossa and Krupka 2001; Avery and Kennickell 1991). Most of these empirical studies on savings are descriptive, and they generally identify that married couples have the highest levels of wealth and lone parents the lowest with singles in between (but with quite low levels of wealth). Taken together, these studies suggest that addressing the endogeneity of assets is the primary challenge in analyzing their effects on the choice of family form.

III. Empirical Methodology

The primary challenge in analyzing the effect of assets on the choice of family form is to address the issue of endogeneity of assets with respect to family status. We undertake a two-pronged approach to address this issue. First, we utilize a time-to-event analysis approach by using proportional hazard model. Second, we implement instrumental variables estimation.

III.1. Time-to-event analysis

The central question we are examining is whether the individual's asset ownership status explains his/her family union transition. The question could be put forward alternatively as whether asset ownership can explain the time elapsed until a family union transition occurs. A natural way to empirically estimate such effects is to apply a time-to-event analysis approach. The benefit of using a time-to-event analysis in our context is that it ensures 'sequential exogeneity' [13] of assets with regard to family status. We are looking at the effect of asset ownership prior to the event of a family union on the probability of a union in

[13] For a discussion on sequential exogeneity, see Wooldridge, 2002, chapter 11.

the next period. Hence, asset ownership is not sequentially dependent on family transition decision.

We consider five sets of family union transitions: a) non-partnered to first partnered union (marriage or cohabitation); b) non-partnered to first cohabitation; c) non-partnered to first marriage; d) cohabitation to first marriage; and e) unmarried (non-partnered or cohabiting) to first marriage. We consider only transitions into the first marriage and first cohabitation in order to keep our analysis simple, since subsequent family unionization is confounded by the choices on dissolving the previous family union. Since the individual's asset ownership status varies over time, for our purposes we utilized a Cox proportional hazard model with time-varying covariates (see Lancaster, 1990 for details) to analyze these transitions. In this model, the instantaneous hazard rate of transitions to family union is specified for individual *i*, *t* years until the family transition occurs, as:

$$h[t, x_i(t)] = h_0(t)\exp[\beta_x x_i(t)]$$

The baseline hazard, $h_0(t)$, is a nonparametric, time-varying function; $x_i(t)$ is a vector of regressors that includes time-varying asset ownership indicators; and β_x is the vector of coefficients to be estimated. We used the maximum-likelihood estimation procedure available in *Stata* to implement the model (Cleves, Gould and Gutierrez, 2004).

III.2. Instrumental variables estimation

Although the time-to-event analysis ensures sequential exogeneity of asset ownership with respect to family status, the individual's intension to form a family in the future may influence his/her asset accumulation behavior in the current period. Therefore, unobserved individual heterogeneity as well the shocks that affect family status could be correlated with asset ownership status. In this sense, asset ownership status may not be strictly exogenous to the family transition decision, and hence, the proportional hazard estimates are potentially inconsistent. In other words, the time-to-event analysis may not provide us with the true effects of asset ownership on family union transitions. Note that the true effect of asset ownership is the effect which would result were it possible to randomly assign asset ownership to a sample of non-partnered men and women. This true effect may be smaller than the effect we estimate by comparing the hazard rates of family union transition for men and women with assets to the rates of transition for men and women without assets, precisely because of the type of endogeneity that is suggested above. Since there is no established instrumental variables framework for hazard models, we implement instrumental variables (IV) probit estimation in a discrete-time analogue of the (continuous time) proportional hazard model to deal with the potential endogeneity of assets.[14]

For the IV estimation, instead of having a standard pooled cross-sectional limited dependent variable, we define the dependent variable as a dichotomous indicator of whether a family union transition occurs in the next period.[15] The advantage of constructing the dependent variable in this fashion is that we are able to retain the sequential exogeneity of the asset ownership status in the time-to-event analysis while we address the concern about strict exogeneity of asset ownership.

IV. NLSY79 Data and Summary Statistics

The National Longitudinal Survey of Youth 1979 (NLSY79) is a nationally representative US sample of young men and women who were 14-22 years old when they were first interviewed (CHRR, 2001). The respondents were interviewed annually until 1994, and biennially since then. Data from the first through the 19th (2000) round are used for this paper. We have used data from the earliest round to determine the respondent's family life history. Detailed information on wealth and assets are, however, available only since the 7th round (i.e., 1985).[16]

---

[14] The other way to do it may be to estimate a 'waiting-time regression model' (Olsen and Wolpin, 1983) by linearizing the probability density function of time to failure, and estimating using linear methods. However, Olsen and Wolpin (1983) requires imposition of an exogenously chosen upper limit for remaining in the non-partnered status, which is deemed to be very restrictive for our analysis.

[15] To use definitions from time-to-event analysis, we use the "failure indicator" in the hazard model as the dependent variable in our discrete time analysis.

[16] Due to budgetary restrictions, wealth questions were not administered in 1991 and 2002 rounds of NLSY79.

We stratify the data by gender, and all our analyses are conducted separately for men and women. We observe the family life transitions of men and women in our sample during the period 1985 to 2000. Table 7 presents some summary statistics on the unmarried spells considered here. As the table shows, the median duration in non-partnered spell for both men and women in our sample is more than 11 years. This is not surprising given that the beginning of this spell is either the date of their first interview in NLSY or their 14 birthday, whichever is later. The median duration in cohabitation is about 2.7 years. Table 8 provides the number of events we observe in our data for each category of transition, along with the median duration prior to any transition. During the period under analysis, there are 1525 transitions into marriage among women; 659 of these transitions are from cohabitation. For men, there are 1807 transitions into marriage of which 683 are from cohabitation. We also have 1304 transitions for women from a non-partnered (i.e., never-married non-cohabiting) status to a family union – either in marriage or in cohabitation; of which 422 are into cohabitation. Among men, we observe 1877 transitions from a non-partnered status to a family union, and 722 of these transitions are into cohabitation.

In Table 9, we present snapshots on the different characteristics of the respondents in our sample. The table is intended to provide a glimpse into the nature of the sample we have for our analysis. The table shows that the proportions that are married increased over time for both men and women, although the fraction of men married converged to the fraction of women married only in the later years. Summary statistics are presented for men and women in three different family statuses: non-partnered (never-married and non-cohabiting), cohabiting and married in 1985.

As we examine the effect of asset ownership on family union transitions in this study, we take three types of assets into consideration[17]: home ownership; liquid monetary assets as indicated by the availability of funds in savings account, certificates of deposit, money market instruments and IRA-Keoghs; and financial investments in stocks, bonds, and mutual funds. In our empirical analysis we include dichotomous indicators of ownership of

---

[17] These ownership data are available from the "Asset" section of the NLSY79, and the relevant questions ask about the ownership status of the respondent and their spouse/partner with respect to the particular categories of assets.

these three types of assets. While data on home ownership and liquid monetary assets are available since 1985, stocks-bonds-mutual funds data are available only from 1988.

Table 9 reveals that both married men and married women are significantly more likely to be a home owner. On average, married men and women are also more likely to own liquid monetary assets, as well as investments in stocks, bonds and mutual funds. This is all the more clear from Figures 1 through 6 showing the average asset ownership status of non-partnered, cohabiting and married men and women at different age. Although the fraction who are homeowners increases with age for men and women in any family status, at every age those who are married are more likely to be homeowners than those who are not married. Cohabiting men are marginally more likely to be homeowners than non-partnered men, and the same is true for young cohabiting women. A much larger fraction of cohabiting women in their mid-30s are homeowners compared to the non-partnered in that age group. In terms ownership of monetary assets, married men in their early 20's appear to be similar to cohabiting men. However, married men who are in their late 20's or older, are more likely to have some monetary assets compared to men in the other two family statuses. Married women, on the other hand, are more likely to have some monetary assets at about any age compared to the unmarried women. The proportion owning financial investments is quite low for both men and women in any of the three family statuses.[18] However, even at that low level, married men and women at any age are more likely to own some financial investments than their unmarried counterparts. Interestingly, cohabiting men and women are either less (in their 20's) or equally (in their 30's) likely to own monetary assets as well as financial investment compared to non-partnered men and women.

The other characteristics that are considered in our analysis and are summarized in Table 9 include age, race and ethnicity, own education, income (wage and business income) in the past calendar year, welfare recipiency in the past calendar year (includes receipt of cash assistance from AFDC or TANF, supplemental security income, food stamps, housing support, or any other benefit), religion, father's and mother's education, region of residence, whether or not the state of residence recognizes common law marriage, and local

[18] In figures 5 and 6, the outlier percentages in the highest age group reflect small sample in that age group.

unemployment rate. Men and women in our sample are between 20 to 28 years old in 1985, the year since which we follow their family union transitions. On average, married men and women are older than their counterparts in the other two family statuses. Married men and women are also likely to be at least high school graduates than others. Table 9 also indicates that while women in general are more likely to be welfare recipients than men, larger fraction of unmarried women received public assistance than their married counterparts.

V. Results

V. 1. Results from time-to-event analysis

Estimates from the Cox proportional hazard model with the asset ownership variables as time-varying covariates are presented in Table 10. As already noted, in analyzing the relationship between family union transitions and asset ownership, we are considering the ownership status of housing assets, of liquid financial assets, and of stocks, bonds and mutual funds. Since data on stocks and bonds ownership is available from a later period (from 1988, instead of 1985), in our analysis we estimate two sets of specifications – one that excludes stocks-bonds-mutual funds ownership indicator (Model 1), and another that includes it (Model 2). Table 10 presents the estimated hazard ratios only for the covariates of interest, and estimates for the complete specification are available upon request. We discuss the estimated results for men and women separately.

Results in section (a) in Table 10 show that home ownership has no statistically significant effect on women's transition from never-married status into a partnered relationship (either in marriage or in cohabitation). Money in the savings account, on the other hand, is positively associated with women's transition into a family union. However, as the ownership of stocks and bonds is included in the specification, thereby reducing the period of analysis and the sample size, the estimated hazard rates on monetary assets no longer remain statistically significant. In the final specification, reported as Model 3 in Table 10, measures of wage and business income in the previous calendar year, and an indicator of the individual's public assistance (AFDC/TANF, food stamps, unemployment insurance benefits, supplemental security income etc.) recipiency in the previous calendar year is added to the specification in Model 1.[19] While income and welfare recipiency are added to remove any reservation regarding the estimated effects of asset ownership, their potential endogeneity with respect to family union transitions convinced us for not including them in our initial specifications. As it appears, inclusion of these two variables does not change the estimated effects of asset ownership on transition to a family union in any substantive way. In addition, welfare recipiency tends to be negatively correlated with transition into a family union, a result that resonates with a large existing literature (see Moffitt, 1998 for a recent review).

Next, we perform a competing-risks analysis that allows for two ways to exit the non-partnered status: form a family in cohabitation, or in marriage. Estimated hazard ratios for transitions from non-partnered to cohabitation are reported in section (b) in Table 10. For women, while model 1 shows no significant effect, model 2 indicates that having liquid assets is negatively associated with forming a cohabiting relationship. Results from section (c) shows that ownership of liquid assets as well as stocks and bonds is positively associated with the rate of marriage for non-partnered women. As it appears, the positive association of liquid asset ownership and women's rate of forming any family union (in section a) is primarily driven by the positive association between ownership of these assets and transition into marriage. This result is further strengthened by the estimates in section (d) which indicate that women who have access to liquid assets are more likely to marry their cohabiting partners. Result in section (d) further show that home ownership has no significant relationship with the rate of marriage among cohabiting women. While the results in this section for liquid assets conform to the qualitative evidence that women would chose to marry their cohabiting partners when there is surplus income (e.g., Gibson, Edin and McLanahan, 2003; Cherlin, 2004), the insignificant relationship between home ownership and marital transition for cohabiting women does not correspond to a so-called 'white picket fence' explanation. As we consider women's transition into marriage from either a never-

---

[19] A specification that added income and welfare recipiency variables to Model 2 is also estimated, and the results (not reported here, but available upon request) show that the estimated coefficients of interest are not significantly different from the initial specification.

married or a cohabiting status (section e, Table 10), surprisingly we find that along with liquid assets, home ownership has a statistically significant positive correlation with such transitions.

Overall, results from Cox proportional hazard model estimates suggest that liquid assets are positively associated with women's rate of transition into marriage. For home ownership, there is weak evidence of such positive relationship. In the case of transition into cohabitation, asset ownership is negatively associated, if not uncorrelated, with women's choice of cohabitation.

For men, the overall evidence indicates that asset ownership is positively associated with rate of transition to marriage, but not to cohabitation. Specifically, the rate of marital transition by non-partnered men has a significant positive relationship with both home and liquid asset ownership, but no significant relationship exist for men's transition from non-partnered status to cohabitation (see sections a, b, c, and e for men in Table 10). More interestingly, cohabiting men's rate of transition to marriage has no significant relationship with home ownership, but there is positive significant association with ownership liquid assets and financial investments. These results give the impression that while ownership of all types of assets is positively correlated with the never-married men's decision to marry; only liquid assets are significantly correlated with cohabiting men's choice to marry. Moreover, men's rate of transition to cohabitation is not at all related to his asset ownership status.

It would be appropriate to note that individual's race and ethnicity indicators have been included in all the specifications reported, along with the other control variables. Overall, for all the transitions analyzed here, compared to non-black non-Hispanic women, black women are less likely to be in any type family union – in marriage or in cohabitation. Black men, on the other hand, are significantly more likely to transit to a cohabiting relationship, and less likely to transit into marriage in comparison with their non-black non-Hispanic counterparts. Our estimates for Hispanic women are similar to those for Black women, but somewhat weaker. For Hispanic men, we have some evidence that unlike black men, they are less likely either to marry or to cohabit than non-black non-Hispanic men.

The empirical estimates discussed so far provide some interesting new evidence. Taken together they show that home ownership as well as access to liquid assets is positively associated with rates of marital transition from a non-partnered status, particularly for men. However, home ownership is not associated with rates of marital transition among cohabiters in the sample, although access to liquid monetary assets is. In addition, both for men and women, asset ownership is not associated with the rate of transition into cohabitation. These results only partially agree with the previously discussed qualitative evidence suggesting a stronger relationship, particularly between home ownership and rates of transition from cohabitation to marriage. Moreover, the results for the transition from cohabitation to marriage might be considered as suggestive evidence that both men and women would accumulate monetary assets while cohabiting, and buy a house when their commitment to a relationship is sealed by the decision to marry. This only reinforces the apprehension about the endogeneity of asset accumulation that has been noted earlier.

V. 2. Results from instrumental variables estimation

To address the potential endogeneity of assets we implement instrumental variables (IV) estimation. To operationalize the procedure, we convert the structure of analysis from a continuous time to a discrete time hazard model with the dependent variable being a dichotomous indicator of a family union transition in the next period. As we treat the asset ownership status as endogenous to family union transitions, the set of excluded instruments is constructed as follows. We use the interaction of monthly averages of the 30-year fixed-rate mortgages, the 1-year adjustable-rate mortgages,[20] the federal funds rate, and NASDAQ stock price index, with individual's age, ethnicity and region of residence as the set of excluded instruments.[21] The individual's homeownership decision is expected to be influenced by the mortgage rates as they play a dual role in the housing market: on the one hand, mortgage rates show the time-value-of money; and on the other hand, they have key

---

[20] The monthly average mortgage rates are collected from the Freddie Mac Survey of Commitment Points and Rates.

[21] Housing price could not be used as an instrument since it directly affects the decision to form independent households (e.g., see, Borsch-Supan, 1986, Haurin et. al, 1993).

roles in determining housing prices. The stock price index and the federal funds rate are proxies for the return to savings and investment during the sample period, and are therefore expected to influence the individual's stock of assets at any point in time. Interacting the mortgage rates, stock price index and the federal funds rate with the individual characteristics provides us with the individual level variation required for identification of the estimating equation.

