Gender Differences in the Marriage and Cohabitation Income Premium

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Audrey Light
Department of Economics
The Ohio State University

410 Arps Hall
1945 North High Street
Columbus, OH 43210
light.20@osu.edu

Abstract

Using data from the NLSY79, I identify causal effects of marriage and cohabitation on total family income. My goals are to compare men's and women's changes in financial status upon entering unions and to assess the relative contributions of adjustments in own-income, income pooling, and changes in family size. Changes in own-income due to intra-household specialization prove to be minor for both men and women relative to the effects of adding another adult's income to the family total. Women gain roughly 55% in needs-adjusted, total family income regardless of whether they cohabit or marry, while men's needs-adjusted income levels remain unchanged when they make these same transitions.

INTRODUCTION

In 2000, the median family income for married couples in the U.S. was $59,099, while the median income for single men and women was $37,727 and $25,716, respectively (U.S. Census Bureau 2002). Statistics such as these are often interpreted as evidence that it "pays" for both men and women to be married. Waite and Gallagher (2000: 109) convey this view in a particularly succinct fashion when they write: "Both men and women, it is fair to say, are financially better off because they marry. Men earn more and women have access to more of men's earnings."

Does marriage *really* confer financial benefits on *both* partners? From a theoretical perspective, the predicted effect of marriage on an individual's income is ambiguous. Consider the traditional behavior envisioned by Becker (1973, 1974, 1991) in which men specialize in market work while their wives specialize in home production. If the intra-household division of labor allows a man to be more productive at work, his earnings increase as a direct result of marrying. His effective income does not necessarily increase, however, because his earnings are now shared with his entire family. A married man is financially better off only if his earnings premium plus his wife's income contribution exceed increases in family need. Similarly, a woman benefits financially only if her husband's income compensates for her lost income as well as changes in family need. Becker's model predicts that both men and women gain "z-goods" produced within the household (and this gain, after all, is what motivates them to marry) but we cannot be sure that each partner's income increases.

It is straightforward to address this issue empirically by tracking men's and women's needs-adjusted family income as they transition into marriage. However, the empirical literature on income-related gains to marriage has focused on slightly different issues. A large group of studies (Cornwell and Rupert 1997; Daniel 1995a, 1995b; Gray 1997; Korenman and Neumark 1991; Loh 1996; Nakosteen and Zimmer 1987; Stratton 2002) examines the causal effect of marriage on men's wages. The consensus is that men receive a modest wage premium upon marrying, but this finding only points to the existence of intra-household specialization—it does not identify the effect of marriage on the financial status of men or women. Another set of studies (Bianchi, Subaiya and Kahn 1999; Burkhauser *et al.* 1991; Duncan and Hoffman 1985; Smock, Manning, and Gupta 1999) reverses the question posed here and identifies the effects of divorce on individuals' financial well-being. Numerous analysts examine the link between marital status and the economic well-being of women with children (Budig and England 2001; 0Lichter, Graefe and Brown 2003; McLanahan and Sandefur 1994; Spain and Bianchi 1996; Thomas and Sawhill 2002). A comprehensive analysis of the effects of union formation on both men's and women's financial well-being is missing from the literature.

In the current study, I fill this gap in the empirical literature by using 1979-2000 data from the National Longitudinal Survey of Youth to analyze changes in family income associated with transitions into first unions. My analysis has the following attributes. First, I consider income effects of both marriage and cohabitation. Given the prominence of cohabitation throughout the observation period (Bumpass and Lu 2000; Bumpass and Sweet 1989), I am interested in learning whether the two types of unions yield different financial benefits. Second, I ask whether marriage and cohabitation *cause* income to increase, or whether it is simply the case that high-income individuals form unions. I exploit within-person variation in the data to isolate true income effects of union formation from the confounding effects of unobserved, time-invariant factors. Third, after identifying overall changes in individuals' needs-adjusted family income, I decompose these changes into the portions due to (a) gaining or losing own-income, (b) adding a partner's income to the family total, and (c) increased family size. My goal is to learn how the financial benefits to union formation and the sources of those benefits differ by gender.

BACKGROUND

Economic marriage models (Becker 1973, 1974, 1991; Weiss 1997) demonstrate how marriage (and perhaps cohabitation) lead to financial gain. Individuals who meet in the marriage market are assumed to assess their combination of attributes, predict the benefit of joining forces, and marry if the expected gain represents their best alternative.[1] While the gain to marriage can span many dimensions, economic models highlight the portion derived from the consumption of commodities produced within the household. Married couples receive consumption-related gains because they jointly consume public goods, pool risk, extend credit to one another, and/or engage in intra-household specialization that enables more goods to be produced—typically, by having the man specialize in market work while the woman concentrates on home production.

The magnitudes of these economic gains to marriage are intrinsically tied to market conditions. For example, the gain to specialization is expected to be positively correlated with the gap between the man's and woman's labor market skills, while the gain to consuming public goods is expected to increase with total family income. Both predictions gained relevance as technological change and other factors caused women's potential earnings to increase in the late 20th century. Put simply, the marriage market no longer consists of "breadwinning" males and "homemaking" females. In light of this trend, theoretical attention has turned to the potential effects of women's increased employment and men's declining labor market prospects on union-forming decisions (Becker 1991; Cherlin 1980; Mulligan and Rubinstein 2002; Oppenheimer 1988, 1994, 1997).

[1]Becker's (1973, 1974) original formulation assumes perfect information. Search-theoretic marriage models assume the decision is made in an environment with imperfect information. See Pollack (2000) and Weiss (1997) for a comparison of these two approaches.

An extensive literature is dedicated to confronting the models' predictions with the data.[2] I will not attempt a comprehensive survey, but it is worth noting that three distinct, empirical approaches dominate the literature. Assortative mating studies look directly at partner choice, typically by examining correlations among partners' attributes. These studies address such issues as the effect of market conditions on partner choice (Lewis and Oppenheimer 2000; Mare 1991; Qian and Preston 1993) and mating differences between married and cohabiting couples (Blackwell and Lichter 2000; Jepsen and Jepsen 2002; Schoen and Weinick 1993). Another class of research asks how union-forming decisions are affected by expected economic gains. In this approach—which can use either aggregate or individual-level data— marital states are the outcomes of interest and proxies for expected gains to marriage or cohabitation are the key covariates. Reduced-form choice models of this type include Brien (1997), Lichter, McLaughlin, and Ribar (2002), Smock and Manning (1997), and Xie *et al.* (2003). A third type of study reverses the causality and models realized gains (*e.g.,* men's wages) as a function of marital status and other covariates. My investigation belongs to this third class of empirical studies, for I ask how individuals' log-income paths are affected by changes in marital status. In the remainder of this section I focus on the "gains to marriage" literature.

