Divorce as Risky Behavior

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Abstract: We extend the orthodox divorce model by assuming risk averse individuals view marriage as risky, and divorce as even (location-independent) riskier. The model predicts that conditional on the expected gains to marriage and divorce, the probability of divorce increases with relative risk tolerance. We assess this prediction by using data for first-married women and men from the NLSY79 to estimate a probit model of divorce; we control for a rich array of covariates, including relative risk tolerance. The estimates reveal that a one-point increase in risk tolerance raises the predicted probability of divorce by 4.3% for a representative man and by 11.4% for a representative woman. These findings are consistent with the model's predictions, and with the well-established view that divorce entails a greater income gamble for women than for men.

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"Better the devil you know than the devil you don't."
—English proverb

For many individuals, divorce is a high-stakes gamble. The gains they will receive by remaining married are far from certain, given that future income, asset values, and nonpecuniary rewards ("love") are susceptible to random shocks. Nonetheless, the value of a current marriage can appear to be a "sure bet" compared to the highly uncertain payoff associated with divorce. The financial well-being of divorced women in particular often depends on the generosity of property settlements, the availability of post-divorce transfers from ex-husbands or other family members, growth of their own labor market earnings, and luck in the remarriage market—all of which are subject to considerable randomness. The extensively-documented financial losses experienced by divorced women (Bianchi *et al*. 1999; Burkhauser *et al*. 1991; Duncan and Hoffman 1985; Smock *et al.* 1999) are consistent with the notion that many women trade the relative financial security of marriage for uncertain future gains. While the inherently risky nature of divorce is widely acknowledged in the policy arena and social science literature, this article is the first to address the following question: How important are individual levels of risk aversion in determining who divorces?

We begin our analysis by recasting a simple model of divorce to highlight the role of individual risk preference. Following the seminal work of Becker *et al.* (1977), we assume individuals compare the expected utilities associated with marriage and divorce on an ongoing basis in response to new information about current match quality, expected divorce costs, prospects for remarriage, and other factors. In contrast to existing studies, we assume individuals are risk averse. If divorce were the *only* alternative to involve risk, then an individual would not divorce unless the expected consumption associated with divorce exceeded the known consumption associated with marriage by an amount at least as large as her risk premium. In fact, we assume *both* alternatives involve risk, and that divorce is location-independent riskier (Jewitt 1989) than marriage. This particular definition of "riskier" ensures that the risk premium that an individual must receive in order to choose divorce over marriage increases monotonically in her Arrow-Pratt index of risk aversion. Simply put, a risk averse individual is predicted to be less likely to divorce than is her more risk tolerant counterpart.

To assess this prediction empirically, we use data from the 1979 National Longitudinal Survey of Youth (NLSY79) to estimate discrete choice models of married women's and men's divorce decisions. Our key regressor is a measure of each individual's relative risk tolerance, which is derived from responses to questions about the willingness to accept alternative income gambles. We also control for an array of variables intended to measure the economic gains to marriage and divorce, attitudes toward marriage and gender roles, characteristics of marriage markets, and state laws governing divorce and property settlements. Our estimates reveal that risk preferences play an important role in the decision to divorce, especially for women. For example, among representative women in their fourth year of marriage, a one-point (1.8 standard deviation) increase in the coefficient of risk tolerance raises the predicted probability of divorce by 11.4%. Among representative men with the same marriage duration, an identical one-point increase in risk tolerance (which equals one standard deviation in the men's distribution) is associated with a 4.3% increase in the predicted probability of divorce. When we consider *identical* men and women (for whom all characteristics equal grand means rather than gender-specific means), the marginal effects just described change to 16.7% for women and 4.0% for men. This gender comparison is consistent with the notion that risk aversion deters divorce for everyone, but that divorce entails a greater income gamble for women than for men.