Test statistics on the joint-significance of the excluded instruments in the first-stage regressions are provided in Table 11. Although the F-statistics for ownership of stocks and bonds is quite small, the values of the F-statistic on the other two asset ownership indicators are sufficiently high to remove the greater portion, if not all, of the potential bias in an analysis conducted without regard to the endogeneity of assets (Bound et. al. 1995, Table A-1; Hahn and Hausman, 2002). The validity of these instruments, particularly for the specifications that includes only home and liquid asset ownership, is also underlined by the Hansen-Sagran J statistic for over-identification tests reported in Table 12 along with the IV-probit estimates.

Table 12 reports both probit and IV-probit estimates for specifications described earlier as Model 1 and Model 2. The sign and significance of the estimated coefficients on asset indicator variables in the probit model are highly comparable to those in the Cox proportional hazard model for every type of family union transitions considered, and for both men and women.

IV-probit estimates for women show that addressing the potential endogeneity of asset ownership removes the positive association between asset ownership and family union transitions. On the contrary, access to liquid assets has a negative influence on the rates of transition to marriage for non-partnered women (sections a, c and e in Table 12). Inclusion of income and welfare recipiency in the specification (not reported in the table) does not change these results in any important way. Section (b) in Table 12 indicates that asset ownership has no statistically significant influence on women's rate of transition to cohabitation as we account for the endogeneity of assets. IV probit estimates in section (d) of Table 12 show that home ownership reduces women's likelihood of marrying their cohabiting partner, although such negative effect does not remain statistically significant when we include ownership of stocks and bonds in our specification.

For men, IV probit estimates show that liquid asset ownership reduces the rate of transition to cohabitation, although inclusion of stocks and bonds in the specification (and thereby, reducing the sample size as well) indicates such effects are not significantly different from zero. For men's transitions into marriage, from non-partnered, or cohabiting, or any unmarried status, the IV-probit coefficients on the asset ownership indicators are not statistically significant, thus indicating that neither homeownership nor liquid assets has any significant effect on men's marriage decisions. However, ownership of stocks and bonds has significant negative influence on the probability of a transition from cohabitation to marriage.

All together, the IV probit estimates either remove the statistical significance of the association between asset ownership and family union transitions, or indicate effects that are in the opposite direction to those derived from the time-to-event analysis. However, it has been suggested in the literature that the influence of economic factors in family life transitions might be different for different education groups (e.g., Moffitt, 2000). To examine whether our results also reflect a similar pattern, we conduct the instrumental variables analysis by stratifying our sample into two groups: those who have never attended college (i.e., at most high-school graduate), and those who have some college or more education (i.e., more than high school education). The F statistics for the joint significance of the instruments in the first stage estimation indicate that as we stratify the sample, the instruments perform satisfactorily only in the case of transitions from unmarried to married status. In Table 13 we report the IV-probit results for this particular transition, for men and women disaggregated by education groups. A comparison of these estimates with those reported in section (e) in Table 12 suggests that the estimated effects of asset ownership on marital transitions do not vary significantly between the two education-groups, for both men and women. While these auxiliary results are available only for the transitions from unmarried to married status, they provide enhanced confidence in our IV estimates for all the other transitions identified across all education levels.

VI. Conclusion

Motivated by a recent set of findings by demographers, the paper presents two broad sets of evidence on the relationship between asset ownership and the family union transition decisions by men and women. The first set of findings, from Cox proportional hazard model with time varying covariates, reveal that home ownership as well as access to liquid assets is positively associated with rates of marital transition from a non-partnered status, particularly so for men. However, home ownership is not associated with rates of marital transition among cohabiting men and women in the sample, although access to liquid monetary assets is positively associated with rates of cohabiting men and women marrying their partners.

The second set of findings stem from instrumental variables probit estimation, implemented to remove the potential bias in the hazard model estimates. The bias is anticipated due to the endogeneity of assets with respect to the family union decisions. The IV probit estimates either remove the statistical significance of the association between asset ownership and family union transitions, or indicate effects that are in the opposite direction to those derived from the estimated hazard model. As with any instrumental variables estimation, the strength of the set of identifying variables in playing the role of instruments is crucial. While we would not claim that these results are definitive, we do believe that they are suggestive. In relating asset ownership to family union behavior, researchers need to be aware that it is likely that common preferences or opportunities underlie both the decisions to accumulate asset and the decision to form a family by marriage or by cohabitation.

Results from the time-to-event analysis indicate that at least as a behavioral regularity we observe a positive relationship between asset ownership and marital transitions. The IV estimates, however, suggest that such behavioral regularity does not indicate a causal relationship. In other words, those who are inherently more likely to marry are the ones who would accumulate assets, and hence asset ownership does not cause their transition into marriage. Indeed, homeownership and monetary assets may reduce the marital surplus by reducing the level of specialization in the household, and thereby influence people to delay marriage, as suggested by the IV results. More importantly, these results have notable policy implications. If the IV estimates are reflective of the true effects of asset ownership, provision of housing subsidy or incentive to accumulate assets may not lead to any significant improvement in the rates of marriage.

Table 7.
Summary Statistics on Spells in Different Family Status Prior to Union Transitions
NLSY79 Women and Men; 1985-2000

Spells	No. of Persons	No. of Observations	Median Duration (in months)
Women			
Non-partnered spells	2056	11073	136.8
Cohabiting spells	1458	3668	32.4
Unmarried spells	3045	14652	116.5
Men			
Non-partnered spells	2875	14717	136.6
Cohabiting spells	1592	4076	34.0
Unmarried spells	3675	18665	130.3

Table 8.
Number of Different Family Union Transitions
NLSY79 Women and Men; 1985-2000

Transitions	No. of Transitions	Conditional Median Duration (months) [a]
Women		
Non-partnered to partnered	1304	25.6
Non-partnered to cohabitation	422	24.5
Non-partnered to marriage	882	26.1
Cohabiting to marriage	659	17.7
Unmarried to marriage	1525	22.1
Men		
Non-partnered to partnered	1877	25.4
Non-partnered to cohabitation	722	24.4
Non-partnered to marriage	1155	25.8
Cohabiting to marriage	683	18.2
Unmarried to marriage	1807	22.9

Note: a. Duration conditional on the fact that the individual made the transition.

Table 9: Summary of Key Variables by Family Status in 1985
Women and Men NLSY79

	1985 Status → / Variables ↓	Women (N=4535)			Men (N=4350)		
		Never Partnered	Cohabit	Married	Never Partnered	Cohabit	Married
Proportion in status [a]	Proportion in 1985	0.40	0.09	0.43	0.57	0.07	0.30
	Proportion in 1990	0.20	0.08	0.57	0.28	0.10	0.50
	Proportion in 2000	0.12	0.05	0.61	0.14	0.07	0.60
Assets	Own house	0.02	0.07	0.38	0.03	0.10	0.33
	Have money asset	0.57	0.49	0.68	0.54	0.46	0.66
	Own stocks & bonds [b]	0.09	0.04	0.14	0.11	0.05	0.16
Age	Age (years)	22.93	23.40	24.17	22.86	23.81	24.52
Ethnicity	Hispanic	0.15	0.14	0.18	0.15	0.23	0.18
	Black	0.38	0.22	0.15	0.31	0.30	0.15
Education	HS grad	0.37	0.47	0.50	0.41	0.46	0.48
	Some College	0.35	0.18	0.20	0.28	0.14	0.15
	College grad	0.16	0.07	0.10	0.12	0.05	0.12
Income	Annual income ($)	6347	6326	6256	8477	11202	14604
Welfare	Welfare recipiency	0.20	0.29	0.18	0.05	0.09	0.13
Religion	Protestant	0.07	0.08	0.08	0.08	0.04	0.08
	Baptist	0.28	0.23	0.19	0.22	0.23	0.17
	Catholic	0.25	0.25	0.26	0.24	0.28	0.26
	Other Christian	0.11	0.08	0.11	0.11	0.07	0.10
	Jew	0.01	0.00	0.00	0.01	0.00	0.00
Father education	HS grad	0.29	0.30	0.29	0.29	0.28	0.30
	Some College	0.08	0.07	0.07	0.09	0.07	0.07
	College grad	0.14	0.09	0.10	0.14	0.09	0.11
	Missing	0.13	0.18	0.10	0.13	0.14	0.11
Mother's education	HS grad	0.35	0.35	0.38	0.38	0.36	0.38
	Some College	0.10	0.09	0.07	0.09	0.07	0.08
	College grad	0.09	0.04	0.05	0.08	0.05	0.06
	Missing	0.05	0.05	0.05	0.07	0.10	0.06
Region of Residence	North Central	0.24	0.27	0.24	0.24	0.22	0.25
	South	0.40	0.33	0.41	0.35	0.29	0.39
	West	0.15	0.22	0.20	0.19	0.28	0.22
State	Common Law Marriage	0.34	0.28	0.35	0.33	0.32	0.37
Local	Unemployment rate	8.03	8.06	8.36	8.06	7.81	8.36

Note: a. The remaining sample include divorced or widowed, and those who are single after cohabiting.
b. For 1988, as data on stocks, bonds and mutual funds is available only since the 1988 round of NLSY79.

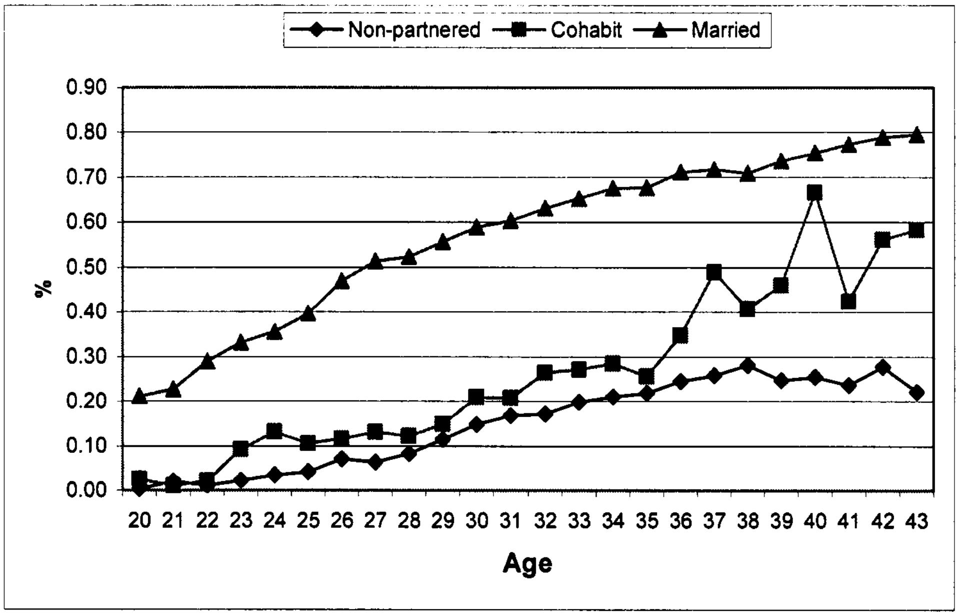

Figure 1. Percent of Women with Home Ownership by Family Status
1985-2000 (NLSY79)

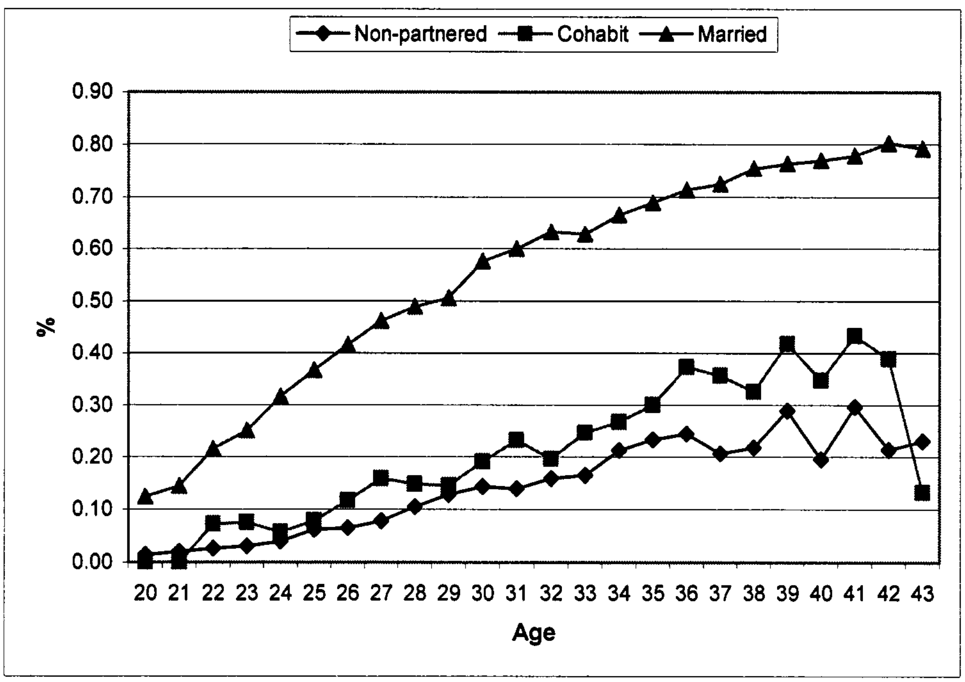

Figure 2. Percent of Men with Home Ownership by Family Status: 1985-2000 (NLSY79)

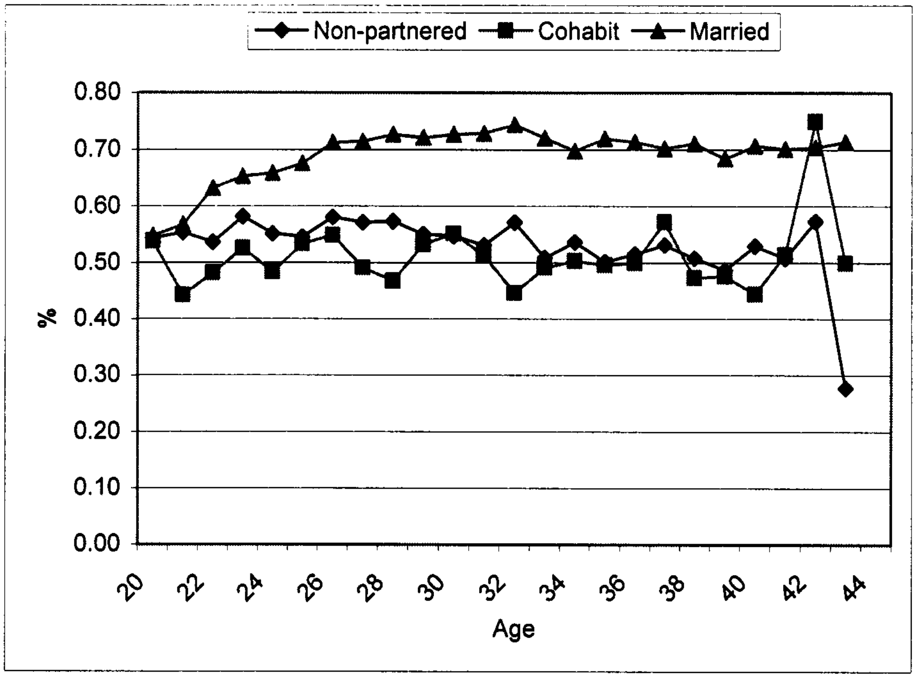

Figure 3. Percent of Women with Money-assets by Family Status: 1985-2000 (NLSY79)