No potential gain to marriage receives more empirical scrutiny than men's wages, for analyses of the marriage-wage link provide a direct test of the economic marriage model. If specialization occurs within marriages, then married men should invest more intensively than their single counterparts in marketable skills and subsequently receive more wage growth. Estimates of the productivity-enhancing effect of marriage on men's wages provide evidence that gains to specialization exist. However, the identification of this effect is nontrivial. While cross-sectional comparisons invariably show that married men have higher earnings than nonmarried men, this finding does not necessarily reflect the gain to specialization. An alternate explanation for the correlation is that men with relatively high levels of labor market productivity are more likely than others to marry.

The standard strategy for distinguishing between selection and the causal, productivity-enhancing effects of marriage is to specify wage models that account for the endogeneity of marital status, often by assuming marriage decisions are driven by unobserved, fixed effects. Most research in this vein concludes that small productivity effects remain after selection effects are eliminated (Daniel 1995a, 1995b; Gray 1997; Korenman and Neumark 1991; Loh 1996; Stratton 2002). Specifically, these studies find that men's wage growth increases after marriage, which is consistent with married men investing more intensively than others in marketable

[2]While Becker's original model (1973, 1974) does not formally distinguish between marriage and cohabitation (see Moffitt (2000) and Weiss (1997) on this point), one mission of empirical analysts is to determine whether cohabitation and marriage decisions differ.

skills.[3]

In principle, the gains to specialization extend to any household where adults team together to increase joint consumption. However, cohabiting men are likely to receive a smaller wage premium than married men if cohabitors have inherently lower levels of trust, commitment, and expected union duration that make them less willing to undertake relationship-specific investments. It is well established that union duration (whether anticipated by the couple or not) is lower, on average, for cohabitors than for married couples (Bumpass and Sweet 1989). South and Spitze (1994) report a smaller male-female differential in weekly hours of housework among cohabitors than among married counterparts. This evidence suggests that intra-household specialization and, in turn, the boost to men's wages may be less pronounced for cohabitors than for married couples. This prediction is supported by Daniel (1995a) and Stratton (2002), who are among the few analysts to assess the causal effects of cohabitation on men's wages.

The assortative mating literature offers an alternative empirical strategy for identifying intra-household specialization. Becker (1973, 1974, 1991) argues that husbands' and wives' potential wages should be negatively correlated if specialization occurs within marriages. The negative correlation arises from their optimal sorting decisions (*e.g.,* a man with high potential wages should seek out a wife with low market productivity) as well as from skill investments made during the marriage. However, Lam (1988) demonstrates that this prediction need not hold if the joint consumption of public goods is included among the gains to marriage. The gain to joint consumption is greater if the partners have similar demands for public goods, and the resulting incentive to pair with similar-skilled individuals may offset the specialization-driven incentive to sort negatively on market skills. In fact, empirical analysts consistently find positive correlations among married couples' wages (Jepsen and Jepsen 2002; Nakosteen and Zimmer 2001; Smith 1979; Suen and Lui 1999); Jepsen and Jepsen (2002) find that cohabiting couples' wages are positively correlated as well.[4]

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[3] This majority view has its detractors: Cornwell and Rupert (1999) and Nakosteen and Zimmer (1987) find that self-selection is the primary source of the observed marriage premium.

[4] By examining correlations in schooling attainment rather than wages, Schoen and Weinick (1993) purport to find evidence that cohabitors specialize less than married couples. They show that cohabiting couples are more educationally homogamous than married couples, and argue that this means each partner makes a relatively equal contribution to family income. They do not test this prediction, which contradicts Becker's (1973) assertion that educational homogamy goes

There is ample evidence that intra-household specialization is becoming less pronounced across successive cohorts of married couples. Blackburn and Korenman (1994), Cohen (2002) and Gray (1997) report that the male marriage premium is declining over time. Numerous analysts document the dramatic increases in married women's labor market activities during the 1970s and 1980s (*e.g.*, Blau 1998; Blau, Ferber and Winkler 1998; Goldin 1989; Spain and Bianchi 1996). Both men and women have changed their work effort to the point where husbands' and wives' earnings are becoming increasingly positively correlated over time (Cancian, Danziger and Gottschalk 1993; Juhn and Murphy 1997). Using cross-sectional data for 1993, Winkler (1998) finds that wives' annual earnings exceed their husbands' in 20% of dual-earner families and account for an average of 35% of the family total. While husbands and wives tend to contribute more equally to family income than in earlier eras, specialization continues to be evident among couples with young children. Researchers who control for the endogeneity of marital status and children generally find negative effects of children on mothers' work effort and wages (Angrist and Evans 1998; Korenman and Neumark 1994; Lundberg and Rose 1998; Waldfogel 1997). Lundberg and Rose (2000) find corresponding increases in fathers' wages and work effort.

One lesson to be learned from these various studies is that we cannot fully understand the gains to union formation by focusing on intra-household specialization. Marriage and cohabitation decisions are driven by the partners' *total* expected gains, which are not limited to gains to specialization. As market forces lead couples to specialize less and collaborate more with respect to labor market activities, men are increasingly likely to benefit financially from their partners' income contributions. With this lesson in mind, I focus on the "overall" income gain that men and women receive upon forming a union, and identify the contributions of partners' income and changes in own-income to the overall gain.

Most studies that explore the link between marital status and family income focus on the well-being of women. One group of studies demonstrates that single mothers are much more likely than married mothers to live in poverty (Lichter *et al*. 2003; McLanahan and Sandefur 1994; Spain and Bianchi 1996; Thomas and Sawhill 2002). These studies provide cross-sectional evidence that women benefit financially from marriage, but relatively little attention is paid to the distinction between selection effects (*e.g.*, the fact that high-income women are more likely to attract a partner) and causal effects of marriage.

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hand in hand with specialization insofar as highly schooled women are relatively more productive in the home. (Benham (1974) suggests that highly schooled women are better able to augment their husbands' productivity.) Blackwell and Lichter (2000) find that cohabitors are slightly *less* educationally homogamous than married couples.

Another group of studies examines the relationship between divorce and financial well-being. Using panel data, Bianchi *et al*. (1999) find that the median income-to-needs ratio for women is barely one-half that of men in the first year after married couples separate. Burkhauser *et al.* (1991) report that the unconditional, median loss in total income associated with divorce is 24% for women and only 6% for men. At the same time, numerous researchers (*e.g.*, Duncan and Hoffman 1985, Smock *et al.* 1999) demonstrate that remarriage goes a long way toward restoring women's economic well-being. In one of the few studies that controls for self-selection into marriage, Smock *et al.* (1999) predict the total family income for remarried women to be more than twice the income level they would attain if they remained divorced, although not as high as the income level of women who never divorce. There appears to be little doubt that women benefit financially from being married, but additional evidence is needed on men's financial benefits and the distinction between cohabitation and marriage.