Although risk and uncertainty are central components of most economic analyses of marriage and divorce, surprisingly little attention has been paid to individual heterogeneity in risk preference. The role of risk aversion is featured prominently in studies that view marriage as a mechanism for insuring against income risk (Anderberg 2001; Chiappori and Reny 2006; Chiappori and Weiss 2007; Hess 2004; Kotlikoff and Spivak 1981; Rosenzweig and Stark 1989). However, Chiappori and Reny (2006) are the first contributors to this literature to consider how individual variation in risk aversion comes into play; they argue that risk-sharing motives should lead to negative assortative matching on risk preference. Studies that take a search-theoretic approach to marital matching (Burdett and Coles 1999; Mortensen 1988) are, by their very nature, concerned with decision-making under uncertainty. Despite this focus, the assumption of risk aversion—let alone heterogeneity in risk preference—has been introduced into search models only recently. Sahib and Gu (2002) argue that unmarried, risk averse individuals establish a higher reservation level for marital partners than for cohabiting partners if marriage is the riskier of the two alternatives. In what may be the first empirical study to use a measure of individual risk preference as a determinant of marital transitions, Spivey (2005) finds that the waiting time to marriage decreases with risk aversion, presumably because risk averse individuals attach less value to continued search.

A lack of data can be blamed for the relative inattention paid to the influence of individual risk preference on marital transitions. The NLSY79 is the only large-scale, U.S. survey to elicit information on respondents' risk preferences while also supporting detailed analyses of transitions into and out of marriage. During three of the 22 interviews conducted to date, NLSY79 respondents were asked whether they would accept two hypothetical, large-stakes income gambles of varying riskiness.[1] We use multiple responses to these questions to estimate an Arrow-Pratt index of relative risk tolerance that accounts for both measurement error and aging effects. While identical income gamble questions were included in multiple rounds of the Health and Retirement Study, that survey's focus on individuals over age 50 makes it less appropriate for an analysis of divorce. The questions were also included in the Panel Study of Income Dynamics, but because they were only asked of *employed* respondents (in a single interview year), the data are of limited use in modeling marital transitions. In 2004, the German Socio-Economic Panel Study (SOEP) asked respondents to rate their willingness to take risks in a number of specific contexts, while also asking about their willingness to participate in a particular, hypothetical lottery. Because the SOEP has followed a large, representative sample of individuals for over 20 years—and collected detailed information on labor market activities and family formation—it appears to be the only viable alternative to the NLSY79 for an analysis of the effects of risk preference on marital dissolution.[2]

The rest of the paper is organized as follows. In section I, we present a model that

[1]Details on the survey and the income gamble questions are provided in section III. The design and validity of the income gamble questions—which originated in the Health and Retirement Study—are discussed in Barsky *et al*. (1997).

[2]Other sources of data on individual risk preferences include the Surveys of Consumers, Italy's Survey of Household Income and Wealth, the Dutch Brabant Survey, and the Dutch DNB Household Survey.

demonstrates how a married person's relative risk tolerance affects her choice between the less risky option of remaining married and the riskier option of divorcing. We also consider how an individual's preference for risk might directly affect the relative gains to marriage, either by affecting her choice of partner or by altering savings and investment behavior during the marriage. In section II, we explain how our theoretical model of divorce decisions can be implemented empirically by estimating a probit model. In section III, we describe the NLSY79 data used for our analysis. We explain how we construct measures of relative risk tolerance from the income gamble questions asked of NLSY79 respondents, and we describe our additional covariates that control for the expected gains associated with marriage and divorce. We present our estimates in section IV, and provide concluding comments in section V.