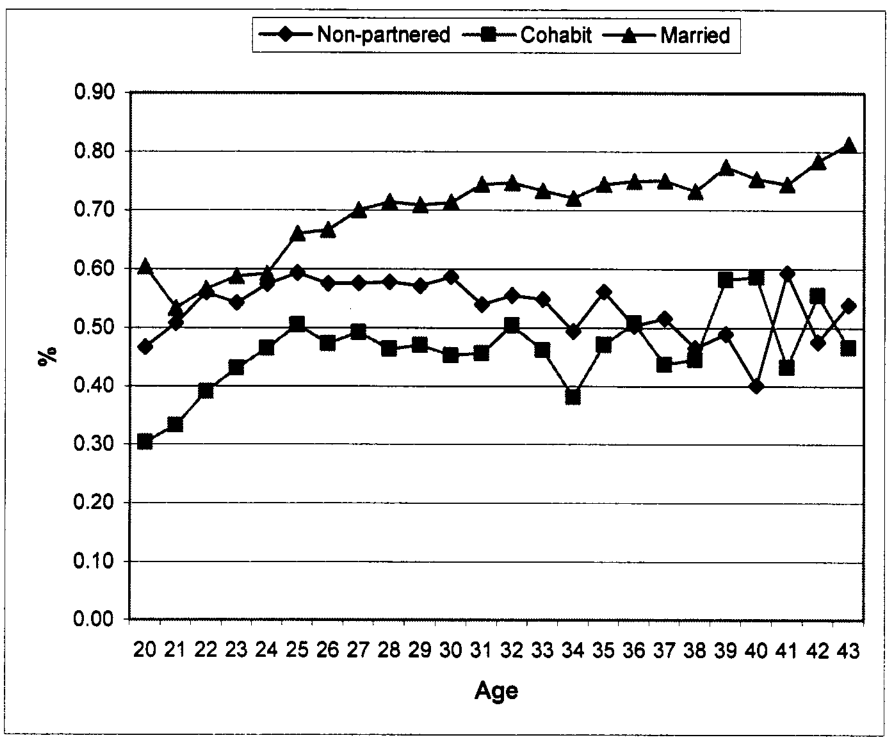

Figure 4. Percent of Men with Money-asset by Family Status: 1985-2000 (NLSY79)

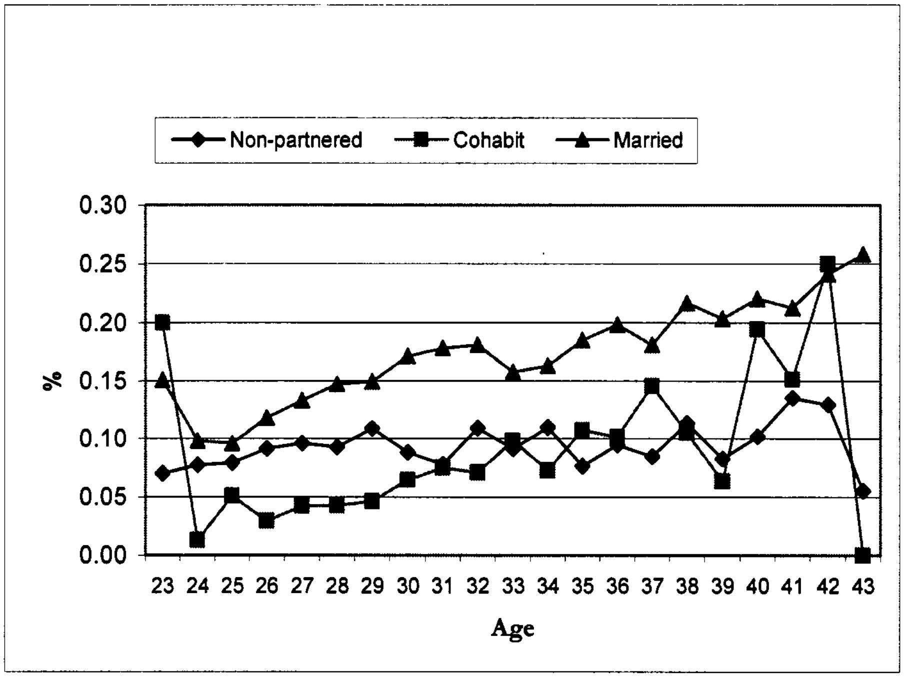

Figure 5. Percent of Women with Financial Investment by Family Status: 1988-2000 (NLSY79)

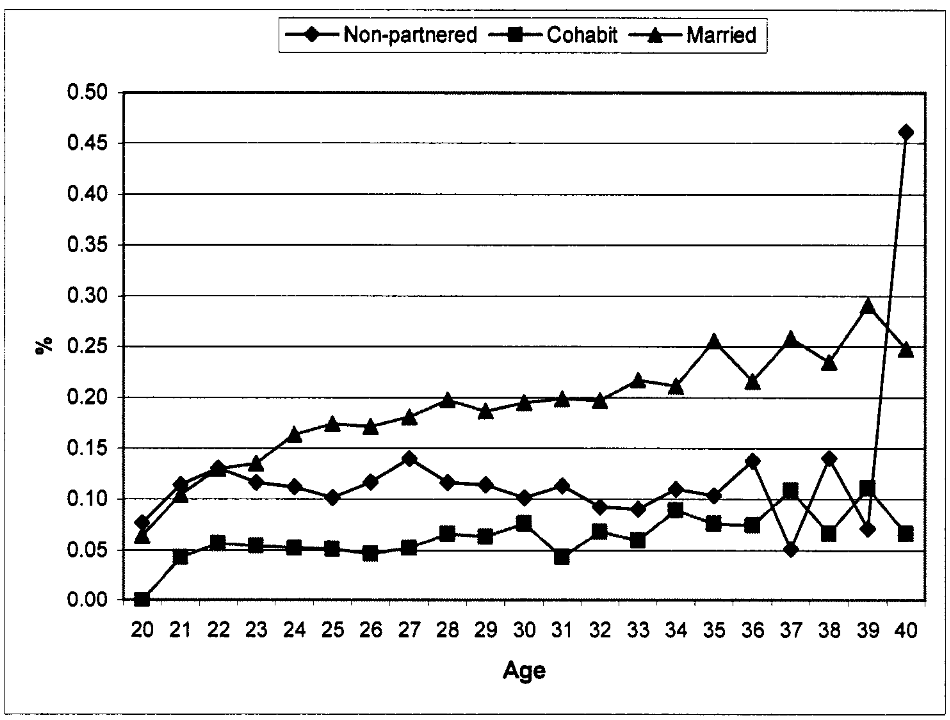

Figure 6. Percent of Men with Financial Investment by Family Status: 1988-2000 (NLSY79)

Table 10. Determinants of the Rate of Transitions
Hazard ratios from Cox Proportional Hazard Model

	Women			Men		
Variables	**Model 1**	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**
a. Transitions from Non-partnered to Partnered						
Own house	0.841	0.901	0.827	1.248	1.233	1.231
	(1.67)	(0.85)	(1.83)	(2.91)***	(2.11)**	(2.72)***
Have money asset	1.187	0.883	1.118	1.267	1.053	1.245
	(2.53)**	(1.13)	(1.65)	(4.10)***	(0.62)	(3.79)***
Own stocks & bonds		1.228			1.158	
		(1.63)			(1.46)	
Hispanic	0.883	0.828	0.879	0.838	0.911	0.836
	(1.19)	(1.19)	(1.24)	(1.95)	(0.80)	(1.97)**
Black	0.518	0.440	0.544	0.826	0.770	0.830
	(7.15)***	(5.91)***	(6.60)***	(2.67)***	(2.42)**	(2.60)***
Annual income			1.000			1.000
			(0.03)			(1.15)
Welfare recipiency			0.725			0.736
			(3.31)***			(2.23)**
N	11073	6342	11067	14717	8280	14691
b. Transitions from Non-partnered to Cohabiting						
Own house	0.887	0.922	0.877	1.119	1.257	1.115
	(0.62)	(0.37)	(0.67)	(0.82)	(1.40)	(0.79)
Have money asset	0.830	0.610	0.803	0.980	0.884	0.968
	(1.53)	(2.58)***	(1.74)	(0.23)	(1.00)	(0.36)
Own stocks & bonds		0.977			1.019	
		(0.10)			(0.10)	
Hispanic	0.835	0.752	0.832	0.635	0.601	0.636
	(0.99)	(1.02)	(1.01)	(3.09)***	(2.68)***	(3.08)***
Black	0.532	0.500	0.550	1.254	1.151	1.263
	(4.15)***	(3.25)***	(3.87)***	(2.14)**	(0.91)	(2.20)**
Annual income			1.000			1.000
			(0.23)			(1.08)
Welfare recipiency			0.840			0.784
			(1.09)			(1.34)
N	11073	6342	11067	14717	8280	14691
c. Transitions from Non-partnered to Married						
Own house	0.827	0.896	0.810	1.317	1.206	1.292
	(1.53)	(0.74)	(1.69)	(2.89)***	(1.47)	(2.68)***
Have money asset	1.422	1.081	1.314	1.526	1.241	1.491
	(4.13)***	(0.56)	(3.26)***	(5.56)***	(1.88)	(5.23)***
Own stocks & bonds		1.357			1.226	
		(2.02)**			(1.62)	
Hispanic	0.899	0.840	0.893	0.963	1.136	0.957
	(0.87)	(0.92)	(0.92)	(0.37)	(0.88)	(0.42)
Black	0.514	0.417	0.545	0.602	0.551	0.605
	(6.25)***	(5.29)***	(5.70)***	(5.38)***	(4.08)***	(5.32)***
Annual income			1.000			1.000
			(0.15)			(1.66)
Welfare recipiency			0.647			0.643
			(3.51)***			(2.06)**
N	11073	6342	11067	14717	8280	14691

Table 10 (Cont'd).

Variables	Women			Men		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
d. Transitions from Cohabiting to Married						
Own house	1.047	1.093	0.995	0.945	0.964	0.921
	(0.46)	(0.77)	(0.05)	(0.59)	(0.35)	(0.85)
Have money asset	1.448	1.394	1.331	1.842	1.686	1.766
	(4.21)***	(3.02)***	(3.00)***	(7.04)***	(4.98)***	(6.53)***
Own stocks & bonds		1.315			1.446	
		(1.64)			(2.90)***	
Hispanic	0.714	0.767	0.710	0.511	0.477	0.520
	(2.59)***	(1.69)	(2.63)***	(4.94)***	(4.53)***	(4.84)***
Black	0.577	0.563	0.578	0.432	0.472	0.433
	(4.27)***	(3.71)***	(4.24)***	(7.19)***	(5.48)***	(7.11)***
Annual income			1.000			1.000
			(2.45)**			(9.60)***
Welfare recipiency			0.913			0.619
			(0.81)			(2.67)***
N	3668	2496	3665	4076	3053	4070
e. Transitions from Unmarried (Non-partnered/Cohabiting) to Married						
Own house	1.237	1.127	1.223	1.402	1.348	1.380
	(2.56)**	(1.29)	(2.42)**	(4.90)***	(3.57)***	(4.65)***
Have money asset	1.344	1.229	1.280	1.586	1.435	1.547
	(4.62)***	(2.32)**	(3.81)***	(7.86)***	(4.40)***	(7.39)***
Own stocks & bonds		1.167			1.114	
		(1.36)			(1.12)	
Hispanic	0.733	0.726	0.735	0.816	0.844	0.813
	(3.55)***	(2.71)***	(3.53)***	(2.52)**	(1.53)	(2.56)**
Black	0.395	0.428	0.409	0.539	0.542	0.541
	(11.51)***	(7.51)***	(10.96)***	(8.50)***	(5.97)***	(8.45)***
Annual income			1.000			1.000
			(0.15)			(2.55)**
Welfare recipiency			0.792			0.701
			(2.69)***			(2.49)**
N	14652	8750	14643	18665	11206	18633

Note:
a. Robust z statistics in parentheses
b. ** significant at 5%; *** significant at 1%
c. In each specification, control variables include Age, age-squared, ethnicity, education, religion, father's and mother's education, region of residence, whether or not state recognizes common law marriage, and local unemployment rate. Estimated hazard ratios for these variables are available upon request.

Table 11.
F-Statistic for the Joint Significance of Instrumental Variables in the First Stage

Variables	Women		Men	
	Model 1	**Model 2**	**Model 1**	**Model 2**
No. of excluded instruments	18	21	18	21
Transition from never-married				
Own home	4.35	2.91	3.18	2.10
Have money assets	4.63	3.91	4.56	3.48
Own stocks, bonds		1.80		1.35
Transition from Cohabiting				
Own home	2.43	2.48	2.17	2.63
Have money assets	1.90	2.06	1.96	1.88
Own stocks, bonds		1.40		1.03
Transitions from unmarried (Non-partnered/cohabiting)				
Own home	5.88	4.22	4.96	3.41
Have money assets	4.99	4.81	6.11	4.80
Own stocks, bonds		2.08		1.58

Note:
a. F-statistic is for a hypothesis that the instrumental variables jointly have no effect.
b. Robust standard errors are calculated to account for clustering on each individual.

Table 12. Determinants of the Rate of Transitions
Probit and IV Probit Coefficients

Variables	Women: Model 1		Women: Model 2		Men: Model 1		Men: Model 2	
	Probit	IV Probit	Probit	IV Probit	Probit	IV Probit	Probit	IV Probit
a. Transitions from Non-partnered to Partnered								
Own home	-0.052	-0.156	-0.017	-0.085	0.112	-0.002	0.132	1.188
	(0.89)	(0.32)	(0.26)	(0.10)	(2.38)**	(0.00)	(2.32)**	(1.27)
Have money assets	0.085	-0.910	0.019	-1.023	0.129	-1.250	0.066	-0.626
	(2.14)**	(2.02)**	(0.33)	(1.49)	(4.00)***	(2.54)**	(1.45)	(0.74)
Own stocks, bonds			0.101	-2.248			0.096	-3.765
			(1.34)	(1.77)			(1.55)	(1.81)
N	11073	11073	6342	6342	14717	14717	8280	8280
J Statistic (Overidentification test) [d]		9.08		18.60		18.32		27.03
P-value for the J statistic		(.910)		(.417)		(.306)		(.078)
b. Transitions from Non-partnered to Cohabiting								
Own home	-0.015	0.121	-0.005	0.829	0.043	0.646	0.121	1.407
	(0.18)	(0.18)	(0.05)	(0.73)	(0.65)	(0.86)	(1.58)	(1.32)
Have money assets	-0.101	0.238	-0.132	0.541	-0.014	-1.705	-0.026	0.373
	(1.77)	(0.39)	(1.63)	(0.62)	(0.35)	(2.54)**	(0.45)	(0.38)
Own stocks, bonds			-0.005	-2.197			0.006	-2.549
			(0.04)	(1.25)			(0.07)	(1.08)
N	11073	11073	6342	6342	14717	14717	8280	8280
J Statistic (Overidentification test) [d]		8.33		13.56		20.95		19.57
P-value for the J statistic		(.938)		(.757)		(.180)		(.357)

Table 12 (Cont'd).

Variables	Women: Model 1		Women: Model 2		Men: Model 1		Men: Model 2	
	Probit	**IV Probit**	**Probit**	**IV Probit**	**Probit**	**IV Probit**	**Probit**	**IV Probit**
c. Transitions from Non-partnered to Married								
Own home	-0.057	-0.309	-0.019	-0.617	0.129	-0.408	0.108	0.294
	(0.87)	(0.56)	(0.26)	(0.64)	(2.41)**	(0.65)	(1.63)	(0.26)
Have money assets	0.175	-1.332	0.112	-1.760	0.205	-0.649	0.129	-1.090
	(3.92)***	(2.55)**	(1.73)	(2.14)**	(5.34)***	(1.19)	(2.35)**	(1.11)
Own stocks, bonds			0.130	-1.930			0.121	-3.880
			(1.57)	(1.33)			(1.76)	(1.55)
N	11073	11073	6342	6342	14717	14717	8280	8280
J Statistic (Overidentification test) [d]		12.21		15.22		24.23		30.78
P-value for the J statistic		(.729)		(.647)		(.085)		(.031)
d. Transitions from Cohabiting to Married								
Own home	0.111	-1.809	0.086	-1.146	0.003	-0.095	0.009	-1.005
	(1.58)	(2.51)**	(1.05)	(1.37)	(0.04)	(0.14)	(0.12)	(1.33)
Have money assets	0.226	0.858	0.224	-0.737	0.388	-1.096	0.319	-0.227
	(3.93)***	(1.28)	(3.11)***	(0.97)	(7.09)***	(1.68)	(4.84)***	(0.25)
Own stocks, bonds			0.170	0.218			0.311	-2.865
			(1.25)	(0.14)			(2.96)***	(1.24)
N	3668	3668	2496	2496	4076	4076	3053	3053
J Statistic (Overidentification test) [d]		15.91		18.03		21.68		19.60
P-value for the J statistic		(.459)		(.454)		(.154)		(.356)

Table 12 (Cont'd).