ANALYTICAL FRAMEWORK

I begin with the following model:

$$\ln Y_{it} = \alpha + \alpha_s S_{it} + \alpha_c C_{it} + \beta A_{it} + \varphi_i + \varepsilon_{it} \tag{1}$$

where $\ln Y_{it}$ is the natural logarithm of income for individual i at time t, S_{it} and C_{it} are dummy variables indicating whether the individual is single or cohabiting at time t (with married the omitted category), and A_{it} is the individual's age. Time-constant, unobserved factors that explain variation in $\ln Y_{it}$ are represented by φ_i, while ε_{it} represents time-varying unobservables. As written, model (1) assumes log-income paths evolve linearly as individuals age and shift up or down according to changes in marital status only. In estimating each income model, I include a quartic function of age and a host of additional demographic and environmental shift factors (race, presence of children, calendar year, *etc*.); my goal is to use a flexible parameterization to minimize the chance that the estimated marital status coefficients reflect the effects of omitted variables.

Ordinary least squares (OLS) estimates of model (1) identify differences in predicted log-income between individuals who are single and cohabiting ($\hat{\alpha}_c - \hat{\alpha}_s$), cohabiting and married ($-\hat{\alpha}_c$) and single and married ($-\hat{\alpha}_s$). However, these estimates have at least two shortcomings. First, OLS does not identify the *causal* effect of changes in marital status on log-income. If unobserved factors subsumed in φ_i and ε_{it} affect individuals' marriage and cohabitation decisions, then OLS estimates confound the value-added of a change in status with the independent, income-enhancing or income-detracting effects of these unobservables. Second, model (1) constrains the slope of the predicted age-income path to be independent of marital status. A change in marital status is assumed to cause a once-and-for-all change in log-income, with no effect on subsequent income growth. This restrictive assumption is inconsistent with evidence that men's wages—which typically account for a large share of family income—increase more rapidly among married men than among nonmarried men (Korenman and

Neumark 1991, Stratton 2002).

I relax the slope restrictions implicit in model (1) with the following, more flexible specification:

$$\ln Y_{it} = \alpha + \alpha_s S_{it} + \alpha_c C_{it} + \gamma_s D_{it}^s + \gamma_c D_{it}^c + \gamma_m D_{it}^m + \beta A_{it} + \varphi_i + \varepsilon_{it} \qquad (2)$$

where D_{it}^s, D_{it}^c and D_{it}^m represent the duration at time t of single, cohabitation, and marriage spells, interacted with corresponding marital status indicators. Model (2) allows log-income paths to vary in slope as well as levels across marital status categories. The predicted, contemporaneous change in log-income associated with a transition from cohabitation to marriage (for example) is $-\hat{\alpha}_c - \hat{\gamma}_c \overline{D}^c$, where $\overline{D}^c$ is the completed duration of the cohabitation spell; this predicted gap continues to grow or shrink by $\hat{\gamma}_m D^m$ as the marriage spell evolves. (An even more flexible specification that allows each state-specific slope to be nonlinear in duration proves to be unwarranted by the data.) Model (1) predicts this same gap to be a uniform $-\hat{\alpha}_c$ regardless of the duration of each spell.

I address the endogeneity issue by differencing the data and using OLS to estimate the model

$$\Delta \ln Y_{it} = \alpha_s \Delta S_{it} + \alpha_c \Delta C_{it} + \gamma_s \Delta D_{it}^s + \gamma_c \Delta D_{it}^c + \gamma_m \Delta D_{it}^m + \beta \Delta A_{it} + \Delta \varepsilon_{it} \qquad (3)$$

where $\Delta \ln Y_{it} = \ln Y_{it} - \ln Y_{it-\tau}$, $\Delta S_{it} = S_{it} - S_{it-\tau}$, *etc.* This transformation of the data eliminates φ_i from the residual and leaves only within-person variation with which to identify the parameters of interest. As long as individuals' decisions to cohabit, marry, and remain single are driven by time-constant unobservables only, the difference estimators for $\alpha_s, \alpha_c, \gamma_s, \gamma_c,$ and γ_m are free from the endogeneity bias inherent in OLS estimators and can be interpreted as causal effects. To test the assumption that S_{it} and C_{it} are exogenous conditional on the fixed effect, I use a test proposed by Heckman and Hotz (1989) and Wooldridge (2002). I estimate a version of (3) in which lead values of the marital status controls ($S_{it+\tau}$ and $C_{it+\tau}$) are included among the regressors. If the estimated coefficients for these lead values are statistically significant, it must be due to correlation between the regressors and the (differenced) time-varying residual. If these estimated coefficients are jointly insignificant, I can conclude that the exogeneity assumption is valid.

In light of the evidence discussed in the preceding section, I expect the effects of marital status transitions on income paths to differ for men and women. However, there are exceptions. When cohabiting couples marry, changes in *family* income are identical for both adult members of the family. More generally, both members of cohabiting and married couples experience identical growth in family income throughout the duration of the union. I estimate models (1) and (2) for a pooled sample of men and women, and interact the regressors with a gender identifier to allow $\alpha_s, \alpha_c, \gamma_s, \gamma_c,$ and γ_m to differ for men and women. In specifications where the dependent variable is total family income, I constrain α_c, γ_c and γ_m to be equal for men and women.[5] With the exception of coefficients for year dummies and the error variances, all other parameters in each model are allowed to vary with gender.

I compute both OLS and difference estimators for the parameters in models (1) and (2). Because each individual contributes multiple observations to the sample, I compute robust standard errors that account for nonindependence of observations within person-specific clusters.

DATA

Sample Selection

The data are from the 1979 National Longitudinal Survey of Youth (NLSY79). The NLSY79 began in 1979 with a sample of 12,686 men and women born in 1957-1964. The sample contains 6,111 individuals who form a representative sample of the civilian, U.S. population in the targeted birth years, an over-sample of 5,295 blacks, Hispanics, and economically disadvantaged whites, and a sample of 1,280 individuals who served in the military prior to the start of the survey. Respondents were interviewed annually from 1979 to 1994, and biennially thereafter; I use data for all interview years from 1979 to 2000.

My strategy for constructing a sample of person-year observations is dictated by the manner in which income is reported in the NLSY79. During every interview, respondents detail their annual income (by source) for the preceding calendar year. Respondents who are currently married or cohabiting report their spouse's or partner's income as well. If a respondent cohabits with the same partner throughout 1990 but ends the relationship prior to the 1991 interview, for example, the income he reports that year does not reflect his partner's contribution to 1990 total family income. Conversely, if the respondent cohabits with a single partner from October 1990 onward, in 1991 he reports his partner's 1990 income despite spending a relatively small portion of that year as a cohabiting couple.