I. The Decision to Divorce

A. Effects of Risk Tolerance on the Choice between Two Risky Options

In modeling the decision to divorce, researchers invariably assume that married couples act to maximize their expected utility (Becker *et al.* 1977; Charles and Stephens 2004; Hoffman and Duncan 1995; Weiss 1997; Weiss and Willis 1997). While existing models explicitly allow unanticipated shocks to affect the decision to divorce, they do not consider the behavior of risk averse individuals in an environment where both alternatives involve risk.[3] In this section, we extend the analysis of divorce by assuming that all individuals are risk averse, but risk aversion differs across individuals; and that both alternatives (marriage and divorce) involve risk, but divorce is the location-independent riskier of the two options. At the same time, we simplify our model by assuming agents myopically consider the expected utility associated with marriage and divorce, rather than using the dynamic programming approach of Charles and Stephens (2004), Weiss (1997) and Weiss and Willis (1997).[4] We treat the individual (assumed for

[3]Becker *et al*. (1977) implicitly acknowledge that marriage is risky when they claim (page 1143) that "(t)he probability of divorce is smaller the greater the expected gain from marriage, and the smaller the variance of the distribution of unanticipated gains from marriage." However, they do not explicitly consider the risky nature of divorce and, in fact, appear to assume (page 1143) that agents are risk neutral.

[4]The dynamic nature of the theoretical model used by Charles and Stephens (2004) and Weiss and Willis (1997) does not carry over to their empirical model, which is a simple probit. Because we use a similar estimation strategy, we eliminate the value function from our theoretical framework, following Hoffman and Duncan (1995).

concreteness to be a woman), rather than the couple, as the decision-maker, while adopting the standard assumption (Becker *et al.* 1977) that transferable utility and wealth leads to efficient resource allocations.

Let $M_{it} = M(X_{it}, X_{it}^{h}, X_{it}^{c}, \varphi_{it}^{c})$ be the lifetime consumption that woman *i* receives if she remains married from time *t* to the end of her horizon. The woman's gain to marriage depends on current and future values of her own characteristics (X_{it}), her husband's characteristics (X_{it}^{h}), tangible factors such as joint financial assets that characterize the couple (X_{it}^{c}), and intangible characteristics of the marriage such as love (φ_{it}^{c}). The lifetime consumption the woman receives if she instead divorces at time *t* is $D_{it} = D(X_{it}, X_{it}^{h}, X_{it}^{c}, Z_{it})$, where Z_{it} represents current and future divorce costs and characteristics of the marriage market. The value of divorce includes X_{it}^{h} and X_{it}^{c} insofar as financial components of these vectors (husband's income, financial assets) affect property settlements, alimony, and child support, while components such as children affect the indirect costs of divorce.

Each woman has an increasing, concave utility function $U(C_{it})$ defined over consumption that implies an Arrow-Pratt measure of relative risk tolerance $\rho_{it} = -U' / C_{it}U''$. Before turning to the case where both marriage and divorce involve risk, we consider the decision-rules that maximize (expected) utility under two simpler scenarios. If the lifetime consumption associated with marriage and divorce are both known with certainty, the woman chooses to divorce whenever $U(D_{it}) > U(M_{it})$; empirical implementation of this model simply requires that we have data for the determinants of M_{it} and D_{it}. Alternatively, if M_{it} is known with certainty but divorce is risky, the woman divorces whenever $EU(D_{it}) > U(M_{it})$—that is, whenever $U[E(D_{it}) - \pi_{it}] > U(M_{it})$, where $\pi_{it} > 0$ is the risk premium the woman is willing to pay to receive $E(D_{it}) - \pi_{it}$ with certainty rather than face the uncertain outcome of divorce. Pratt (1964) establishes that under this scenario, π_{it} increases monotonically with the index of relative risk aversion $(1/\rho_{it})$, or decreases monotonically with ρ_{it}. Because this particular model predicts that the probability of divorce increases in ρ_{it}, its empirical analog should include a control for ρ_{it} in addition to controls for the determinants of M_{it} and D_{it}. Even if we have no direct interest in the role of risk preference as a determinant of divorce, the omission of ρ_{it} from our empirical specification could cause us to incorrectly assess the effects of factors that are correlated with risk preference.