Variables	Women: Model 1		Women: Model 2		Men: Model 1		Men: Model 2	
	Probit	IV Probit	Probit	IV Probit	Probit	IV Probit	Probit	IV Probit
e. Transitions from Unmarried (Non-partnered/Cohabiting) to Married								
Own home	0.099	0.108	0.100	-0.127	0.167	-0.148	0.162	0.783
	(2.12)**	(0.25)	(1.84)	(0.17)	(4.00)***	(0.32)	(3.31)***	(0.81)
Have money assets	0.179	-0.917	0.151	-1.081	0.251	-0.704	0.197	0.184
	(5.12)***	(2.11)**	(3.15)***	(1.74)	(7.90)***	(1.66)	(4.61)***	(0.18)
Own stocks, bonds			0.095	-2.093			0.119	-7.716
			(1.39)	(1.57)			(2.10)**	(2.44)**
N	14652	14652	8750	8750	18665	18665	11206	11206
J Statistic (Overidentification test) [d]		16.84		21.42		13.40		22.25
P-value for the J statistic		(.396)		(.259)		(.643)		(.221)

Note:
a. z statistics in parentheses. Robust standard errors are calculated to account for clustering on each individual.
b. ** significant at 5%; *** significant at 1%
c. In each specification, control variables include Age, age-squared, ethnicity, education, religion, father's and mother's education, region of residence, whether or not state recognizes common law marriage, and local unemployment rate.
d. The Hansen-Sargan J Statistic is derived from a linear estimate of the binary dependent variable model.

Table 13.
Determinants of the Rate of Transition from Unmarried to Married: Women and Men by Education

Probit and IV Probit Coefficients in Model 1

Variables	HS or less educated		More than HS educated	
	Probit	**IV Probit**	**Probit**	**IV Probit**
a. Women				
Own house	0.104	-0.072	0.085	0.193
	(1.49)	(0.11)	(1.35)	(0.38)
Have money asset	0.228	-1.079	0.131	-0.947
	(5.08)***	(1.50)	(2.52)**	(2.08)**
First-stage F statistic				
Own house		3.06		4.65
Have money asset		1.81		4.52
N	7777	7777	6875	6875
b. Men				
Own house	0.166	0.087	0.164	-0.819
	(2.89)***	(0.13)	(2.70)***	(1.44)
Have money asset	0.276	0.101	0.241	-0.477
	(7.25)***	(0.18)	(4.50)***	(1.04)
First-stage F statistic				
Own house		3.03		2.79
Have money asset		3.43		5.89
N	11505	11505	7160	7160

Note:
a. The results reported in this table are estimated from specifications similar to model 1 in Table 12 (includes dummy indicators for homeownership and money assets).
b. z statistics in parentheses. Robust standard errors are calculated to account for clustering on each individual.
c. ** significant at 5%; *** significant at 1%

Chapter 3

A Two Period Model of Cohabitation and Marriage

I. Introduction

There have been dramatic changes in the demographic trends in the U.S. over the past several decades: delayed marriage, increased rates of divorce, increase in childbearing outside of formal marriage, and growth of cohabitation. Among these, the increasing prevalence of cohabitation has dwarfed the other demographic changes during this period. The number of cohabiting-couple households in the U.S. increased from 1.1 million (1.5 percent of all households) in 1977 to more than 4.6 million (4.2 percent of all households) in 2003 (Casper and Cohen, 2000; Fields, 2003). Moreover, the percentage of marriages preceded by cohabitation rose from about 10 percent for those marrying during the period 1965-1974 to well over 50 percent for those marrying during 1990-1994 (Bumpass and Sweet 1989; Bumpass and Lu 2000). Bramlett and Mosher (2002), using the National Survey of Family Growth 1995 data, also identify that about a third of all ever married women have cohabited at some point in their lives.[22] Taken together, these trends suggest that cohabitation has evolved in to a major mode of living arrangement – either as a transitional phase prior to a more permanent arrangement in marriage, or as an alternative family arrangement.

In light of these changes in the family structure, the objective of this paper is to develop an economic model of nonmarital cohabitation and marriage that is consistent with the current data on the formation and dissolution of relationships. A simple two-period theoretical model is developed in which participants do not fully realize the quality of their match with a potential partner in the first period. Based on their knowledge of the

[22] Similar demographic changes have been occurring in the other developed countries, (see, e.g., Cherlin, 2004, 2005; Kiernan 2000, 2002). Since this dissertation focuses on the population in the U.S., we discuss only the trends in the U.S.

distribution from which the match is drawn, agents must decide whether to form a marital relationship which may be relatively costly to dissolve, or a more informal nonmarital cohabiting relationship which may not have some of the benefits associated with marriage. Once in a relationship, in the second period, the partners observe the quality of the match and they must decide whether to dissolve the relationship. The model provides a simple framework where cohabitation and marriage may coexist, and captures the relative stability of the two types of relationships. Moreover, the model identifies some of the factors that influence the choice of family structure, and indicate how the choice of family form might affect women's labor supply decision.

In the following section we provide a brief review of the recent theoretical research which incorporates cohabitation as a distinct form of family union than marriage. Section III presents the basic framework and the implications of the model. Summary and conclusions follow in section IV.

II. Previous Theoretical Research Incorporating Cohabitation

Until very recently, the theory of the family largely ignored the distinction between marriage and cohabitation. In his seminal work, Gary Becker acknowledges the existence of nonmarital cohabitation, but does not distinguish between marriage and cohabitation (Becker, 1973, 1974). In fact, for his theoretical work, "marriage" is considered to mean that two individuals "share the same household". Since Becker's groundbreaking work, methodologically the economic analyses of the family have shifted from a household-centered approach of the unitary model, to an approach that is focused on the individual decision-maker. The latter approach includes the theoretical literature in a (cooperative or non-cooperative) bargaining framework (Manser & Brown 1980; McElroy & Horney 1981; Lundberg & Pollak 1994) as well as the collective framework put forward by Pierre Chiappori (Chiappori, 1988, 1992).[23] However, similar to Becker, this subsequent theoretical

[23] See Weiss (1997) for a recent survey of the theoretical literature on household formation and dissolution.

literature also continues to use 'marriage' and 'household formation' synonymously, without distinguishing between married and cohabiting households.

Only a handful of recent theoretical papers have attempted to analyze household decision-making while keeping in mind the individual's choice between forming a family union by marriage or by cohabitation. Brien, Lillard and Stern (2004) develop a theoretical model of union formation and dissolution in the presence of uncertain match quality. The information-gathering framework in their matching model provides an explanation for cohabitation, marriage, dissolution of relationships, and the relative stability of the two family forms. They estimate the structural parameters of the model using longitudinal data from a sample of female high school seniors in the U.S. They show that a significant cause of cohabitation is the need to learn about potential partners and to hedge against future bad shocks. The estimated parameters are used to conduct several comparative dynamic experiments. For example, they show that policy experiments changing the cost of divorce have little effect on relationship choices.

Drewianka (2003) classifies the decisions about marriage, fertility, divorce etc. together as "commitment decisions", and develops a general framework to analyze these decisions. In the model, which is built on a search framework, match quality is grounded on the idea of marital/partnership surplus. The paper analyzes the merit of a number of hypotheses about the recent demographic changes by comparing them with the comparative static results coming out of the model of commitment. The paper argues that decreasing opportunities of generating marital surplus (due to less specialization within the household, or decreasing employment opportunities for men, or substitution of "household commodities" with market goods and public institutions) can not explain the falling rates of marriage as well as the rising rates of cohabitation and divorce. The paper rather suggests that increased uncertainty about future match quality and lower search cost can explain the full pattern of demographic trends. However, as the paper focuses on explaining the major demographic changes by looking at marital surplus in a commitment framework, it pays no attention to how intra-household choices may vary by family structure, and how these choices may in turn influence the overall incentives to marry, cohabit, and divorce.

Murphy (2002) presents a model of marriage as an institution that changes the incentives of a mating game between men and women. The model considers that decisions to invest in children are not contractible ex ante, but must be sub-game perfect given that intimacy and pregnancy are sequential. Marriage and divorce, which are publicly observable, create costs for exiting a match; nonmarital cohabitation does not. Marriage, by providing an institution which makes a match observable, improves incentives for men to invest costly unobservable effort in their children. While the paper shows a rational basis of marriage as an institution, and also identifies cohabitation as an equilibrium outcome, it does not explain how the families resolve their household allocation decisions. Moreover, in this model, the choice between marriage and cohabitation are intertwined with the couple's fertility decision, which may not be the case in general.

Sahib and Gu (2002) uses a matching framework to develop a model of premarital cohabitation in which cohabitation serves as an information collecting period for potential partners, who at the end of the period decide if they should transform their relationship into a formal marriage. The model implies that couples set higher standards when forming marital unions than when forming cohabiting unions. The model also suggests positive assortative mating during premarital cohabitation, and indicates that some of the premarital cohabiting unions end in marriage while others end in separation.

Although intra-household resource allocation decisions are considered to provide the basis for a forward looking rational choice of family structure, in all of the studies discussed above the payoff of being in a particular type of family union is considered to be exogenous to the resource allocation decisions within the household. The current paper attempts to identify whether intra-household choices vary by the type of family union and also distinguish the factors that influence the choice of family type.

III. The Model

Let us consider a two-person household, consisting of a man (m) and a women (f), who make decisions about consumption and time allocation over two periods, $j = 1, 2$. Let us assume transferable utility, where for each partner

$$U_{ij} = G_j x_{ij} + \theta \ . \tag{1}$$

The utility of each individual i depends upon consumption of a household public good, G, and consumption of a private good, x_i. Let G be the quality of child produced. In each period j, the household public good is produced with inputs of m and/or f's time, $0 \le l_{ij} \le 1$, $(i \in m, f)$ and determined by the household production function

$$G_j = (h_m l_{mj} + h_f l_{fj}) \ , \tag{2}$$

θ is a random match parameter drawn from a probability distribution with cumulative distribution function $prob\{\theta \le s\} = F(s)$. In the first period, the couple does not observe the match parameter, but the cumulative distribution function that the match parameter is drawn from is common knowledge. Hence, in the first period they are only able to form expectations about the true matching parameter. In the second period, the true value for θ is realized. We assume that θ has the following distribution:

$$\theta \sim N(\mu, \sigma_\theta^2) \tag{3}$$

Both partners are interested in making inferences about θ. After drawing θ, each married partner expects to receive a "marriage-bonus" of μ in the first period. However, for a cohabiting couple, while they can estimate the mean of the matching parameter, just as a married couple would, they would not be able to materialize a payoff out of the expected value of the match. The "marriage bonus" may reflect greater social approval and heightened prestige (Cherlin, 2004), or the value of legal rights associated with marriage. Later on, in the second period, having observed the true θ, the partners must decide whether or not to break up. θ is set to zero if the couple is not in a union. Otherwise, they receive a "union-bonus" in the second period equal to the true matching parameter irrespective of the type of union they are in.

In period 1, when the partners move into a union, the initial wages w_{m1} and w_{f1} are given, and we assume that $w_{m1} \ge w_{f1}$. In the second period, wages are determined according to

$$w_{i2} = W_{i2}(t_{i1}) \tag{4}$$

where $t_{i1}(=1-l_{i1})$, is time spent at market work by partner i in the first period and $W_{i2}(t_{i1})$ is a monotone increasing function of t_{i1}. This relationship reflects that experience in labor market influences wages, representing a process of learning by doing where current work in the market affects future wages. Each individual's time endowment is normalized to one and all time not allocated to household public good production is spent in market work. We assume that individual f has higher productivity at home than m, and that she has a comparative advantage in home production:

$$h_f > h_m \text{ , and } \frac{h_f}{w_{fj}} > \frac{h_m}{w_{mj}} \text{, for } j=1,2 \tag{5}$$

We assume that preferences are not interdependent, and hence there is no altruism. To simplify further, we assume that saving and borrowing is not an option. We assume that within a union, the allocation of resources on the production of household public good is determined in a cooperative fashion, while if they live apart the allocation will be determined non-cooperatively. We solve the family's problem backwards, starting in the second (and last) period. Having observed the true θ and given the new wages, there are two possible states: either the partners remain in the union or they decide to separate. However, the scenario under break-up would vary depending on whether it was a marital or nonmarital union in the first place.

III.1. Decision making in the second period

A. Remain in the union (married or cohabiting)

The couple maximizes the objective function:

$$\max_{\{l_{i2},x_{i2}\}} U_{m2}+U_{f2} = G_2x_{m2}+\theta+G_2x_{f2}+\theta$$

$$= (h_m l_{m2}+h_f l_{f2})[x_{m2}+x_{f2}]+2\theta \tag{6}$$

such that, $x_{m2}+x_{f2} = w_{m2}(1-l_{m2})+w_{f2}(1-l_{f2})$

Solutions to the necessary first order conditions indicate the following:

$l^{*}_{m2}=0$

$$l_{f2}^{*}=1 \tag{7}$$

$$x_{m2}^{*}+x_{f2}^{*}=w_{m2}$$

If the partners remain in the union, under our assumption of f having absolute and comparative advantages in home production, there will be complete specialization in the household: the man (*m*) will specialize in market work and the woman (*f*) will spend her time working at home. The total indirect utility for the family is given by

$$U_{m2}^{*}+U_{f2}^{*}=h_f w_{m2}+2\theta \tag{8}$$

B. The union is dissolved

(i) Divorce from a marital union

Once a marriage is formed, dissolving it is costly (Weiss and Willis, 1997). First, there are legal costs associated with the divorce process and the division of property. Second, there is potential "social stigma" associated with being someone who had a "divorce".[24] This may also be considered in light of each individual's diminished value in the "(re)-marriage market" (in a broader sense of the term which encompasses prospects for forming nonmarital families as well). We assume that the partners m and f each face d as the financial cost of divorce, and further assume that cost of divorce is less than their earnings in the second period. In other words, we assume that $d < w_{i2}(0)$, $i = m, f$. We also assume that dissolution of families occurs through mutual consent.

A common arrangement in the event of a break-up is that one partner is selected as a custodian who determines the expenditure on the public good. We assume that individual f is the custodian in case of a break up. We, however, do not consider any transfer from the non-custodian parent to the custodian parent. In other words, child support transfers are not captured in our model.

The woman's allocation is determined by

[24] Other costs of dissolving a family union include loss of union-specific capital such as information about the partner's preference, as well as inefficiently low level of child-care expenditure (Weiss, 1997). These costs may not be specific to the dissolution of a marital union, rather are expected to be similar for dissolution of any family union.

$$\max_{\{l_{f2},x_{f2}\}} U_{f2} = G_2 x_{f2}$$

$$= (h_f l_{f2}) x_{f2} \tag{9}$$

such that, $x_{f2} + d = w_{f2}(1 - l_{f2})$

Solutions to the necessary first order conditions provide:

$$l_{f2}^{DM} = \frac{1}{2w_{f2}}(w_{f2} - d)$$

$$x_{f2}^{DM} = \frac{1}{2}(w_{f2} - d) \tag{10}$$

The following is the man's problem:

$$\max_{\{x_{m2}\}} U_{m2} = (\lambda G_2) x_{m2}$$

$$= (\lambda h_f l_{f2}^{DM}) x_{m2} \tag{11}$$

such that, $x_{m2} + d = w_{m2}$

Where $1 > \lambda > 0$ indicates that the man receives a fraction of the utility from the continued investment in child quality by the women. We may interpret λ as the measure of a man's tendency to free-ride on his previous spouse/partner's inputs to the production of child quality. Since the investment on the household public good is independently determined by the women, the man's resource allocation choice becomes:

$$x_{m2}^{DM} = w_{m2} - d \tag{12}$$

The implied utilities in the divorce state would be

$$U_{f2}^{DM} = \frac{h_f}{4w_{f2}}(w_{f2} - d)^2 \tag{13}$$

$$U_{m2}^{DM} = \lambda[\frac{h_f}{2w_{f2}}(w_{f2} - d)](w_{m2} - d) \tag{14}$$

These outcomes reflect the loss of efficiency in the allocation of the public good.