To ensure that respondents' reported family income matches their marital status in the preceding year, I proceed as follows. First, I use information on marital status, starting dates for marriage and cohabitation spells, and partner/spouse identifiers to determine when each partner-specific cohabitation and marriage begins and ends. Second, I classify a respondent's status during each calendar year as cohabiting or married if he spends at least 10 months with a *single* partner, and as single if he spends at least 10 months without a partner. Remaining cases are classified as "mixed." Third, I determine the respondent's total family income (own income plus the spouse's or partner's, if applicable) for each calendar year using information reported

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[5] A small number of cohabiting-to-married transitions in the data are accompanied by partner changes, so in principle I can identify different values of α_c for men and women. In each specification, I fail to reject the null hypothesis of equality of coefficients using a 5% significance level.

during the next year's interview; married and cohabiting respondents must still be living with the *same* spouse or partner for this information to be available. A given person-year remains in the sample if (a) the respondent is classified as single, cohabiting, or married, (b) his own income for the year and his spouse's or partner's (if applicable) are reported to be between $100 and $1 million, (c) his 20th birthday precedes the end of the calendar year, and (d) he has not yet dissolved his first marriage.[6]

I use person-year observations that are two years apart to compute $\Delta \ln Y_{it}, \Delta S_{it},$ *etc.* for the differenced versions of models (1) and (2). That is, I use lag-two differences of annual person-year observations for interview years 1979-1994 and lag-one differences of biannual observations for interview years 1996-2000. This strategy has two advantages, in that I handle the data uniformly throughout the period of observation, and I skip the "mixed" years when respondents change their marital status. My sample consists of 41,078 differenced person-year observations for 9,839 individuals. To estimate the levels (OLS) models, I "undifference" the data to obtain 50,917 observations for the same 9,839 individuals.

An unavoidable consequence of using annual income data—and especially of using two-year differences—is that short spells are excluded from the sample. This causes cohabitation spells to be under-represented in my sample because they tend to be shorter than marriages (Bumpass and Sweet 1989). If individuals who experience short cohabitation spells and marriages differ from others only in their time-invariant, personal characteristics (*e.g.*, if they have lower levels of commitment), then the omission of their spells does not pose a problem because I rely on within-person variation in the data. However, if the structural relationship between marital status and family income differs systematically with completed spell durations—*e.g.,* if unions in which one partner contributes relatively little to family income are less likely than others to last—then the parameters I estimate do not necessarily reflect the relationships that prevail for the overall population.

Measuring Income and Adult Equivalents

I use three alternative income measures in estimating models (1) and (2). Total family income (Y^F) is the sum of earnings (military income, wages, salary, tips, farm and business income), unemployment benefits, public assistance (AFDC, food stamps, SSI), educational benefits, veteran's benefits, child support, and "other income" received by the individual or couple during

[6] I eliminate about 1% of observations by imposing $100 and $1 million cut-offs. A number of implausibly large income measures (on the order of $10 million) are reported by cohabitors, and these outliers cause a significant increase in the estimated gain to cohabitation if left in the sample.

the calendar year. The definition of own income (Y^o) is less clear-cut because certain sources of income (veteran's benefits, each form of public assistance, child support, and "other income") are not reported separately for individuals and their spouses. I define own income to be the sum of the respondent's earnings, unemployment benefits, educational benefits, and these shared sources of income. The estimates reported in the next section are insensitive to whether Y^o includes or excludes "shared" income.

The third dependent variable is total family income per adult equivalents (Y^F / AE), or effective income. The value that a given income affords different families depends on the presence of children and the extent to which they exploit scale economies.[7] To account for these cross-family differences, I use an adult equivalence scale proposed by Fuchs (1986a, 1986b). The number of adult equivalents is a weighted sum of the number of adults and children in the family, where the first adult is given a weight of one, the second adult's weight is 0.8, the first child's weight is 0.4, and all subsequent children are weighted 0.3.

I also consider two alternative controls for family size. One is a measure proposed by Citro and Michael (1995) in which adult equivalents are defined as $(A + 0.75C)^{0.75}$, where A and C are the number of adults and children in the family. The measure $(A+aC)^b$ is widely used in the literature (*e.g.,* Cutler and Katz 1992; Deaton and Paxson 1998). Citro and Michael (1995) propose the values $a=b=0.75$ as part of a new poverty measure recommended by a panel of the National Research Council Committee on National Statistics; hence, I refer to the measure as the NRC scale. A second adult adds 0.68 in adult equivalents to a family with no children and 0.57 if two children are present; the first and second child add 0.46 and 0.42, respectively, assuming two adults are present. The NRC measure always puts less weight on adults than does Fuchs's scale, and it puts more weight on additional children for families with no more than two adults and up to 14 children. My second alternative to Fuchs's scale defines adult equivalents as the total number of individuals in the family ($A+C$). By putting a weight of one on each family member, this per capita measure accounts for neither economies of scale nor age-specific differences in consumption.

Two additional comments are in order regarding the definition of family income per adult equivalent. First, because the numerator (Y^F) is restricted to income earned by the respondent and his/her spouse or partner, I include only those two adults and their children in the measures

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[7] These factors are unlikely to depend on whether a man and woman cohabit or marry, for they invariably pay for only one dwelling, one refrigerator, *etc.,* regardless of marital status. I am unaware of evidence suggesting that different adult equivalence scales should be used for cohabiting and married couples.

of adult equivalents. I wish to focus on changes in financial status associated with gaining a partner (and possibly children), and not on changes caused by additional reconfigurations of household composition.

Second, in order to use family income to assess *individual* well-being, it is necessary to account for the within-family allocation of income. By using total income per adult equivalent as a dependent variable, I assume couples share their needs-adjusted family income equally. This assumption is appropriate if family members jointly maximize a common utility function, as in Becker's (1973, 1991) decision-making framework. However, this aspect of Becker's model receives little empirical support. One prediction of the "unitary" model is that allocation decisions are affected by total family income, and not by individual income—but this "income pooling" hypothesis is rejected for samples of married couples (Browning *et al.* 1994; Lundberg, Pollak and Wales 1997; Phipps and Burton 1998; Thomas 1990) as well as cohabiting couples (Winker 1997). The data instead appear to support bargaining models (Manser and Brown 1980; McElroy and Horney 1981) and the efficient, collective decision-making framework proposed by Browning *et al*. (1994). While bargaining is widely acknowledged to occur within families, analysts have yet to identify household allocation parameters that can be used to improve upon the equal division rule implicit in "per capita" or "per adult equivalent" income measures. Virtually all analysts assume equal division because it is a useful benchmark, but it is unlikely to characterize the behavior of married or cohabiting couples.[8]

Table 1 contains sample means and standard errors for the three dependent variables. Each income measure is expressed in thousands of dollars and deflated by the GDP implicit price deflator, with 1996 as the base year. The mean level of log-income is the same regardless of whether I use Fuchs's scale or the NRC measure to adjust for adult equivalents.

Explanatory Variables

Table 1 also reports summary statistics for the explanatory variables. I control for the respondent's age (and higher-order terms), dummy variables indicating whether the respondent is black or Hispanic (with nonblack, non-Hispanic the omitted group), the respondent's highest grade completed, a dummy variable indicating whether the respondent was enrolled in school during the calendar year, and the number of weeks he/she worked during the year. I also include dummy variables indicating whether children are present in the household, whether the

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[8] An alternative income measure is $A \cdot Y^o / AE$, which is consistent with each adult consuming his *own* income adjusted for his *share* of adult equivalents (*e.g.,* Fuchs 1986a, 1986b). Because own income proves to change relatively little when individuals marry and cohabit, changes in this "no sharing" measure are driven almost entirely by the addition of children to the household.

respondent lives with his/her parents, and the calendar year (with 1985 the omitted year).