Having established the role of risk preference when divorce is the only risky option, we turn to the scenario that we believe most accurately describes the divorce decision: marriage is risky, and divorce is even riskier. We assume divorce is the riskier option for two reasons. First, the woman's consumption while married (M_{it}) depends on the evolution of her current husband's characteristics, while her consumption while divorced (D_{it}) depends on the current and future attributes of a potential new husband; thus, D_{it} is riskier than M_{it} because its depends on which second husband (if any) is realized as well as realizations of his characteristics over time. Second, while X_{it}^{h} and X_{it}^{c} are determinants of both M_{it} and D_{it}, their contribution to D_{it} depends on how they will change over time *and* how they will be distributed after the dissolution of the current marriage. Because women's financial well-being is typically more dependent than men's on spousal income, alimony and child support (Bianchi *et al*. 1999; Burkhauser *et al*. 1991; Cancian *et al.* 1993; Light 2004; Winkler 1998), these arguments also imply that divorce entails a greater income gamble for women than for men.

In order to show that the probability of divorce increases in ρ_{it} when both options are risky, we must be explicit about the sense in which divorce is riskier than marriage. Rather than describe a stochastic process by which each factor $X_{it}, X_{it}^{h}, X_{it}^{c}, \varphi_{it}$, and Z_{it} evolves over time, we simplify the discussion by assuming that both M_{it} and D_{it} are random variables with cumulative distribution functions F_M and F_D, respectively. We further assume that F_D is location-independent riskier than F_M as defined by Jewitt (1989). This condition holds if and only if

$$\int_{-\infty}^{F_D^{-1}(p)} F_D(c)dc \geq \int_{-\infty}^{F_M^{-1}(p)} F_M(c)dc \quad \forall\ p \in (0,1).$$

As Chateauneuf *et al*. (2004) demonstrate, an alternative definition is that F_D single-crosses F_M such that the (negative) horizontal distance $F_D^{-1}(c) - F_M^{-1}(c)$ is nondecreasing in every interval below the crossing point. Location-independent risk is the most general stochastic order to ensure that the premium a risk averse individual will pay for partial insurance is monotonically decreasing in her Arrow-Pratt coefficient of risk tolerance (Chateauneuf *et al.* 2004; Jewitt 1989; Landsberger and Meilijson 1994).[5]

When both divorce and marriage are risky, the woman divorces whenever $E_D U(D_{it}) > E_M U(M_{it})$, where the expectations are formed over F_D and F_M. This condition is met whenever $E_M U(D_{it} - \pi_{it}) > E_M U(M_{it})$, where $\pi_{it} > 0$ is now the risk premium the woman is willing to pay to draw $D_{it} - \pi_{it}$ from the less-risky distribution F_M rather than face the riskier outcome of divorce. Because the assumption of location-independent risk assures that π_{it} decreases monotonically in ρ_{it}, we continue to predict that, all else equal, the probability of divorce rises with a woman's level of relative risk tolerance.

B. Effects of Risk Tolerance on the Gains to Marriage and Divorce

The preceding discussion provides our primary rationale for including a measure of relative risk tolerance among the determinants of divorce: ρ_{it} is inversely related to the risk premium needed to compensate women for the extra risk associated with divorce. Of course, a woman's risk preference can also affect her search for a husband both before and after her current marriage, and the extent to which she engages in within-household risk sharing. The matching process and risk sharing contribute to the relative gains associated with marriage which, in turn, affect the probability of divorce. In this subsection, we consider how risk preference might affect the probability of divorce through these additional channels.

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[5]Ross (1981) demonstrates that a mean-preserving spread does not guarantee that the risk premium is monotonic in the Arrow-Pratt index unless additional distributional assumptions are made. The distributional assumption of location-independent risk ensures the monotonicity of the risk premium for every nondecreasing and concave utility function. In order to include risk lovers (for whom utility functions are nonconcave), we would have to assume the more general definition of "riskiness" proposed by Bickel and Lehmann (1979); see also Landsberger and Meilijson (1994).