Proposition 1. *The divorce rule:* The couple will remain married for realizations of $\theta \geq \theta^M$,

where: $$\theta^M = \frac{h_f}{8w_{f2}}(w_{f2}-d)^2 + \frac{\lambda h_f}{4w_{f2}}(w_{f2}-d)(w_{m2}-d) - \frac{h_f}{2}w_{m2} \qquad (15)$$

and divorce otherwise.

Proof: Since divorce occurs through mutual consent, the joint utility in the divorce status has to be greater than that in the married status. The divorce rule derives from the comparison of aggregate indirect utility in the two statuses.

The conditions in proposition 1 suggest that a couple would remain married for some negative values of θ. That is, to avoid the loss of efficiency, the partners will stay in marriage despite the "mis-match" in their relationship, provided, of course, that the negative shock is not too large.

(ii) Separation from a cohabiting union

If the partners were in a cohabiting relationship in the first period, and if with the realization of the true θ in the second period they choose to separate, the decision making problems would be similar to those faced by a divorcing couple. The fundamental difference would lie only in the cost of dissolution. We consider that there would be no financial cost in dissolving a cohabiting union so that $d = 0$. Considering the same preference structure and the set of constraints as in (9) and (11), the optimal choice functions for the man (*m*) and the woman (*f*) would be:

$$l_{f2}^{DC} = \frac{1}{2}$$

$$x_{f2}^{DC} = \frac{1}{2}w_{f2} \qquad (16)$$

$$x_{m2}^{DC} = w_{m2}$$

Proposition 2. $l_{f2}^{DM} < l_{f2}^{DC} < l_{f2}^{*} = 1$

Proof: The proposition is easily verified by comparing the choices l_{f2}^{DM} in (10), l_{f2}^{DC} in (16), and l_{f2}^{*} in (7).

Thus, time spent on home production in the second period by a divorced woman is less than that by a woman who dissolved her cohabitating union. This difference is generated by the positive cost of divorce. But time spent in home production in any of the dissolved status is less than that in any partnered (married or cohabiting) family.

Since the household public good being considered here is child quality, such differences in inputs to the production of the household good may have important implications for children's outcomes. The proposition captures the essence of an extensive body of literature which identifies childhood family structure as a key factor associated with the children's achievements in their early adulthood. Experiences of life in a single-parent family during the childhood are usually associated with negative outcomes of children as young adults: lower educational attainments, higher risks of non-activity and early birth, and higher chances of smoking and experiencing psychological distress (Haveman and Wolfe, 1995; McLanahan, 1997; Ermisch and Francesconi, 2001). Our model suggests that such differences in outcomes can be explained in terms of the differences in allocation of resources under different family structures.

The individual's indirect utilities in the separated state would be

$$U_{f2}^{DC} = \frac{h_f}{4} w_{f2}, \text{ and } U_{m2}^{DC} = \lambda \frac{h_f}{2} w_{m2} \tag{17}$$

Proposition 3. *The separation rule:* The couple will remain cohabiting for realizations of $\theta \geq \theta^C$, where:

$$\theta^C = \frac{h_f}{8} w_{f2} + \lambda \frac{h_f}{4} w_{m2} - \frac{h_f}{2} w_{m2} \tag{18}$$

and break-up otherwise.

Proof: The joint utility in the separated status has to be greater than the utility in cohabitation as separation occurs through mutual agreement. The separation rule comes out of comparison of aggregate utility in cohabiting versus separated status.

Once again, from the conditions in proposition 3, we note that the cohabiting partners will remain in the union for some negative values of θ. In other words, the cohabiting couple would endure a "mis-match" to avoid the loss of efficiency, provided that the negative shock is not too large.

Proposition 4. $\theta^C > \theta^M$

Proof: See Appendix 2.

Proposition 4 implies that the quality of matching required to keep a marriage intact can be less than that required to keep a cohabiting couple together. And this arises due to the fact that dissolving a cohabiting family does not entail the material costs associated with a divorce. As a corollary, proposition 4 implies that $F(\theta^C) > F(\theta^M)$, since $F(\theta)$ is a positive monotone transformation of θ. Thus, in their first period decision making, if a couple choose to cohabit instead of being married, they bear a higher risk of breaking up in the future even with a better matching than would be required to keep a marriage intact.

There is ample empirical evidence which suggest that cohabiting unions tend to be shorter-lived than marriages. Data from the United States (Lundberg, 2005; Brien, Lillard and Stern, 2004; Willis and Michael, 1994; Bumpass and Sweet, 1989)[25] show that between 40 to 50 percent of all cohabiting relationships end in either marriage or separation within a year, and only a third of cohabiting couples are still cohabiting after two years. On the other hand, more than 95 percent of first marriages remain intact at the end of the first anniversary, and by the second anniversary, approximately 90 percent of marriages remain intact. About 50 percent of the marriages are dissolved only by the end of the fifth year. These estimates suggest indirectly that a cohabiting union is more likely to be dissolved than a marriage.

[25] Lundberg (2005) uses data from the National Longitudinal Survey of Youth 1979; Brien, Lillard and Stern (2004), and Willis and Michael (1994) use the National Longitudinal Study of High School Class of 1972; and Bumpass and Sweet, (1989) use the National Survey of Families and Households.

III.2. Decision making in the first period

Anticipating the possibility of a break up, the partners need to decide on their respective choice of work in the market and at home in the first period, when they only have information about the distribution of θ. We assume that the partners coordinate their work activities, and maximize their collective utilities with equal weights on each partner:

$$\max_{\{l_{f1}\}} \sum_i E \sum_j U_{ij}$$

$$\begin{aligned} &\equiv (h_f l_{f1})[w_{m1} + w_{f1}(1-l_{f1})] + 2E(\theta)M + [1-F(\theta^*)]h_f w_{m2} \\ &+F(\theta^*)[\frac{h_f}{4w_{f2}}(w_{f2}-d)^2 + \frac{\lambda h_f}{2w_{f2}}(w_{f2}-d)(w_{m2}-d)] + 2\int_{\theta^*}^{\infty} \theta dF(\theta) \end{aligned} \tag{19}$$

where, θ^* is either θ^M or θ^C. The appropriate reservation value between θ^M and θ^C is determined by the partners' decision on the type of union they choose to be in (marriage vs. cohabitation). Recall that d is considered to be zero for first-period cohabiters, and that $F(\theta)$ is the distribution of θ. We exploit the assumption on f's comparative advantage in home production and set m's work at home to zero (i.e., $l^*_{m1} = 0$).[26] M=1 if the couple is married, and 0 otherwise, so that the choice of the type union also determines whether the couple can materialize the "marriage-bonus", $E(\theta) = \mu$.

The female partner's choices of work at home under marriage and cohabitation, respectively, are: [27]

---

[26] The necessary and sufficient conditions are provided in Appendix 2.

[27] Note that $l^M_{f1} = 1$ if

$$w_{m1} > w_{f1} - \frac{F(\theta^M)}{4w^2_{f2}} \frac{\partial w_{f2}}{\partial l_{f1}} \{w^2_{f2} - d^2 + 2\lambda d(w_{m2} - d)\}$$

$$-\left\{\frac{\partial F(\theta^M)}{\partial \theta^M} \frac{\partial \theta^M}{\partial w_{f2}} \frac{\partial w_{f2}}{\partial l_{f1}}\right\}\left\{\frac{1}{4w_{f2}}(w_{f2}-d)^2 + \frac{\lambda}{2w_{f2}}(w_{f2}-d)(w_{m2}-d) - w_{m2}\right\}$$

and $l^C_{f1} = 1$ if

$$\begin{aligned}
l_{f1}^{M} &= \frac{1}{2}+\frac{1}{2}\frac{w_{m1}}{w_{f1}} \\
&+F(\theta^{M})\frac{1}{8w_{f1}w_{f2}^{2}}\frac{\partial w_{f2}}{\partial l_{f1}}[w_{f2}^{2}-d^{2}+2\lambda d(w_{m2}-d)] \\
&+\frac{1}{2w_{f1}}(\frac{\partial F(\theta^{M})}{\partial \theta^{M}}\frac{\partial \theta^{M}}{\partial w_{f2}}\frac{\partial w_{f2}}{\partial l_{f1}})[\frac{1}{4w_{f2}}(w_{f2}-d)^{2}+\frac{\lambda}{2w_{f2}}(w_{f2}-d)(w_{m2}-d)-w_{m2}]
\end{aligned} \tag{20}$$

$$\begin{aligned}
l_{f1}^{C} &= \frac{1}{2}+\frac{1}{2}\frac{w_{m1}}{w_{f1}}+F(\theta^{C})\frac{1}{8w_{f1}}\frac{\partial w_{f2}}{\partial l_{f1}} \\
&\quad +\frac{1}{2w_{f1}}(\frac{\partial F(\theta^{M})}{\partial \theta^{M}}\frac{\partial \theta^{M}}{\partial w_{f2}}\frac{\partial w_{f2}}{\partial l_{f1}})[\frac{1}{4}w_{f2}+\frac{\lambda}{2}w_{m2}-w_{m2}]
\end{aligned} \tag{21}$$

Note that because of the possibility of a dissolved family in the second period and because first period experience in the labor market influences future wages, women work less at home in the first period than the level which maximizes current family utility, both under cohabitation and marriage. This is generally referred to as the "investment effect" on labor supply (Weiss, 1997). In case of a marital union, the potential cost of divorce intensifies such investment effects.

Proposition 5. There will be less intra-household specialization under cohabitation than under marriage (i.e., $l_{f1}^{M}-l_{f1}^{C}>0$) if $\lambda<\frac{d}{2(w_{m2}-d)}$, and $0<d<\frac{2}{3}w_{m2}(1)$.

Proof: See Appendix 2.

In other words, cohabiting women would spend less time at home, and more time at market work in the first period compared to married women if divorce-cost is sufficiently small, and the man in dissolved family status receives sufficiently small share of utility from the household public good. In the context of the U.S., the condition of cost of divorce to be

---

$$w_{m1}>w_{f1}-\frac{1}{4}F(\theta^{C})\frac{\partial w_{f2}}{\partial l_{f1}}-\left\{\frac{\partial F(\theta^{C})}{\partial \theta^{C}}\frac{\partial \theta^{C}}{\partial w_{f2}}\frac{\partial w_{f2}}{\partial l_{f1}}\right\}\left\{\frac{1}{4}w_{f2}+\frac{\lambda}{2}w_{m2}-w_{m2}\right\}$$

sufficiently low compared to men's future earnings may not be too restrictive, especially when divorce is considered to occur through mutual agreement.

In general, however, the sign of the difference between women's market labor supply under marriage and cohabitation is ambiguous. This is because, in general, we cannot determine whether the "investment effect" on labor supply would be larger under cohabitation or under marriage. In the second period, the possibility of a dissolved family is higher for cohabiting families, and hence cohabiting women are potentially more likely to work in the second period than married women. But in case of dissolution, the level of market work is higher for divorced women than for women separated from cohabitation ($l_{f2}^{DM} < l_{f2}^{DC}$, as in proposition 2). So, cohabiting women enjoy the benefits of "investing" in labor market experience indirectly through their greater likelihood of being in a dissolved status. Married women, on the other hand, expect to enjoy the benefits of such "investment" directly through higher wages for higher labor supply in the future. Under the specified parametric conditions, the expected benefit of increasing first-period labor market experience will be smaller for married women than those who are cohabiting. Hence, in the first period, marriage would entail more specialization in the household than cohabitation.

There is some empirical evidence indicating that cohabiting women are more involved in the labor market than married women. Ressler and Waters (1995) using 1990 U.S. census data aggregated at the State level, and Henkens et. al. (2002) using data from the Netherlands show that increased rates of female labor force participation are associated with higher rates of cohabitation.

Total Indirect Utilities and Choice of Family Type. This difference in the allocation of resources that marriage and cohabitation entail, does not completely determine a couple's decision about marriage vis-à-vis cohabitation.[28] Comparison of total indirect utility under marriage and cohabitation suggest that despite the scope of differential levels of specialization under marriage and cohabitation, the individual may decide to cohabit or marry depending on which choice results in greater expected utility (see Appendix 2).

---

[28] Total indirect utility functions for marriage and cohabitation, and their difference are provided in the Appendix 2.

Proposition 6a. If $\lambda < \frac{d}{2(w_{m2}-d)}$ and $0 < d < \frac{2}{3}w_{m2}(1)$, any exogenous increase in men's first-period wages increases utility in both marriage and cohabitation. Moreover, the increase in utility is greater for marriage, which is indicated by an increase in the difference between the expected utilities in the two statuses.

Proof: Consider that both l_{f1}^{M} and l_{f1}^{C} have interior solutions. Using the envelope theorem,

$$\frac{\partial EU^M}{\partial w_{m1}} = h_f l_{f1}^M > 0\,;\ \frac{\partial EU^C}{\partial w_{m1}} = h_f l_{f1}^C > 0\,;\text{ and } \frac{\partial (EU^M - EU^C)}{\partial w_{m1}} = h_f (l_{f1}^M - l_{f1}^C)$$

If $\lambda < \frac{d}{2(w_{m2}-d)}$ and $0 < d < \frac{2}{3}w_{m2}(1)$, then $l_{f1}^{M} > l_{f1}^{C}$ which implies that

$$\frac{\partial EU^M}{\partial w_{m1}} > \frac{\partial EU^C}{\partial w_{m1}},\text{ and } \frac{\partial (EU^M - EU^C)}{\partial w_{m1}} > 0$$

This proposition suggests that people in any family union enjoy improved utility when the labor market conditions for men improve. But people in married families are better off than people in cohabiting families. In other words, there is more incentive to be married than to cohabit when men's labor market conditions improve. The proposition can be explained as a modified version of the Wilson hypothesis (Wilson, 1987, 1996) which considers decreased employment opportunities for men as the exogenous "prime mover" in reducing the marriage rates in the U.S., and argues that increasing employment opportunities for men would encourage marriage. It is well recognized that the primary challenge in testing this hypothesis empirically is to address the issue of endogeneity of men's labor market opportunities with respect their choice of family status. Only a small number of empirical studies have carefully attended the issue of endogeneity. The evidence coming out of these studies are mixed – some are supportive of the Wilson hypothesis that men's labor market opportunities are fundamental in determining the marriage rates, particularly in the low-income group (Olsen and Farkas, 1990; Black, McKinnish and Sanders, 2003), while others (e.g., Wood, 1995) find that decline in men's current economic status explains little of the decline in marriage rates, thereby contradicting the Wilson hypothesis..

Proposition 6b. Exogenous increase in women's wages in the first period increases utility in marriage and cohabitation, but such increases may not be equal for the two statuses.

Proof. Consider that both l_{f1}^{M} and l_{f1}^{C} have interior solutions. Using the envelope theorem,

$$\frac{\partial EU^M}{\partial w_{f1}} = h_f l_{f1}^M (1 - l_{f1}^M) > 0\,; \ \frac{\partial EU^C}{\partial w_{f1}} = h_f l_{f1}^C (1 - l_{f1}^C) > 0\,; \text{ and}$$

$$\frac{\partial (EU^M - EU^C)}{\partial w_{f1}} = h_f (l_{f1}^M - l_{f1}^C)[1 - (l_{f1}^M + l_{f1}^C)] \neq 0$$

As it appears, a change in women's first-period wages increases the marginal expected utilities for marriage as well as for cohabitation. Moreover, such changes in women' wages will change the gap between the utilities in the two statuses, which suggests that changes in women's first-period wages will have differential impact on the choice of marriage and cohabitation. Without any further specific assumptions, we cannot decide whether changes in women's first period wages will work in favor or against the choice of marriage compared to cohabitation.

Proposition 7. First period female labor supply increases (i.e., work at home decreases) when cost of divorce decreases if $\lambda \leq \dfrac{d}{w_{m2} - 2d}$ and $0 < d < \dfrac{1}{3} w_{m2}(1)$.