The covariates related to marital status include dummy variables indicating whether the respondent is single or cohabiting, with married the omitted group. Married respondents are identified via marital status questions asked in every NLSY79 interview. In the 1979-81 interviews, cohabitors are identified indirectly from the household roster: any respondent who lives with one unrelated, opposite-sex adult is assumed to be cohabiting. In 1982-86, respondents with this living arrangement who report themselves to be unmarried are asked directly whether they live with a partner. From 1987 onward, cohabitation status is asked of all unmarried respondents as a follow-up to the questions on marital status.[9] The remaining marriage-related covariates are the current durations of each single, cohabiting, and marriage spell. I measure the duration of single spells from the respondent's 20th birthday onward. Starting dates of cohabitation spells are not reported prior to 1990, so I assume early cohabitation spells begin midway between interview dates.

RESULTS

Descriptive Statistics

Table 2 provides preliminary evidence on the relationship between annual income and marital status. Married respondents in my sample have a mean family income of $53,437, which is almost three times as much as the mean income for single individuals and 20% more than the mean income for cohabitors. Adjusting for adult equivalents reduces the mean family income among married respondents to $25,318—a change that is consistent with a typical family having two adults and one child—and reduces the unconditional income premium associated with marriage. With family size taken into account, married respondents' mean, family income exceeds the corresponding means for single and cohabiting respondents by only 46% and 15%, respectively. One goal of the ensuing analysis is to determine how much of this income differential remains after I control for observed and unobserved factors that are correlated with marital status.

Table 2 also reveals that the average, individual income among married men is $34,436. This figure exceeds the mean income levels for single and cohabiting men by 80% and 50%, respectively. At the same time, the mean income for married women is only 12% higher than the mean for single women, and 3% *lower* than the mean for cohabiting women. Even though these statistics do not identify causal effects of marriage and cohabitation on income, they highlight

[9] These changes in measurement do not affect the estimated effect of cohabitation on log-income. I estimate a version of each income model in which the coefficient for the cohabiting dummy is allowed to vary over the three time periods (1979-81, 1982-86, and 1987-2000). In all cases, I fail to reject the null hypothesis that the three coefficients are equal.

the fact that any productivity gains realized by men are likely to come at the expense of their partners' income. Another goal of the multivariate analysis is to learn whether changes in own-income constitute significant gains to marriage and cohabiation, and how these gains compare to the benefits of pooling two adults' incomes.

Multivariate Results

I first estimate models 1 and 2 using the log of family income per adult equivalent (Y^F/AE) as the dependent variable, where adult equivalent is defined by Fuchs's scale. Table 3 reports the predicted changes in log-income associated with single-to-cohabiting, single-to-marriage, and cohabiting-to-marriage transitions. Additional estimates for each model are in the appendix.

The right-most column of table 3 contains my preferred estimates of the effects of marital transitions on the financial status of men and women. These difference estimates are unaffected by the correlation between marital status and unobserved, fixed effects. In addition, they are based on model 2, which allows marital status to affect both the level and slope of the log-income path. I predict that single women gain 0.440 in log family income when they cohabit, and 0.416 when they marry.[10] The difference between these two estimates is statistically indistinguishable from zero at a 5% significance level. I predict that single men have the same total family income per adult equivalent regardless of whether they are single, cohabiting, or married. Based on these estimates, I conclude that marriage and cohabitation confer sizeable—and identical—financial benefits on women, while men "break even" upon entering either type of union.

Before exploring the sources of these predicted changes in financial status, I return to two specification issues raised earlier. First, the evidence in table 3 indicates that the estimated gain to marriage is overstated when unobserved, fixed effects are not taken into account. Using OLS estimates for model 2, I predict that single women gain 0.437 in log family income when they cohabit and 0.542 when they marry; the corresponding figures for men are 0.116 and 0.220. The change in predicted log-income associated with cohabitation-to-marriage transitions—which I constrain to be equal for men and women—is 0.108.[11] If these estimated effects could be

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[10] To compute predicted changes in log-income for model 2, I assume individuals spent three years in the previous state (single or cohabiting) and have been in their current state for one year.

[11] Because models 1 and 2 use two parameters to estimate three transitions, the predicted effect of single-to-cohabitation transitions is constrained to be the difference between the predicted changes for the other two transitions. I test this over-identifying restriction in the differenced version of each model by adding a parameter (α_{sc}) that identifies change in log-wages associated

considered causal, I would conclude that marriage leads to an income premium that is 10 percentage points larger than the substantial income gain associated with cohabitation. However, this additional marriage premium disappears once I difference the data because it is due to unobserved, time-invariant factors that vary systematically with marital status.

The difference estimates represent causal effects only if individuals' marital status decisions are also independent of time-varying unobservables that influence log-income. To test the assumption that the dummy variables S_{it} and C_{it} are exogenous conditional on the fixed effects, I use the strategy described earlier—that is, I reestimate the differenced version of model 2 after adding dummy variables indicating whether the individual is single or cohabiting in each of the three succeeding calendar years. These dummy variables are expected to have statistically significant coefficients only if they are correlated with time-varying components of the residual. I test the null hypothesis that the coefficients for these lead values are jointly zero; the p-values are 0.902 for men and 0.278 for women. This is evidence that, in fact, the difference estimates are free of endogeneity bias.

The second specification issue concerns the contrast between models 1 and 2. Model 1 constrains the slope of the log-income path to be invariant to marital status. Table 3 indicates that this restriction has no effect on the predicted log-income gaps for men, but reduces women's predicted gains to cohabiting or marrying by five log-points (although this prediction depends on assumed union durations).[12] Putting aside the implied gender difference for the moment, the bottom line is that *income* gains to marriage and cohabitation change very little with union duration. This is in direct contrast to the finding that men's *wages* grow more rapidly after marriage (Korenman and Neumark 1991; Loh 1996; Stratton 2002). As I demonstrate below, changes in needs-adjusted family income are dominated by the addition of partners' income and changes in family size. Both factors lead to large, contemporaneous effects on log-income levels that dwarf any subsequent changes. Moreover, with (annual) family income as the dependent variable, the gains to marriage and cohabitation also reflect changes in labor supply related to marital status. This dimension explains the gender difference revealed by model 2. The

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with single-to-cohabitation transitions. For every specification presented in this paper, I fail to reject the null hypothesis $\alpha_{sc} = \alpha_c - \alpha_s$ at a 5% significance level; the smallest p-value I obtain is 0.119.

[12] I fail to reject the null hypothesis that the slope of the log-income path is identical for single, cohabiting, and married men $(\gamma_s = \gamma_c = \gamma_m)$; the p-value is 0.654 for the OLS estimates and 0.252 for the difference estimates. For women, the corresponding p-values are both less than 0.005.

parameter estimates in the appendix table show that single women receive more income growth than married women. After examining the components of annual income more closely, I attribute most of this differential growth to greater work effort among single women, and not to differences in average hourly wages and nonlabor income. This finding is consistent with evidence that women reduce their work effort upon marrying and, especially, upon having children (Angrist and Evans 1998; Lundberg and Rose 2000; Waldfogel 1997).