Consider a situation where single women search for marriage partners; for now, we set aside the option to cohabit rather than marry, as well as the ability to engage in assortative matching on risk preference. Given this simple scenario, we expect the value of search and, therefore, the reservation level for an acceptable husband to increase with the woman's degree of relative risk tolerance. This argument, which originates in the job search literature (Pissarides 1974) and is applied to marital search by Spivey (2005), suggests that select components of M_{it} ($X_{it}, X_{it}^{h}, X_{it}^{c}$, and/or φ_{it}^{c}) increase in ρ_{it}. In other words, the probability of divorce is predicted to *decrease* in ρ_{it} to the extent that ρ_{it} is positively correlated with *unmeasured* components of M_{it}.

This naïve prediction does not necessarily hold once we acknowledge that cohabitation is another option available to single women. As shown formally by Sahib and Gu (2002), a risk averse woman can mitigate the risk inherent in marriage by forming a cohabiting union with her potential mate. In essence, cohabitation can be used as a testing ground for conducting what Becker *et al.* (1977) term "intensive search." If the likelihood of cohabitation decreases in ρ_{it}, then match quality might be higher among relatively low-ρ women who cohabit before marriage than among relatively high-ρ women who transition directly from single to married. Moreover, because women can expect to re-launch the search process after a divorce, any relationship between risk preference and M_{it} can also exist between risk preference and D_{it}.

Risk sharing motives provide another mechanism by which a woman's risk preference can potentially affect the gains associated with marriage and divorce—and, in turn, the probability of divorce. Given the consumption-smoothing opportunities inherent in a two-adult household (Weiss 1997), a naïve prediction is that a highly risk averse woman simply derives a higher level of expected utility from marriage than does a more risk tolerant woman. However, Chiappori and Reny (2005) argue that the desire to share risk leads to negative assortative matching on risk preference. If high-ρ women are matched with low-ρ husbands and vice versa, then the additional marital consumption accruing to the *couple* as a result of risk sharing behavior is unlikely to be tied to the *woman's* risk preference. Given our maintained assumption that marital resources are allocated efficiently, the woman's gain due to risk sharing should be independent of her risk preference as well. Of course, to the extent that not all women find the "perfect" match in the risk preference dimension, we can expect elements of M_{it} that represent intra-household risk sharing to be correlated with ρ_{it}.[6]

II. Estimation of the Divorce Model

To implement our model empirically, we assume that the function $S_{it} = E_D U(D_{it}) - E_M U(M_{it})$ is linear in the factors that determine the gains to marriage and divorce. That is,

$$S_{it} = \beta_1 \rho_{it} + \beta_2 X_{it} + \beta_3 X_{it}^h + \beta_4 X_{it}^c + \beta_5 \varphi_{it} + \beta_6 Z_{it} + \varepsilon_{it}, \tag{1}$$

where ρ_{it} continues to represent the Arrow-Pratt coefficient of relative risk tolerance. The model presented in section I.A predicts that β_1 is positive: holding constant the determinants of M_{it} and D_{it}, we expect the probability of divorce to *increase* in ρ_{it} because ρ_{it} is negatively correlated with the premium needed to accept the greater risk associated with divorce. As discussed in section I.B, correlations between ρ_{it} and the unmeasured components of M_{it} and D_{it} can also affect our estimate of β_1. These correlations arise because an individual's level of risk tolerance affects marital search and risk sharing behavior which, in turn, affect the relative gains to marriage. However, we do not expect these indirect effects to be systematically positive or negative, and in light of the richness of our data we do not expect them to dominate the risk premium interpretation of β_1. In section III we explain how we measure ρ_{it} and how we control for $X_{it}, X_{it}^h, X_{it}^c, \varphi_{it}$, and Z_{it}; although not made explicit in equation (1), our controls also include dummy variables indicating the current duration of the marriage.