Proof. See Appendix 2.

Note that in general we can not determine the direction of the effect of a change in divorce cost on first-period labor supply by married women. Intuitively this is because, a change, say a decrease, in divorce cost increases the probability of divorce, but reduces the level of market work for divorced women in the second period. This creates ambiguity about the benefits of investing in first-period labor market experience. When divorce costs are low enough, the reduction in women's labor supply in a divorced status is outweighed by the increase in the probability of a future divorce, and women would clearly increase their first-period labor supply as divorce-cost decreases.

The proposition is consistent with the empirical evidence in Gray (1998) which suggest that wives who are favored by changes in divorce laws increased their labor supply. However, the proposition contradicts the findings reported in Chiappori, Fortin and Lacroix (2002) who find that adoption of divorce laws deemed favorable to women reduces wives' labor supply.

IV. Conclusion

In this paper we develop a two period model of family union to identify whether intra-household choices vary by the type of family union and also to distinguish some of the factors that influence the choice of family type. The analytical results from the model suggest that cohabiting unions would be shorter lived than marital unions; there will be less intra-household specialization under cohabitation than under marriage; improved labor market conditions for men provides stronger incentive for family union by marriage than by cohabitation; and cost of divorce affect married women's labor supply choice.

In the structure of the model analyzed here, cohabiting individuals did not have the option to get married in the second period. To incorporate this option within the current structure, one may consider a pre-draw stage (before decision making in the first period) when the cohabiting couples, without observing the match quality, could choose between continuing their relationship or be married in the first period. However, we did not consider such a pre-draw stage in our analysis to keep the problem tractable.

The analysis here was substantially simplified by the assumption of transferable utility which implies that partners have a mutual interest to coordinate their work activities if they stay in a family union. One might argue that while such coordination might be observed in marital unions, it would be less likely in a cohabiting union. It was shown that women in cohabiting unions will be more involved in the labor market than their married counterparts, essentially reflecting defensive investment anticipating the possibility of dissolution of the family union in the future. This investment effect will be exacerbated if the partners in a cohabiting union cannot coordinate their actions, and hence will increase the wedge between labor supply by married and by cohabiting women.

List of References

Aassve, Arnstein. 2003. "The Impact of Economic Resources on Premarital Childbearing and Subsequent Marriage Among Young American Women". *Demography.* 40(1). pp. 105-26.

Aizcorbe, Ana and Kennickell Arthur and Kevin Moore. 2003. "Recent changes in U.S. family finances: Evidence from the 1998 and 2001 Survey of Consumer Finance". *Federal Reserve Bulletin.* January. 89(1).

Antonovic, Kate and Town, Robert. 2004. "Are All Good Men Married? Uncovering the Sources of the Marital Wage Premium". *American Economic Review: Papers and Proceedings of the American Economic Association.* May. pp. 317 – 21.

Avery, Robert and Kennickell, Arthur. 1991. "Household Savings in the US". *Review of Income and Wealth.* December. 37(4). pp. 409-32.

Bardasi, Elena and Taylor, Mark. 2005. "Marriage and Wages". Working Papers of the Institute for Social and Economic Research. Paper # 2005-1. Colchester: University of Essex.

Bartlett, Robin and Callahan, Charles. 1984. "Wage Determination and Marital Status: Another Look". *Industrial Relations.* 23(1). pp. 90-96.

Basu, Kaushik; Narayan, Ambar and Ravallion, Martin. 2002. "Is Literacy Shared within the Households? Theory and Evidence for Bangladesh". *Labour Economics.* 8(October). pp. 649-65.

Becker, Gary S. 1973. "A Theory of Marriage: Part I". *Journal of Political Economy.* 81(4). pp. 813-46.

Becker, Gary S. 1974. "A Theory of Marriage: Part II". *Journal of Political Economy.* 82(2-II). pp. S11-S26.

Becker, Gary S. 1985. "Human Capital, Effort, and the Sexual Division of Labor". *Journal of Labor Economics.* 3(1, Part 2). pp. S33-58.

Becker, Gary S. 1991. *A Treatise on the Family.* Enlarged Edition. Harvard University Press. Cambridge, MA.

Benham, Lee. 1974. "Benefits of Women's Education within Marriage". *Journal of Political Economy.* 82(2, Part 2). pp. S57 – 71.

Black, Dan; McKinnish, Terra; and Sanders, Seth. 2003. "Does the Availability of High-Wage Jobs for Low-Skilled Men Affect Welfare Expenditures? Evidence from Shocks to the Steel and Coal Industries". *Journal of Public Economics.* 87. pp. 1921- 42.

Blackburn, McKinley and Sanders Korenman. 1994. "The Declining Marital-Status Earnings Differential." *Journal of Population Economics.* 7(3). July. pp. 247-270.

Blau, Francine; Kahn, Lawrence; and Waldfogel, Jean. 2000. "Understanding Young Women's Marriage Decisions: The Role of Labor and Marriage Market Conditions". *Industrial and Labor Relations Review.* July. 53(4). pp. 624-47.

Borsch-Supan, Axel. 1986. "Household Formation, Housing Prices, and Public Policy Impacts". *Journal of Public Economics.* 30. pp. 145-64.

Bound, John; Jaeger, David and Baker, Regina. 1995. "Problems with Instrumental Variables Estimation When the Correlation Between the Instruments and the Endogenous Explanatory Variable is Weak". *Journal of American Statistical Association.* 90(430). pp. 443-50.

Bowman, Cynthia G. 2004. "Legal Treatment of Cohabitation in the United States". *Law & Policy.* January. 26(1). pp. 119-51.

Bramlett, Mathew D. and Mosher, William. D. 2002. "Cohabitation, Marriage, Divorce, and Remarriage in the United States. National Center for Health Statistics". *Vital and Health* Statistics. 23(22). July.

Bratsberg, Bernt and Terrell, Dek. 1998. "Experience, Tenure and Wage Growth of Young black and white Men". *Journal of Human Resources.* 33(3). pp. 658-82.

Brien, Michael J. 1997. "Racial differences in marriage and the role of marriage markets" *Journal of Human Resources.* 32(4). pp. 741-78.

Brien, Michael; Lillard, Lee and Stern, Steven. 2004. "Cohabitation, Marriage and Divorce in a Model of Match Quality". University of Virginia Working Paper. (forthcoming in *International Economic Review*).

Brown, S., 2000. "Union Transitions among Cohabitors: The Significance of Relationship Assessments and Expectations. *Journal of Marriage and the Family.* 62 (3). pp. 833–47.

Browning, Martin and Annamaria Lusardi. 1996. "Household Saving: Micro Theories and Micro Facts". *Journal of Economics Literature.* 34(December). pp1797-1855

Bulcroft, Richard A. and Bulcroft, Kris A. 1993. "Race Differences in Attitudinal and Motivational Factors in the Decision to Marry." *Journal of Marriage and Family.* 55(2). pp. 338-55.

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

Bumpass, Larry L. and James A. Sweet. 1989. "National Estimates of Cohabitation". *Demography.* 26. pp. 615-25.

Carlson, Marcia; McLanahan, Sarah and England, Paula. 2004. "Union Formation in Fragile Families". *Demography.* 41(2). May. pp. 237-61.

Casper, Lynne M. and Cohen, Philip. N. 2000. "How does POSSLQ Measure Up? Historical Estimates of Cohabitation". *Demography.* 37(2). pp. 237-45.

Center for Human Resource Research (CHRR). 2001. *NLSY79 User's Guide: A Guide to the 1979-2000 National Longitudinal Survey of Youth Data.* The Ohio State University. Columbus, Ohio.

Cherlin, Andrew. 2004. "The Deinstitutionalization of American Marriage". *Journal of Marriage and Family.* 66. November. pp. 848-61.

Cherlin, Andrew. 2005. "American Marriage in the Early Twenty-First Century". *The Future of Children.* 15(2). Fall.

Chiappori, Pierre A. 1988. "Rational Household Labor Supply". *Econometrica.* 56. pp. 63-90.

Chiappori, Pierre A. 1992. "Collective Labor Supply and Welfare". *Journal of Political Economy.* 100 (June) pp. 437-67.

Chiappori, Pierre A.; Fortin, Bernard; and Lacroix, Guy. 2002. "Marriage Market, Divorce Legislation, and Household Labor Supply". *Journal of Political Economy.* 110(1). pp. 37-72.

Clarkberg, M., 1999. The Price of Partnering: The Role of Economic Well-Being in Young Adults' First Union Experiences. *Social Forces.* 77 (3). pp. 945–68.

Cleves, Mario, William Gould and Gutierrez, Roberto. 2004. *An Introduction to Survival Analysis Using Stata* (Revised Edition). Stata Corporation. College Station, TX

Cohen, Philip N. 1999. "Racial-Ethnic and Gender Differences in Returns to Cohabitation and Marriage: Evidence from the Current Population Survey". *U.S. Bureau of Census,* Population Division Working Paper # 35.

Cohen, Philip N. 2002. "Cohabitation and the Declining Marriage Premium for Men". *Work and Occupations.* 29(3). pp. 346-63.

Cornwell, Chirstopher and Rupert, Peter. 1997. "Unobservable Individual Effects, Marriage and Earnings or Young Men". *Economic Inquiry.* April. 35(2). pp. 285-294.

Daniel, Kermit. 1992. "Does Marriage Make Men More Productive?" *Economics Research Center, NORC,* Working Paper 92-2. and University of Chicago.

Drewianka, Scott. 2003. "Theories of Family Structure: Implications from a Generalized Model of Commitment". Manuscript. University of Wisconsin – Milwaukee.

Easterlin, Richard A. "Relative Economic Status and the American Fertility Swing". In E.B. Sheldon (ed.) *Family Economic Behaviour: Problems and Prospects.* Lippincott. Toronto and Philadelphia.
Edin, Kathryn. 2000. "What Do Low-Income Single Mothers Say About Marriage?". *Social Problems.* 47. pp. 112-33.

Ellwood, David and Crane, Jonathan. 1990. "Family Change among Black Americans: What Do We Know?". *Journal of Economic Perspective.* 4(4). pp. 65-84.

Ermisch, John and Francesconi, Marco. 2001. "Family Structure and Children's Achievements". *Journal of Population Economics.* 14. pp. 249–70.

Fields, Jason and Casper, Lynne M.. 2001. "America's Families and Living Arrangements: March 2000". *Current Population Reports*, P20-537. U.S. Census Bureau, Washington, DC.

Fields, Jason. 2003. "America's Families and Living Arrangements: 2003". *Current Population Reports*, P20-553. U.S. Census Bureau, Washington, DC.

Gibson, Christina; Kathryn Edin and Sara McLanahan. "High Hopes But Even Higher Expectations: The Retreat From Marriage Among Low-Income Couple". *Center for Research on Child Wellbeing Working Paper* # 03-06-FF. June 2003 (Revised: June 2004).

Ginther, Donna K. and Zavodny, Madeline. 2001. "Is the Male Marriage Premium due to Selection? The Effect of Shotgun Weddings on the Return to Marriage". *Journal of Population Economics*. 14(2). pp. 313-28.

Gray, Jeffrey S. 1997. "The Fall in Men's Return to Marriage: Declining Productivity Effects or Changing Selection?" *Journal of Human Resources*, 32(3). pp. 481-504.

Gray, Jeffrey. 1998. "Divorce-Law Changes, Household Bargaining, and Married Women's Labor Supply". *American Economic Review*. 88(3). pp. 628-42.

Grossbard-Shechtman, Shoshana. 1993. *On the Economics of Marriage: A Theory of Marriage, Labor and Divorce*. Westview Press. Boulder, Colorado.

Gupta, Nabanita Datta; Smith, Nina; and Stratton, Leslie S. 2005. "Is Marriage Poisonous? Are Relationships Taxing? An Analysis of the Male Marital Wage Differential in Denmark". IZA Bonn Discussion Paper No. 1591.

Hahn, Jinyong and Hausman, Jerry. 2002. "Notes on bias in estimators for simultaneous equation models". Economics Letters. 75. pp. 237–241.

Haurin, Donald; Hendershott, Patric; and Kim, Dongwook, 1993. "The Impact of Real Rents and Wages on Household Formation". *The Review of Economics and Statistics*. 75(2). pp. 284-93.

Haveman, Robert and Wolfe, Barbara. 1995. "The Determinants of Children's Attainments: A Review of Methods and Findings". *Journal of Economics Literature*. 33. pp. 1829-78.

Henkens, Kene; Yolanda Grift and Jacques Siegers. 2002. "Changes in Female Labor Supply in the Netherlands 1989-1998: The Case of Married and Cohabiting Women". *European Journal of Population*. 18(1). pp. 39-57.

Hersch, Joni and Stratton, Leslie S. 2000. "Household Specialization and the Male Marriage Wage Premium". *Industrial and Labor Relations Review*. 54(1). pp. 78-94.

Hill, Martha. 1979. "The Wage Effects of Marital Status and Children". *Journal of Human Resources*. 14(4). pp. 579-94.

Jepsen, Lisa. 2005. "The Relationship between Wife's Education and Husband's Earnings: Evidence from 1960-2000". *Review of Economics of the Household*. 3(2). pp. 197-214.

Kenny, Lawrence. 1983. "The Accumulation of Human Capital During Marriage by Males". *Economic Inquiry.* 21(2). pp. 223-31.

Kiernan, Kathleen. 2000. "European Perspectives on Union Formation". In Linda Waite (ed.). *The Ties that Bind: Perspectives on Marriage and Cohabitation,.* New York: Aldine de Gruyter.

Kiernan, Kathleen. 2002. "Cohabitation in Western Europe: Trends, Issues, and Implications". In A. Booth and A. C. Crouter (eds). *Just Living Together: Implication of Cohabitation on Families, Children, and Social Policy.* pp. 3–31. Mahwah, NJ: Erlbaum.

Korenman, Sanders and Neumark, David. 1991. "Does Marriage Really Make Men More Productive?". *The Journal of Human Resources.* 26(2). pp. 282-307.

Korenman, Sanders and Neumark, David. 1992. "Marriage Motherhood, and Wages". *The Journal of Human Resources.* 27(2). pp. 233-55.

Lam, David. 1988. "Marriage Markets and Assortative Mating with Household Public Goods: Theoretical Results and Empirical Implicaitons". *Journal of Human Resources.* 23(4). pp. 462-87.

Lancaster, Tony. 1990. *The Econometric Analysis of Transition Data.* Cambridge University Press. Cambridge, UK.

Loh, Eng Seng.1996. "Productivity Differences and the Marriage Wage Premium for white Males". *Journal of Human Resources.* 31(3). pp. 566–89.

Lundberg, Shelly and Pollak, Robert. 1993. "Separate Spheres Bargaining and the Marriage Market". Journal of Political Economy. December 1993. pp. 988-1010.

Lundberg, Shelly and Pollak, Robert. 1994. "Noncooperative Bargaining Models of Marriage". American Economic Review Papers and Proceedings. May 1994. pp. 132-137.

Lundberg, Shelly and Pollak, Robert. 1996. " Bargaining and Distribution in Marriage". Journal of Economic Perspectives. Fall 1996. pp. 139-158.

Lundberg, Shelly. 2005. "Men and Islands: Dealing with the Family in Empirical Labor Economics". Forthcoming in *Labour Economics*, 2005. (Presented as the Adam Smith Lecture, European Association of Labour Economists Annual Meeting, 2004).

Lupton, Joseph and Smith, James. 2003. "Marriage, Assets and Saving". Chapter 6 in *Marriage and the Economy: Theory and Evidence from Advanced Industrial Societies.* by Shoshana Grossbard-Shechtman (ed). Cambridge University Press.

Lusardi, Annamaria; Cossa, Ricardo and Krupka, Erin. 2001. "Savings of Young Parents". *Journal of Human Resources.* (Special Issue on Early Results from the National Longitudinal Survey of Youth 1997). Autumn. 36(4). pp. 762-94.

MacDonald, Maurice M. and Rindfuss, Ronald R. 1981. "Earnings, Relative Income, and Family Formation". *Demography.* 18(2). pp. 123-36.