Next, I assess the contribution of intra-household specialization to the financial gains associated with union formation. To isolate the effect of specialization, I reestimate model 2 using respondents' own income, unadjusted for adult equivalents, as the dependent variable. The difference estimates in table 4 reveal that women's predicted own-income declines by 8% when they move from cohabitation to marriage and by 13% when they move directly from single to marriage. When men make these same transitions, their income is predicted to increase by 3-4%, although these gains are imprecisely estimated. (Because own income is not shared, I no longer constrain the estimated cohabitation-to-marriage effects to be identical for men and women.) Both men and women see their own-income levels change by negligible, statistically insignificant amounts when they form cohabiting unions.

I draw three conclusions from the "own income" estimates in table 4. First, married couples continue to choose a division of labor that augments men's labor market earnings but detracts from women's. Second, specialization is more pronounced among married couples than among cohabiting couples. These findings are consistent with evidence that men receive a small wage premium upon marrying (Daniel 1995a, 1995b; Gray 1997; Korenman and Neumark 1991; Loh 1996; Stratton 2002) that exceeds the premium associated with cohabiting (Daniel 1995a, Stratton 2002).[13] Finally, I conclude that gains to specialization are a minor part of men's and especially women's overall income gains. A single woman is predicted to gain 52% (exp(0.416)-1) in effective income upon marrying, despite the fact that her own income falls by 13%. The fact that intra-household specialization allows her spouse to earn 3% more than he otherwise would is not the driving force behind her financial gain.

While specialization is not a major source of the income gain associated with union

[13] To my knowledge, only Daniel (1995a, 1995b) provides parallel evidence on the causal effect of marriage and cohabitation on women's wages. (As noted earlier, several studies identify causal effects of *motherhood* on women's wage.) Daniel reports that white women's wages are unaffected by marital status while black women receive a small wage premium upon marrying (but not cohabiting). Because the author provides few details about his data and model specification, it is difficult to reconcile his findings with mine.

formation, gaining access to another adult's income is important for both men and women. To demonstrate this, I reestimate model 2 using the log of total family income *unadjusted* for adult equivalents as the dependent variable. Each predicted change in needs-adjusted income shown in table 3 reflects a change in the log of total family income ($\log(Y^F)$) net of a corresponding change in the log of adult equivalents ($\log(AE)$). By comparing predictions that use $\log(Y^F/AE)$ and $\log(Y^F)$ as dependent variables, I can decompose the former into its components.

The right-hand columns of table 4 show predicted changes in the log of total family income associated with each marital status transition. The predictions reveal that both men and women receive substantial gains in total family income when they form unions. Focusing on the difference estimates in table 4, the predicted effect of single-to-cohabitation transitions on log-income is 0.959 for women and 0.532 for men. The predicted changes associated with single-to-marriage transitions are only slightly smaller (0.934 for women and 0.506 for men), while the predicted change associated with cohabitation-to-marriage changes is not significantly different than zero. These gains are largely due to the addition of a second adult's income to the family total. They are consistent with the fact that "breadwinner/homemaker" specialization has given way to increasingly collaborative unions in which women make significant contributions to family income. The predicted changes in log-income shown in table 4 correspond to a female-male income ratio of 0.60-0.70, which is close to the gender ratio in wages and earnings reported elsewhere (*e.g.*, Blau 1998).

Table 5 contains transformations of the same predicted changes shown tables 3-4. The top row gives the predicted changes in total family income. For example, women making single-to-cohabitation transitions are predicted to gain 0.959 in log-income, and table 5 shows this as a 161% change in income (exp(0.959)-1). The second row of table 5 reports predicted percent changes in family income per adult equivalent. For example, the predicted change of 0.440 for women making single-to-cohabitation transitions (table 3) appears in table 5 as a 55% increase in income (exp(0.440)-1). The third row shows predicted percent changes in adult equivalents. If women gain 0.959 in total log-income and 0.440 in needs-adjusted log income when they form cohabiting unions, their predicted change in log(AE) is 0.519 (0.959-0.440). This appears in table 5 as a 68% (exp(0.519)-1) change in adult equivalents. The bottom row of table 5 shows predicted changes in own-income, based on the estimates in table 4.

By comparing all four rows of estimates in table 5, I can offer the following conclusions. First, productivity enhancement accounts for an insubstantial portion of the gains to marriage. When single individuals marry, women's predicted income levels decrease by 13% while men's increase by 4%. Added together, a given male-female pair may or may not earn more income as a married couple than they would as single individuals—but any productivity-related gain pales in comparison to the expected gain in total income of 154% for women and 66% for men. The contrast is even starker for single-to-cohabitation transitions: women and men are predicted to gain 161% and 70% in total income, respectively, despite the fact that neither partner can expect his or her own income to change. Second, because the lion's share of the income premium comes from gaining another adult's income, women typically benefit far more than men. Women receive a predicted 52-55% gain in family income even after increased family size is taken into account, whereas men's increase in predicted family income is exactly offset by gains in family size. Third, because the estimated effects represent gains to "coupling" rather than gains to intra-household specialization, cohabiting couples receive at least as large an income boost as married couples. Married and cohabiting couples may differ in other dimensions, but my analysis indicates that men's and women's expected changes in financial status do not depend on the type of partnership they form.

Robustness Checks

I have found that women gain slightly more than 50% in needs-adjusted family income when they cohabit or marry, while men neither gain nor lose effective income. To assess the sensitivity of these results to the measure of adult equivalents used, I reestimate the difference version of model 2 with $\log(Y^F / AE)$ as the dependent variable, but I consider the two alternative measures of adult equivalents defined earlier: the NRC measure, and the per capita measure. The left-hand columns of table 6 show the estimated effects of marital status transitions on these alternative measures of log-income per adult equivalent. In table 7, I use these estimates to reproduce a portion of the decomposition shown in table 5.

Using the NRC measure, I predict a 66-70% increase in needs-adjusted income when single women move into cohabitation or marriage, and a 5-7% increase when single men do so. These numbers are slightly larger than the corresponding estimates (52-55% for women and roughly zero for men) obtained with Fuchs's scale. The NRC measure puts less weight on the second adult and more weight on each child than does the Fuchs measure. Because the former weight dominates (*i.e.*, the addition of one adult is the modal, observed change in family size at the time of union formation), use of the NRC measure leads me to predict that both men and women gain needs-adjusted income upon forming a union. When I switch to per capita family income, the predicted changes in needs-adjusted income are smaller than the corresponding estimates based on Fuchs's scale: table 7 shows that women gain only 47-50% in adjusted income when they cohabit or marry, while men lose 7-9%. By weighting adults' and children's' needs equally and not allowing for scale economies, the per capita measure understates gains in needs-adjusted income associated with gaining a partner. I cannot assess the merits of alternative measures of family need in the absence of detailed consumption data, but it appears that reasonable alternatives to Fuchs's measure have relatively small effects on the estimates.