In equation (1), ε_{it} represents unobserved factors that influence the probability of divorce. We assume ε_{it} is a normally distributed random variable and that, conditional on the control variables, ε_{it} has a zero mean and constant variance. With these

[6]If women "self insure" against a potential divorce by increasing their labor supply (Greene and Quester 1982, Johnson and Skinner 1986, Stevenson 2007), relatively risk averse women may contribute a relatively high share of total household income.

assumption in place, we estimate the probability of divorce (the probability that $S_{it} > 0$) as a probit model. We compute standard errors that account for nonindependence of ε_{it} across observations for a given individual.

III. Data

A. Sample selection

Our primary data source is the 1979 National Longitudinal Survey of Youth (NLSY79), although as described in section III.C we use data from other sources to characterize the marriage market and prevailing divorce laws facing each individual. The original NLSY79 sample consists of a nationally representative subsample of 6,111 individuals born between 1957 and 1964, an over-sample of blacks, Hispanics, and disadvantaged nonblacks/non-Hispanics ("whites") born between 1957 and 1964, and a sample of 1,280 individuals born between 1957 and 1961 who enlisted in the military prior to September 30, 1978. All 12,686 sample members were interviewed in 1979, and subsequent interviews were conducted every year through 1994 and biennially thereafter.[7] We use data from survey years 1979 through 2004 for this study.

We construct separate samples of men and women for our analysis. Of the 6,283 female and 6,403 male respondents in the NLSY79, we drop 1,093 women and 512 men from our samples because they marry prior to their first interview in 1979. We impose this selection rule because a subset of our covariates (*e.g.,* premarital cohabitation) cannot be identified for in-progress marriages. We eliminate 1,108 (1,704) of the remaining women (men) because they remain "never married" at the time of their last interview date. This leaves us with 4,082 women and 4,187 men whose first marriages are observed from their beginning to their dissolution or to the respondent's last interview date. We are forced to drop 850 (866) of these women (men) from our samples because we lack responses to at least one series of income gamble questions asked in 1993, 2002 and 2004; 811 of these women and 824 of these men leave the survey prior to 1993, while the remaining 39 women and 42 men simply failed to respond to the questions.

[7]Most members of the military subsample were dropped from the survey in 1985, and all members of the disadvantaged white over-sample were dropped in 1991. Thus, none of these respondents appear in our sample because we require valid responses to income gamble questions asked in 1993 and beyond.

Finally, we drop 18 women and 23 men because their marriage begins the same year as their last interview, which prevents us from observing the marriage over at least one 12-month interval. After imposing these selection criteria, we are left with samples of first marriages for 3,214 women and 3,298 men.

In modeling the decision to divorce, we use a sample of 38,733 person-year observations for the 3,214 women, and 37,662 person-year observations for the 3,298 men. Each individual (*i.e.*, each marriage) contributes one observation per year from its onset until the time it ends in divorce *or* the individual is last interviewed. We include annual observations for those years (1995, 1997, *etc*.) when NLSY79 respondents were not interviewed by imputing values for select time-varying covariates from information reported during adjacent interviews. Each marriage contributes between one and 24 observations to the sample, with a mean (standard deviation) of 12.1 (6.9) observations per marriage.

B. Measuring risk tolerance

In 1993, 2002 and 2004, NLSY79 respondents were asked the following question about their willingness to accept income risk:

> *Suppose that you are the only income earner in the family, and you have a good job guaranteed to give you your current (family) income every year for life. You are given the opportunity to take a new and equally good job, with a 50-50 chance that it will double your (family) income and a 50-50 chance that it will cut your (family) income by a third. Would you take the new job?*

Respondents who answered "yes" were asked as a follow-up whether they would still take the new job if the chances were 50-50 that it would double their income and 50-50 that it would cut their income by *one-half.* Respondents who answered "no" to the initial question were asked a follow-up question in which the gamble was changed to a 50-50 chance of doubling income and a 50-50 chance of cutting it by *20 percent*.