Manning, W.D., Smock, P.J., 1995. "Why Marry? Race and the Transition to Marriage among Cohabitors" *Demography.* 32 (4). pp. 509–20.

Manning, Wendy and Smock, Pamela J. 2003. "Measuring and Modeling Cohabitation: New Perspectives from Qualitative Data," presented at the Population Association of American Annual Meeting, Minneapolis, MN, May 3, 2003.

Manser, M. and Brown, M. 1980. "Marriage and Household Decision Theory: A Bargaining Analysis". *International Economic Review.* February. 21(1):21. 34.
Mare, R. D. and Winship, C. 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. The Urban Institute. Washington, DC.

McElroy, M. B. and Horney, M. J. 1981. "Nash Bargained Household Decision". *International Economic Review.* June. 22(2).

McLanahan, Sara and Casper, Lynne. 1995. "Growing Diversity and Inequality in the American Family", In *State of the Union: America in the 1990s.* Edited by Reynolds Farley. Vol. 2. pp. 1-45. Russell Sage Foundation. New York.

McLanahan, Sarah. 1997. "Parent Absence or Poverty: Which Matter More?". In GJ Duncan and J. Brooks-Gunn (eds.) *Consequences of Growing Up Poor.* Russell Sage Foundation. New York.

Moffitt, Robert A. 1998. "The Effect of Welfare on Marriage and Fertility". In Robert Moffitt edited *Welfare, the Family, and Reproductive Behavior.* pp. 50-97. National Academy Press. Washington, DC.

Moffitt, Robert A. 2000. "Female Wages, Male Wages, and the Economic Model of Marriage: The Basic Evidence". In *Ties That Bind: Perspectives on Marriage and Cohabitaiton.* Edited by Linda Waite, Christine Bachrach, Michelle Hindin, Elizabeth Thomson, and Arland Thornton. Aldine de Gruyter. New York.

Mulder, Clara and Wagner, Michael. 2001. "The Connections between Family Formation and First-time Home Ownership in the Context of West Germany and the Netherlands". *European Journal of Population.* 17(2). pp. 137-64.

Murphy, Russell D. Jr. 2002. "A Good Man is Hard to Find: Marriage as an Institution". *Journal of Economic Behavior and Organization.* 47. pp. 27-53.

Nakosteen, Robert and Zimmer, Michael. 1987. "Marital Status and Earnings of Young Men". *Journal of Human Resources.* 22(2). pp. 248-68.

Neuman, Shoshana and Ziderman, Adrian. 1992. "Benefits of Women's Education Within Marriage: Results for Israel in a Dual Labor Market Context". *Economic Development and Cultural Change.* January. 40(2). pp. 413-24.

Olsen, R., Farkas, G., 1990. The effect of economic opportunity and family background on adolescent cohabitation and childbearing among low-income blacks. Journal of Labor Economics 8 (3), 341–362.

Olsen, Randall and Wolpin, Kenneth. 1983. "The Impact of Exogenous Child Mortality on Fertility: A Waiting Time Regression with Dynamic Regressors". *Econometrica.* 51. pp. 731-49.

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

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

Oppenheimer, Valerie K. 1997. "Women's Employment and the Gains to Marriage: The Specialization and Trading Model of Marriage". *Annual Review of Sociology.* 23. pp. 431-53.

Peters, Elizabeth. 1986. "Marriage and Divorce: Informational Constraints and Private Contracting". *American Economic Review.* 76. pp. 437 -54.

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. pp. 973–83.

Reed, Robert W. and Harford, Kathleen. 1989. "The Marriage Premium and Compensating Wage Differentials". *Journal of Population Economics.* 2(4). pp. 237-65.

Ressler, Rand and Waters, Melissa. 1995. "The Economics of Cohabitation". *Kyklos.* 48(4). pp. 577-92.

Ribar, David. 2004. "What Do Social Scientists Know About the Benefits of Marriage? A Review of Quantitative Methodologies". *IZA Discussion Paper* No. 998. January.

Richardson, Katarina. 2000. "The Evolution of the Marriage Premium in the Swedish Labor Market 1968-1991." Unpublished manuscript, Labor Market Policy Evaluation. Uppsala, Sweden, June.

Rivers, Douglas and Vuong, Quang. 1988. "Limited Information Estimators and Exogeneity Tests for Simultaneous Probit Models". *Journal of Econometrics.* Novemeber. 39(3). pp. 347-66.

Sahib, Padma Rao and Gu, Xinhua. 2002. ""Living in Sin" and Marriage: A Matching Model". *Journal of Population Economics.* 15. pp. 261-82.

Sanchez, L., Manning, W., Smock, P. 1998. "Sex-specialized or Collaborative Mate Selection. Union Transitions among Cohabitors". *Social Science Research.* 27. pp. 280–94.

Sassler, S. and McNally, J. 2003. "Cohabiting Couples' Economic Circumstances and Union Transitions: A Re-examination Using Multiple Imputation Techniques". *Social Science Research.* 32. pp. 533-78.

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

Schultz, T. Paul. 1994. "Marital Status and Fertility in the United States: Welfare and Labor in the Market effects". *Journal of Human Resources.* 29(2). Pp, 636-69.

Seff, Monica A. 1995. "Cohabitation and the Law". *Marriage and Family Review.* 21(3-4). pp. 141-65.

Seidman, Lawrence. "Assets and the Tax Code". In *Assets for the Poor: The Benefits of Spreading Asset Ownership.* Edited by Thomas Shapiro and Edward Wolff. pp. 324-56. Russell Sage Foundation. New York.

Seltzer, Judith A. 2000. "Families Formed Outside of Marriage". *Journal of Marriage and the Family.* 62(4). pp. 1247-68.

Sherraden, Michael. 2001. "Asset-Building Policy and Programs for the Poor". In *Assets for the Poor: The Benefits of Spreading Asset Ownership.* Edited by Thomas Shapiro and Edward Wolff. pp. 302-23. Russell Sage Foundation. New York.

Smock, Pamela J., and Wendy D. Manning. 1997. "Cohabiting Partners' Economic Circumstances and Marriage." *Demography* 34(3). pp. 331-41.

Spitko, E. Gary. 2002. "An Accrual/Multi-Factor Approach to Intestate Inheritance Rights for Unmarried Committed Partners". *Oregon Law Review.* 255. Summer. 81(2).

Stern, Mark. 2001. "The Un(credit)worthy Poor: Historical Perspectives on Policies to Expand Assets and Credit". In *Assets for the Poor: The Benefits of Spreading Asset Ownership.* Edited by Thomas Shapiro and Edward Wolff. pp. 269-301. Russell Sage Foundation. New York.

Stratton, Leslie S. 2002. "Examining the Wage Differential for Married and Cohabiting Men". *Economic Inquiry*, 40(2), pp. 199–212.

Thornton, Arland and Linda Young-DeMarco. 2001. "Four Decades of Trends in Attitudes Toward Family Issues in the United States: The 1960s Through the 1990s." *Journal of Marriage and the Family* 63(4):1009-1037.

Tiefenthaler, Jill. 1997. "The Productivity Gains of Marriage: Effects of Spousal Education on Own Productivity Across Market Sectors in Brazil". *Economic Development and Cultural Change.* April. 45(3). pp. 633-50.

Tucker, Belinda M. and Mitchell-Kernan, Claudia. 1995a. "Trends in African American Family Formation: A Theoretical and Statistical Overview". In Belinda Tucker and Claudia Mitchell-Kernan (Eds). *The Decline in Marriage Among African-Americas.* Chapter 1. pp. 3 -26. New York: Russell Sage Foundation.

Tucker, M. Belind. 2000. "Marital Values and Expectations in Context: Results from a 21-city Survey". In *Ties That Bind: Perspectives on Marriage and Cohabitaiton.* Edited by Linda Waite, Christine Bachrach, Michelle Hindin, Elizabeth Thomson, and Arland Thornton. Aldine de Gruyter. New York.

Weiss, Yoram and Willis, Robert J. 1985. "Children as Collective Goods". *Journal of Labor Economics.* vol. 3.July. pp. 268-92.

Weiss, Yoram. 1997. "The Formation and Dissolution of Families: Why Marry? Who Marries Whom? And What Happens Upon Divorce?". Ch. 3 in Rosenzweig, Mark and Stark, Oded (eds.) *Handbook of Population and Family Economics*. Elsevier.

Weiss, Yoram and Willis, Robert J. 1997. "Match Quality, New Information, and Marital Dissolution". *Journal of Labor Economics*. vol. 15. no. 1 part 2. January. pp. S293-S329.

Willis, Robert and Michael, Robert. 1994. "Innovation in Family Formation: Evidence on Cohabitation in the United States". In John Ermisch and Naohiro Ogawa (Eds). *The Family, the Market and the State in Aging Societies*. Oxford University Press. pp. 9-45.

Wilson, William J. 1987. *The Truly Disadvantaged: The Inner City, the Underclass, and Public Policy*. University of Chicago Press, Chicago.

Wilson, William J. 1996. *When Work Disappears: The World of the New Urban Poor*. Alfred A. Knopf, New York.

Wolff, Edward N. 2001. "Recent Trends in Wealth Ownership, from 1983 to 1998". In *Assets for the Poor: The Benefits of Spreading Asset Ownership*. Edited by Thomas Shapiro and Edward Wolff. pp. 34-73. Russell Sage Foundation. New York.

Wood, Robert. 1995. "Marriage Rates and Marriageable Men: A Test of Wilson Hypothesis". *Journal of Human Resources*. 30(1). pp. 163-93.

Wooldridge, Jeffrey. 2002. *Econometric Analysis of Cross Section and Panel Data*. The MIT Press. Cambridge, MA.

Wu, Z., Pollard, M., 2000. "Economic Circumstances and the Stability of Nonmarital Cohabitation". *Journal of Family Issues*. 21(3). pp. 303-28.

Xie, Yu; Raymo, James; Goyette, Kimberly and Thornton, Arland. 2003. "Economic Potential and Entry into Marriage and Cohabitation". *Demography*. May. 20(3). pp. 351-67.

Appendix 1. Appendix to Chapter 1

Table A-1. Summary Characteristics of Sample by Family Union Status NLSY White Men, 1990

Variables	Never Married	Cohabiting [a]	Married	Divorced or Separated
Hourly wage (in dollars) [b]	11.70	10.33	12.41	9.99
Hours worked per week	43.58	45.62	46.51	45.09
Weeks worked per calendar year	46.51	47.09	49.64	46.58
Age (in years)	28.31	28.46	29.35	29.46
Yrs in marriage, total	-	2.02	6.17	5.17
Yrs in current marriage	-	-	5.77	-
Yrs div. or sep.	-	1.01	-	2.32
Yrs cohabiting	0.32	2.17	-	0.32
Yrs cohabited with wife before marriage	-	-	2.04	-
Has schooling level				
Less than high-school (<12)	0.10	0.18	0.10	0.16
High-school grad (=12)	0.41	0.53	0.51	0.63
Some college (>12 & <16)	0.18	0.17	0.17	0.15
College grad (>=16)	0.32	0.13	0.23	0.06
Partner's schooling level				
Less than high-school (<12)	-	0.15	0.08	-
High-school grad (=12)	-	0.54	0.48	-
Some college (>12 & <16)	-	0.18	0.22	-
College grad (>=16)	-	0.10	0.22	-
Missing	-	0.03	0.01	-
Lives in the				
North east	0.24	0.22	0.20	0.16
North central	0.36	0.32	0.37	0.34
South	0.21	0.22	0.27	0.31
West	0.19	0.24	0.16	0.18
Lives in SMSA	0.79	0.83	0.71	0.76
Local unemployment rate	2.34	2.28	2.39	2.33
Years in current job	2.88	2.93	4.39	3.02
Is covered by union	0.14	0.15	0.18	0.16
Sample Size (1990)	395	78	734	137
Proportion of Sample in different status:				
1990	0.31	0.06	0.58	0.11
2000	0.15	0.05	0.69	0.17

Note: a. About 63 percent of currently cohabiting white men are never married.
b. For our sample we dropped the cases where hourly wages fell outside the $1 - $100 range.

Table A-2. Summary Characteristics of Sample by Family Union Status NLSY Black Men, 1990

Variables	Never Married	Cohabiting [a]	Married	Divorced or Separated
Hourly wage (in dollars) [b]	7.99	8.36	9.98	8.07
Hours worked per week	42.05	43.14	44.04	43.85
Weeks worked per calendar year	42.98	42.95	47.51	41.64
Age (in years)	28.56	28.78	29.40	29.54
Yrs in marriage, total	-	1.31	5.13	5.39
Yrs in current marriage	-	-	4.89	-
Yrs div. or sep.	-	0.40	-	1.32
Yrs cohabiting	0.68	3.49	-	0.38
Yrs cohabited with wife before marriage	-	-	2.13	-
Has schooling level				
Less than high-school (<12)	0.21	0.24	0.14	0.20
High-school grad (=12)	0.53	0.57	0.49	0.58
Some college (>12 & <16)	0.17	0.12	0.20	0.18
College grad (>=16)	0.09	0.07	0.17	0.05
Partner's schooling level				
Less than high-school (<12)	-	0.17	0.07	-
High-school grad (=12)	-	0.51	0.50	-
Some college (>12 & <16)	-	0.17	0.22	-
College grad (>=16)	-	0.12	0.20	-
Missing	-	0.03	0.03	-
Lives in the				
North east	0.18	0.17	0.13	0.14
North central	0.18	0.20	0.17	0.18
South	0.57	0.59	0.63	0.60
West	0.06	0.04	0.07	0.07
Lives in SMSA	0.77	0.79	0.77	0.83
Local unemployment rate	2.23	2.25	2.15	2.16
Years in current job	2.39	2.09	3.62	2.14
Is covered by union	0.21	0.18	0.30	0.21
Sample Size (1990)	493	109	288	111
Proportion of Sample in different status:				
1990	0.55	0.12	0.32	0.13
2000	0.35	0.09	0.43	0.22

Note: a. About 76 percent of currently cohabiting black men are never married.
b. For our sample we dropped the cases where hourly wages fell outside the $1 - $100 range.

Table A-3. Summary Characteristics of Sample by Family Union Status NLSY Hispanic Men, 1990

Variables	Never Married	Cohabiting [a]	Married	Divorced or Separated
Hourly wage (in dollars) [b]	9.84	10.40	10.67	9.01
Hours worked per week	41.87	44.52	44.02	43.67
Weeks worked per calendar year	43.69	44.40	46.88	42.80
Age (in years)	28.39	28.21	29.14	29.30
Yrs in marriage, total	-	1.47	6.25	6.83
Yrs in current marriage	-	-	5.90	-
Yrs div. or sep.	-	0.52	-	1.62
Yrs cohabiting	0.77	3.72	-	0.42
Yrs cohabited with wife before marriage	-	-	2.21	-
Has schooling level				
Less than high-school (<12)	0.30	0.40	0.29	0.41
High-school grad (=12)	0.40	0.38	0.43	0.49
Some college (>12 & <16)	0.17	0.17	0.19	0.08
College grad (>=16)	0.12	0.04	0.09	0.02
Partner's schooling level				
Less than high-school (<12)	-	0.36	0.27	-
High-school grad (=12)	-	0.38	0.41	-
Some college (>12 & <16)	-	0.17	0.20	-
College grad (>=16)	-	0.04	0.11	-
Missing	-	0.04	0.03	-
Lives in the				
North east	0.15	0.15	0.15	0.16
North central	0.05	0.04	0.05	0.08
South	0.28	0.17	0.28	0.34
West	0.52	0.64	0.52	0.41
Lives in SMSA	0.87	0.81	0.85	0.85
Local unemployment rate	2.51	2.49	2.55	2.64
Years in current job	2.35	3.41	3.78	3.01
Is covered by union	0.14	0.21	0.25	0.28
Sample Size (1990)	193	47	318	61
Proportion of Sample in different status:				
1990	0.34	0.08	0.56	0.11
2000	0.20	0.09	0.63	0.17

Note: a. About 75 percent of currently cohabiting Hispanic men are never married.
b. For our sample we dropped the cases where hourly wages fell outside the $1 - $100 range.