Previous research points to significant racial differences in the wage effects of marriage and cohabitation (Daniel 1995a, 1995b). Thus far in my analysis, I have constrained the income effects of marital transitions to be uniform across race groups. To determine whether my estimates mask heterogeneity across races in the gains to cohabitation and marriage, I reestimate the differenced version of model 2 for separate samples of black and nonblack, non-Hispanic ("white") respondents.[14] I use $\log(Y^F / AE)$ as the dependent variable and Fuchs's measure of adult equivalents. Difference estimates for each race-specific sample appear in the right-hand columns of table 6.

The estimates in table 6 indicate that blacks and whites receive virtually the same income premium when they cohabit and marry. The largest black-white difference is seen among men who transition from single to married: white men are predicted to lose 0.026 in log-income, while black men are predicted to lose 0.101. However, neither coefficient is precisely estimated and, in fact, each race difference shown in table 6 is statistically indistinguishable from zero at conventional significance levels. Other analysts suggest that blacks and whites receive different gains in *own* income upon forming unions, presumably because of race differences in intra-household specialization. My estimates indicate that expected changes in *total* income—which are primarily due to the addition of a partner's income—are similar for the average black and white individual.

CONCLUDING COMMENTS

My investigation reveals that union formation confers sizeable financial benefits on women, but not on men. A portion of this conclusion is "old hat," for evidence abounds that women gain financially upon marrying. What my investigation highlights, however, is that the average woman can expect to receive a virtually *identical* income premium regardless of whether she cohabits or marries. Marriage appears to be more beneficial than cohabitation when self-selection is ignored, but causal effects of marriage and cohabitation are indistinguishable.

My conclusion that men neither gain nor lose financial status upon forming unions is more surprising. Analysts tend to focus on evidence that marriage makes men (slightly) more productive in the labor market than their nonmarried counterparts, but it is incorrect to interpret this gain to intra-household specialization as an improvement in men's financial well-being. After all, men are destined to lose financially if their family's division of labor is sufficiently extreme—a man's family need increases substantially (on the order of 70%) when an adult partner joins his household, and no amount of productivity enhancement will increase his earnings enough to offset that change. My evidence shows that in an era when intra-household specialization has diminished and women's income contributions have increased, men can be expected to "break even" financially upon marrying or cohabiting.

I conclude by considering one final question: Do gains in "total family income per adult equivalent" matter? On one hand, this outcome comes closer to capturing the economic gains to union formation than do more limited measures, such as average hourly wages and hours

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[14]Hispanic respondents account for only 15% of the person-year observations and only 147 cohabitation-to-marriage transitions, so I exclude them from the race-specific analysis.

worked. Moreover, it is easily measured and compared across families. If we are intent on cataloging the various dimensions in which marriage is beneficial to individuals, this is one outcome where guesswork is unnecessary. As my analysis demonstrates, it is relatively straightforward to identify the causal effects of marriage and cohabitation on women's *and* men's financial well-being.

On the other hand, "family income per adult equivalent" does not measure individual consumption or well-being. It excludes the value of nonmarket time, and also assumes effective income (which accounts for economies of scale and the presence of children) is shared equally among adult equivalents within the household. Relatively few couples are likely to use such a resource allocation scheme regardless of whether they are married or cohabiting. As I documented earlier, the data consistently reject a model in which couples maximize a single, joint utility function (Browning *et al.* 1994; Lundberg *et al.* 1997; Phipps and Burton 1998; Thomas 1990; Winker 1997). Instead, the data indicate that intra-family resource allocations are determined through bargaining. We have no evidence that "equal shares" is the dominant rule for married or cohabiting couples.

From a policy perspective, what matters most is how families use their income to acquire such shared goods as safe neighborhoods and schooling for their children, as well as individually-consumed goods such as health care. I believe my analysis—as well as other research on the income effects of marital transitions—is a key component of a broader literature that focuses on these essential outcomes. If marital status is found to have a causal effect on these outcomes, the final question to be addressed is whether the effect arises *as a result* of the income premium gained by married and cohabiting women, or whether it exists even after income differences are taken into account. The answer to this question will reveal whether the link between marital status and income "matters" insofar as it fuels consumption-related outcomes, and whether married and cohabiting women exploit their income gains with equal success.

APPENDIX

Table A-1. OLS and Difference Estimates for Specifications in Table 3
(Dependent variable is log of total family income per adult equivalent)

	Model 1				Model 2			
	OLS		Difference		OLS		Difference	
Variable	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Women								
Intercept	-14.60	3.391			-13.44	3.398		
1 if black	-.153	.010			-.161	.011		
1 if Hispanic	-.038	.012			-.040	.012		
Age	1.857	.469	1.712	.470	1.721	.469	1.496	.473
$Age^2/10$	-.805	.239	-.694	.242	-.745	.239	-.601	.243
$Age^3/100$	.156	.053	.127	.054	.144	.053	.109	.054
$Age^4/1000$	.011	.004	.009	.004	-.010	.004	-.007	.004
1 if enrolled in school	-.259	.012	-.188	.012	-.252	.012	-.186	.012
Highest grade completed	.089	.002	.120	.009	.089	.002	.115	.009
Annual weeks worked	.014	.000	.012	.000	.014	.000	.011	.000
1 if children present	-.420	.010	-.351	.012	-.412	.010	-.343	.012
1 if live with parents	-.199	.012	-.081	.013	-.192	.012	-.077	.013
1 if single	-.495	.011	-.368	.018	-.566	.017	-.468	.033
Duration if single					.008	.002	.015	.004
Men								
Intercept	-6.189	3.301			-6.623	3.305		
1 if black	-.182	.010			-.182	.010		
1 if Hispanic	-.034	.012			-.034	.012		
Age	.805	.458	.637	.598	.854	.459	.656	.458
$Age^2/10$	-.332	.235	-.198	.236	-.352	.234	-.207	.236
$Age^3/100$	.061	.053	.026	.053	.065	.053	.029	.053
$Age^4/1000$	-.004	.004	-.001	.004	-.005	.004	-.001	.004
1 if enrolled in school	-.368	.013	-.257	.013	-.368	.013	-.257	.013
Highest grade completed	.093	.002	.090	.009	.093	.002	.090	.009
Annual weeks worked	.018	.000	.018	.000	.018	.000	.018	.000
1 if children present	-.392	.011	-.385	.011	-.392	.011	-.383	.011
1 if live with parents	-.283	.011	-.120	.011	-.283	.011	-.119	.011
1 if single	-.233	.011	.021	.016	-.213	.018	.031	.032
Duration if single					-.003	.002	-.002	.004
Women and Men								
1 if cohabiting	-.101	.015	.019	.017	-.105	.021	.021	.021
Duration if cohabiting					-.002	.005	-.005	.007
Duration if married					-.001	.001	-.009	.004
Root MSE	.657		.623		.657		.623	
Number of observations	50,917		41,078		50,917		41,078	
Number of individuals	9,839		9,839		9,839		9,839	

Note: All specifications also include dummy variables for each calendar year. Standard errors account for nonindependence of observations within person-specific clusters.