Appendix 2. Appendix to Chapter 3

Table A-4: Assumptions for the Model

1. Transferable utility; no altruism; no saving and borrowing.
2. Individual *f* has the comparative advantage in home production.
3. The first period wage is given.
4. Decision making within the family union occurs in a collective framework, with equal weights on each partner's utility.
5. In the first period, there is a "marriage bonus".
6. Wages in the second period depend on the labor market experience in the first period.
7. Within a union, the allocation on the public good is determined in a cooperative fashion, while if they live apart the allocation will be determined non-cooperatively.
8. In a dissolved status, the non-custodial partner receives only a fraction of the utility from the household public good produced by the custodial partner. However, there is no transfer of resources from the non-custodial parent to the custodial parent.
9. Dissolution of families occurs through mutual agreement.
10. Cost of divorce is positive, but there is no cost to dissolving a cohabiting union. Further, cost of divorce is assumed to be less than the second period earnings.

Proposition 4. $\theta^C > \theta^M$

Proof: By comparing equations (15) and (18),

$$\theta^M - \theta^C = \frac{h_f}{8}\left\{\frac{[w_{f2}(t_{f1}^M)-d]^2}{w_{f2}(t_{f1}^M)} - w_{f2}(t_{f1}^C)\right\} + \frac{\lambda h_f}{4}\left\{\frac{[w_{f2}(t_{f1}^M)-d]}{w_{f2}(t_{f1}^M)}[w_{m2}(1)-d] - w_{m2}(1)\right\}$$

It is easily verified that with $0 < d < w_{i2}(0)$ (which is true by assumption),

$\theta^M - \theta^C < 0$ if

$$w_{f2}(t_{f1}^C) > w_{f2}(t_{f1}^M) - d \qquad \text{(A-1)}$$

We will consider that condition (A-1) holds for now, and show later that it is indeed satisfied by the choice of labor supply under marriage and cohabitation.

Conditions for $l_{m1}^{*} = 0$.

The first period utility maximization problem without substituting $l_{m1}^{*} = 0$ is:

$$\max_{\{l_{f1}\}} \sum_i E \sum_j U_{ij}$$

$$\equiv (h_m l_{m1} + h_f l_{f1})[w_{m1}(1-l_{m1}) + w_{f1}(1-l_{f1})] + 2E(\theta)M + [1-F(\theta^*)]h_f w_{m2}$$

$$+F(\theta^*)[\frac{h_f}{4w_{f2}}(w_{f2}-d)^2 + \frac{\lambda h_f}{2w_{f2}}(w_{f2}-d)(w_{m2}-d)] + 2\int_{\theta^*}^{\infty} \theta dF(\theta)$$

$$\frac{\partial EU}{\partial l_{m1}} \equiv h_m[w_{m1}(1-l_{m1}) + w_{f1}(1-l_{f1})] - h_m w_{m1} l_{m1} - h_f w_{m1} l_{f1}$$

$$+\begin{bmatrix} [1-F(\theta^M)]h_f + F(\theta^M)\dfrac{\lambda h_f}{2w_{f2}}(w_{f2}-d) \\ +\dfrac{\partial F(\theta^M)}{\partial \theta^M}\dfrac{\partial F\theta^M}{\partial w_{m2}}\left\{\dfrac{h_f}{4w_{f2}}(w_{f2}-d)^2 + \dfrac{\lambda h_f}{2w_{f2}}(w_{f2}-d)(w_{m2}-d) - h_f w_{m2}\right\} \end{bmatrix}\frac{\partial w_{m2}}{\partial l_{m1}}$$

We understand that the necessary condition for $l_{m1}^{*} = 0$ is $\dfrac{\partial EU}{\partial l_{m1}} < 0$, which implies that $l_{m1}^{*} = 0$ if (considering an interior solution for l_{f1}):

$$h_m[w_{m1} + w_{f1} - 2w_{m1}l_{m1} - w_{f1}l_{f1}]$$

$$< h_f\begin{bmatrix} w_{m1}l_{f1} - \dfrac{\partial w_{m2}}{\partial l_{m1}} \\ \begin{Bmatrix} [1-F(\theta^M)]h_f + F(\theta^M)\dfrac{\lambda h_f}{2w_{f2}}(w_{f2}-d) \\ +\dfrac{\partial F(\theta^M)}{\partial \theta^M}\dfrac{\partial F\theta^M}{\partial w_{m2}}\left\{\dfrac{h_f}{4w_{f2}}(w_{f2}-d)^2 + \dfrac{\lambda h_f}{2w_{f2}}(w_{f2}-d)(w_{m2}-d) - h_f w_{m2}\right\} \end{Bmatrix} \end{bmatrix}$$

Then, a sufficient condition for $l_{m1}^{*} = 0$ is

$$h_m = 0. \qquad \text{(A-2)}$$

Also note that in the case where $l_{f1}^{*} = 1$, it can be shown that $l_{m1}^{*} = 0$ will necessarily be true.

Proposition 5. $l_{f1}^M - l_{f1}^C > 0$ if $\lambda < \dfrac{d}{2(w_{m2}-d)}$ and $0 < d < \dfrac{2}{3} w_{m2}(1)$.

Proof: Consider that both l_{f1}^M and l_{f1}^C have interior solutions. Then, comparison of the choices regarding work at home under marriage and cohabitation, in equations (20) and (21) respectively, indicates:

$$l_{f1}^M - l_{f1}^C = \frac{1}{8w_{f1}w_{f2}^2}\frac{\partial w_{f2}}{\partial l_{f1}}\left\{F(\theta^M)[w_{f2}^2 - d^2 + 2\lambda d(w_{m2}-d) - F(\theta^C)w_{f2}^2\right\}$$

$$+\frac{1}{2w_{f1}}\frac{\partial w_{f2}}{\partial l_{f1}}\left\{\begin{array}{l}\dfrac{\partial F(\theta^M)}{\partial \theta^M}\dfrac{\partial \theta^M}{\partial w_{f2}}[\dfrac{1}{4w_{f2}}(w_{f2}-d)^2 + \dfrac{\lambda}{2w_{f2}}(w_{f2}-d)(w_{m2}-d) - w_{m2}] \\ -\dfrac{\partial F(\theta^C)}{\partial \theta^C}\dfrac{\partial \theta^C}{\partial w_{f2}}[\dfrac{1}{4}w_{f2} + \dfrac{\lambda}{2}w_{m2} - w_{m2}]\end{array}\right\} \quad \text{(A-3)}$$

We understand that $\dfrac{\partial w_{f2}}{\partial l_{f1}} < 0$. The necessary condition for the first element on the right hand side to be positive is

$$F(\theta^M)[w_{f2}^2 - d^2 + 2\lambda d(w_{m2}-d)] - F(\theta^C)w_{f2}^2 < 0 \quad \text{(A-4)}$$

Since $F(\theta^M) < F(\theta^C)$, a sufficient condition for (A-4) to hold, and thereby for the first element on the right hand side to be positive, is:

$$w_{f2}^2 - d^2 + 2\lambda d(w_{m2}-d) < w_{f2}^2$$

$$\Rightarrow 2\lambda d(w_{m2}-d) < d^2$$

$$\Rightarrow \lambda < \frac{d}{2(w_{m2}-d)}. \quad \text{(A-5)}$$

Now, $0 < \lambda < 1$. Hence, we need

$$\frac{d}{2(w_{m2}-d)} > 0, \text{ and } \frac{d}{2(w_{m2}-d)} < 1$$

$$\Rightarrow d > 0, \text{ and } \Rightarrow d < \frac{2}{3}w_{m2}(1) \quad \text{(A-6)}$$

Again, since by assumption $0 < d < w_{i2}(0)$, it is easily verified that

$$\frac{1}{4w_{f2}}(w_{f2}-d)^2+\frac{\lambda}{2w_{f2}}(w_{f2}-d)(w_{m2}-d)-w_{m2}<\frac{1}{4}w_{f2}+\frac{\lambda}{2}w_{m2}-w_{m2}$$

Also, $\frac{\partial F(\theta^M)}{\partial \theta^M}<\frac{\partial F(\theta^C)}{\partial \theta^C}$ since $\theta^M<\theta^C$ and $\theta \sim Normal$.

Now, $\frac{\partial \theta^M}{\partial w_{f2}}=\frac{h_f}{8}[\frac{w_{f2}^2-d^2+2\lambda d(w_{m2}-d)}{w_{f2}^2}]$, and $\frac{\partial \theta^C}{\partial w_{f2}}=\frac{h_f}{8}$.

Hence, $\frac{\partial \theta^M}{\partial w_{f2}}<\frac{\partial \theta^C}{\partial w_{f2}}$ if

$$\frac{w_{f2}^2-d^2+2\lambda d(w_{m2}-d)}{w_{f2}^2}<1$$

$$\Rightarrow 2\lambda d(w_{m2}-d)<d^2$$

$$\Rightarrow \lambda<\frac{d}{2(w_{m2}-d)}.$$

Thus, with $\frac{\partial w_{f2}}{\partial l_{f1}}<0$, the second element on the right hand is also positive if

$$\lambda<\frac{d}{2(w_{m2}-d)}.$$

Therefore, $l_{f1}^M-l_{f1}^C>0$ if $\lambda<\frac{d}{2(w_{m2}-d)}$, and $0<d<\frac{2}{3}w_{m2}(1)$.

$l_{f1}^M-l_{f1}^C>0$ implies that $w_{f2}(t_{f1}^C)>w_{f2}(t_{f1}^M)$. Thus, the condition in (A-1) holds for the choice of labor supply under marriage and cohabitation.

Total Indirect Expected Utilities under Marriage and Cohabitation.

$$EU^M = (h_f l_{f1}^M)[w_{m1} + w_{f1}(1 - l_{f1}^M)] + 2\mu + [1 - F(\theta^M)]h_f w_{m2}(1)$$
$$+F(\theta^M)\left\{\frac{h_f}{4w_{f2}(t_{f1}^M)}[w_{f2}(t_{f1}^M) - d]^2 + \lambda\frac{h_f}{2}w_{f2}^{-1}(t_{f1}^M)[w_{f2}(t_{f1}^M) - d][w_{m2}(1) - d)]\right\}$$
$$+2\int_{\theta^M}^{\infty}\theta dF(\theta)$$

$$EU^C = (h_f l_{f1}^C)[w_{m1} + w_{f1}(1 - l_{f1}^C)] + [1 - F(\theta^C)]h_f w_{m2}(1)$$
$$+F(\theta^C)\left\{\frac{h_f}{4}w_{f2}(t_{f1}^C) + \lambda\frac{h_f}{2}w_{m2}(1)\right\} + 2\int_{\theta^C}^{\infty}\theta dF(\theta)$$

Note that the implicit relationship between first period labor supply and second period wages are written explicitly in the indirect utility functions. The difference between indirect utilities in the two family union statuses is:

$EU^M - EU^C = A + B$, where

$$A \equiv 2\mu + h_f w_{m1}(l_{f1}^M - l_{f1}^C) + h_f w_{f1}(l_{f1}^M - l_{f1}^C) + 2\int_{\theta^M}^{\theta^C}\theta dF(\theta)$$

$$B \equiv -h_f w_{f1}(l_{f1}^{M^2} - l_{f1}^{C^2})$$
$$+\left[\begin{array}{l} F(\theta^M)\left\{\frac{h_f}{4w_{f2}(t_{f1}^M)}[w_{f2}(t_{f1}^M) - d]^2 + \lambda\frac{h_f}{2}w_{f2}^{-1}(t_{f1}^M)[w_{f2}(t_{f1}^M) - d][w_{m2}(1) - d)]\right\} \\ -F(\theta^C)\left\{\frac{h_f}{4}w_{f2}(t_{f1}^C) + \lambda\frac{h_f}{2}w_{m2}(1)\right\} \end{array}\right]$$

With $0 < d < w_{i2}(0)$, it is easily verified that

$$\left\{\frac{h_f}{4w_{f2}(t_{f1}^M)}[w_{f2}(t_{f1}^M) - d]^2 + \lambda\frac{h_f}{2}w_{f2}^{-1}(t_{f1}^M)[w_{f2}(t_{f1}^M) - d][w_{m2}(1) - d)]\right\}$$
$$< \left\{\frac{h_f}{4}w_{f2}(t_{f1}^C) + \lambda\frac{h_f}{2}w_{m2}(1)\right\}$$

Also, $F(\theta^M) < F(\theta^C)$.

Hence, $B<0$.

Now, $A>0$, and $B<0$.

Therefore, if $A<B \Rightarrow EU^M - EU^C < 0$, which implies that the individual will decide to cohabit. On the other hand, if $A>B$ the individual will chose to marry since in that case $EU^M - EU^C > 0$.[29]

Proposition 7. First period female labor supply increases (i.e., work at home decreases) when cost of divorce decreases if $\lambda \leq \frac{d}{w_{m2}-2d}$, and $0<d<\frac{1}{3}w_{m2}(1)$.

Proof:

$$\frac{\partial l_{f1}^M}{\partial d} = \frac{1}{8w_{f1}w_{f2}^2}\frac{\partial w_{f2}}{\partial l_{f1}}\{2F(\theta^M)[\lambda w_{m2} - 2\lambda d - d]\}$$

$$+\frac{1}{8w_{f1}w_{f2}^2}\frac{\partial w_{f2}}{\partial l_{f1}}\left\{\frac{\partial F(\theta^M)}{\partial \theta^M}\frac{\partial \theta^M}{\partial d}[w_{f2}^2 - d^2 + 2\lambda d(w_{m2}-d)]\right\}$$

$$-\frac{1}{2w_{f1}}\frac{\partial F(\theta^M)}{\partial \theta^M}\frac{\partial \theta^M}{\partial w_{f2}}\frac{\partial w_{f2}}{\partial l_{f1}}\frac{1}{4w_{f2}}\left\{2(w_{f2}-d)+\frac{\lambda}{2w_{f2}}(w_{f2}+w_{m2}-2d)\right\}$$

It is easily observed that $[w_{f2}^2 - d^2 + 2\lambda d(w_{m2}-d)] > 0$, and that $\left\{2(w_{f2}-d)+\frac{\lambda}{2w_{f2}}(w_{f2}+w_{m2}-2d)\right\} > 0$.

Since $0<d<w_{i2}(0)$, $\frac{\partial F(\theta^M)}{\partial \theta^M} > 0$, $\frac{\partial \theta^M}{\partial w_{f2}} > 0$, $\frac{\partial w_{f2}}{\partial l_{f1}} < 0$, and $\frac{\partial \theta^M}{\partial d} < 0$, the sum of the last two elements in the above polynomial on the right hand side are positive.

The first element is also non-negative always if,

$\lambda w_{m2} - 2\lambda d - d \leq 0$

---

[29] Note that $\frac{\partial(EU^M - EU^C)}{\partial d} < 0$ which suggests that reducing the cost of divorce lowers the gap between expected utilities in marriage and cohabitation. This serves as a consistency check for the model where the positive cost of divorce is one of the primary criteria that distinguish marriage from cohabitation.

$$\Rightarrow \lambda \le \frac{d}{w_{m2} - 2d}$$

But, $0 < \lambda < 1$. Hence, we need $0 < d < \frac{1}{3} w_{m2}(1)$.

Therefore, the sufficient condition for $\frac{\partial l_{f1}^{M}}{\partial d} > 0$ is

$\lambda \le \frac{d}{w_{m2} - 2d}$ and $0 < d < \frac{1}{3} w_{m2}(1)$.

Vita

Arif A. Mamun was born in Brahmanbaria and brought up in Dhaka, Bangladesh. At the University of Dhaka, he earned a Bachelor of Social Sciences degree (with Honours) in Economics in 1995, and a Master of Social Sciences in Economics in 1996. He also earned a Master of Arts degree in Economics at the University of Washington in 2003. In 2005, he earned a Doctor of Philosophy at the University of Washington in Economics.