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Table 1. Summary Statistics for Variables Used in Income Models

	Women		Men	
Variable	Mean	S.E.	Mean	S.E.
Dependent variables				
Log of own income (Y^O)	2.45	.007	2.85	.006
Log of total family income (Y^F)	3.21	.006	3.09	.006
Log of total family income per adult equivalent (Y^F / AE)				
Using Fuchs's scale	2.71	.005	2.73	.006
Using NRC measure	2.70	.005	2.73	.006
Per capita	2.46	.006	2.67	.006
Explanatory variables				
Age	27.85	.035	27.50	.033
1 if black	.25		.26	
1 if Hispanic	.15		.15	
1 if enrolled in school	.16		.15	
Highest grade completed	13.35	.014	12.99	
Annual weeks worked	41.83	.102	41.76	.103
1 if children present in household	.52		.61	
1 if live with parents	.13		.12	
1 if single	.45		.58	
1 if cohabiting	.09		.08	
Duration if single	3.18	.032	3.84	.030
Duration if cohabiting	1.11	.025	1.12	.025
Duration if married	3.41	.031	2.36	.026
Number of observations	24,569		26,348	
Number of individuals	4,700		5,139	

Note: Each income measure is expressed in thousands of dollars divided by the GDP implicit price deflator (1996=1.00). Each model also includes higher-order terms in age and dummy variables for each calendar year. Standard errors account for nonindependence of observations within person-specific clusters.

Table 2. Sample Means of Alternative Income Measures by Marital Status

Variable	Marital Status		
	Single	Cohabiting	Married
Total family income (Y^F)	$18,656	$44,025	$53,437
	(114.42)	(787.55)	(254.28)
Total family income per adult equivalent (Y^F / AE)	$17,386	$22,101	$25,318
	(111.48)	(410.05)	(120.50)
Men's own income (Y^o)	$19,091	$22,892	$34,436
	(162.92)	(556.60)	(267.30)
Women's own income (Y^o)	$16,399	$18,958	$18,418
	(134.26)	(449.45)	(157.60)
Number of observations	26,347	2,154	22,425

Note: Adult equivalent is defined by Fuchs's scale. Incomes figures are divided by the GDP implicit price deflator (1996=1.00). Standard errors of means are in parentheses.

Table 3. Estimated Effects of Marital Status Transitions on Log of Total Family Income Per Adult Equivalent

	Model 1		Model 2	
Marital Status Transition	OLS	Difference	OLS	Difference
Women				
Single to cohabiting	.394	.387	.437	.440
	(.017)	(.020)	(.021)	(.027)
Single to married	.495	.368	.542	.416
	(.011)	(.018)	(.014)	(.026)
Men				
Single to cohabiting	.132	-.002	.116	-.011
	(.017)	(.019)	(.021)	(.026)
Single to married	.233	-.021	.220	-.035
	(.011)	(.016)	(.015)	(.025)
Women and Men				
Cohabiting to married	.101	-.019	.108	-.013
	(.015)	(.017)	(.016)	(.019)

Note: Computed from estimates in appendix table. Adult equivalent is defined by Fuchs's scale. Model 2 estimates assume completed duration in the last state is 3 years and current duration in the current state is 1 year. Standard errors of predictions are in parentheses.

Table 4. Estimated Effects of Marital Status Transitions on Log of Total Family Income and Log of Own Income (unadjusted for adult equivalents)

	Own Income		Total Family Income	
Marital Status Transition	OLS	Difference	OLS	Difference
Women				
Single to cohabiting	-.010	-.022	.950	.959
	(.032)	(.037)	(.021)	(.026)
Cohabiting to married	-.024	-.086		
	(.028)	(.032)		
Single to married	-.067	-.142	1.033	.934
	(.017)	(.035)	(.014)	(.025)
Men				
Single to cohabiting	.076	-.008	.660	.532
	(.030)	(.035)	(.021)	(.026)
Cohabiting to married	.174	.034		
	(.027)	(.032)		
Single to married	.256	.044	.744	.506
	(.018)	(.033)	(.015)	(.025)
Women and Men				
Cohabiting to married			.083	-.018
			(.016)	(.019)

Note: Estimates used to compute these predictions are available from the author. Estimates are for model 2, and assume completed duration in the last state is 3 years and current duration in the current state is 1 year. Standard errors of predictions are in parentheses.

Table 5. Predicted Percent Changes in Income and Adult Equivalents

Marital Status Transition	Single to cohabiting		Cohabiting to married		Single to married	
	Women	Men	Women	Men	Women	Men
Total family income	+161%*	+70%*	-2%		+154%*	+66%*
Total family income per adult equivalent	+55%*	-1%	-1%		+52%*	-3%
Adult equivalent	+68%*	+72%*	+0%		+67%*	+71%*
Own income	-2%	-1%	-8%*	+3%	-13%*	+4%

Note: The figures are computed from the difference estimates for model 2 shown in tables 3 and 4. Adult equivalent is defined by Fuchs's scale.

*Predicted changes are significantly different than zero at a 5% significance level.

Table 6: Estimated Effects of Marital Status Transitions on Log of Total Family Income Per Adult Equivalent, by Alternative Adult Equivalent Measures and Race

	AE Measure[a]		Race Subsample[b]	
Marital Status Transition	NRC	Per cap.	Whites	Blacks
Women				
Single to cohabiting	.530	.406	.451	.459
	(.027)	(.027)	(.031)	(.069)
Single to married	.506	.384	.433	.403
	(.026)	(.026)	(.029)	(.069)
Men				
Single to cohabiting	.070	-.071	-.008	-.045
	(.026)	(.026)	(.031)	(.062)
Single to married	.047	-.093	-.026	-.101
	(.025)	(.026)	(.029)	(.064)
Women and Men				
Cohabiting to married	-.011	-.008	-.012	-.034
	(.019)	(.019)	(.023)	(.045)

Note: Estimates used to compute these predictions are available from the author. All estimates are for differenced versions for model 2, and assume completed duration in the last state is 3 years and current duration in the current state is 1 year. Standard errors of predictions are in parentheses.

[a]NRC measure is $(A+0.75C)^{0.75}$ and per capita measure is $(A+C)$, where A is the number of adults and C is the number of children.

[b]Adult equivalent is defined by Fuchs's scale.

Table 7: Predicted Percent Changes in Income and Adult Equivalents, by Alternative Adult Equivalent Measures

Marital Status Transition	Single to cohabiting		Cohabiting to married		Single to married	
	Women	Men	Women	Men	Women	Men
Total family income per adult equivalent (NRC)	+70%*	+7%*	-1%		+66%*	+5%*
Adult equivalents (NRC)	+54%*	+59%*	+0%		+53%*	+58%*
Total family income per capita	+50%*	-7%*	-1%		+47%*	-9%*
Adult equivalents (per capita)	+74%*	+83%*	+0%		+73%*	+82%*

Note: The figures are computed from the difference estimates for model 2 shown in tables 4 and 6.

*Predicted changes are significantly different than zero at a 5% significance level.