We form a four-way, ordinal ranking based on individuals' direct responses to this series of income gamble questions. The first category identifies the least risk tolerant individuals who decline gambles that could cut their income by one-third and one-fifth. Category 2 identifies individuals who decline the gamble with a downside risk of one-third, but accept the downside risk of one-fifth. Individuals who decline the gamble with a downside risk of one-half but accept the "one-third" gamble are in category 3, and category 4 represents the most risk tolerant individuals who accept gambles that entail a potential loss of both one-third and one-fifth of their income.

We use these categorical variables to estimate each individual's Arrow-Pratt coefficient of relative risk tolerance in each year. The resulting variable (*Risk Tolerance*) is a cardinal measure of risk that can be compared in a meaningful fashion across individuals, and is inversely related to the risk premium as described in section I.A. To compute *Risk Tolerance*, we modify the estimation procedure proposed by Barsky *et al.* (1997) to incorporate the multiple income gamble responses available in the NLSY79; this allows us to attribute within-person variation in risk tolerance to both aging and measurement error (or other time-varying shocks).[8]

The first step in the estimation procedure is to assume that each individual's utility over lifetime consumption (*C*) exhibits constant relative risk aversion:

$$U(C_i) = \frac{C_i^{1-1/\rho_{it}}}{1-1/\rho_{it}}$$

where ρ_{it} is the coefficient of relative risk tolerance for individual *i* at time *t*; this parameter is allowed to vary across individuals and over time for a given individual. We do not observe ρ_{it} directly, but we can infer lower and upper bounds for each individual's risk parameter from her categorical responses to the income gamble questions asked in the NLSY79. For example, if a respondent accepts the first gamble (*i.e.,* is willing to risk her current income for a 50-50 chance of doubling income or cutting income by one-third) but rejects the second (is unwilling to gamble on a 50-50 chance of doubling her income or cutting it in half), the following must hold:

$$\frac{1}{2}\mathrm{U(2I)} + \frac{1}{2}\mathrm{U}(\frac{2}{3}\mathrm{I}) \geq \mathrm{U(I)} \quad \text{and} \quad \frac{1}{2}\mathrm{U(2I)} + \frac{1}{2}\mathrm{U}(\frac{1}{2}\mathrm{I}) < \mathrm{U(I)}.$$

Given our parameterization of the utility function, we can infer that this individual's true ρ_{it} must lie between 0.5 and 1.0.[9]

We further assume that an individual's true ρ_{it} can be modeled as:

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[8]By incorporating these features, our estimation strategy is identical to Ahn's (2007) adaption of the strategy used in Barsky *et al.* (1997). See Kimball *et al*. (2007) and Sahm (2007) for other implementations of this computational method.

[9]This particular example refers to an individual in risk category 3. The lower and upper bounds for individuals in risk categories 1, 2 and 4 are (0,0.27), (0.27,0.5), and (1.0,∞).

$$\log\rho_{it} = \beta \text{AGE}_{it} + \alpha_i + u_{it}, \tag{2}$$

where $\alpha_i \sim N(\overline{\alpha}, \sigma_\alpha^2)$ and $u_{it} \sim N(0, \sigma_u^2)$. We allow variation in $\log\rho_{it}$ to depend on age, unobserved, time-constant individual factors (α_i), and time-varying shocks (u_{it}) due to measurement error and other factors not captured by the age trend. While risk preference is often viewed as an innate characteristic that is constant over an individual's life, we include age in our model in light of evidence presented in Ahn (2007), Mandal and Roe (2007) and Sahm (2007) that individuals tend to grow more risk averse with age; we present our own evidence of this pattern below. Moreover, we estimated (2) separately for men and women in light of the evidence presented by Mandal and Roe (2007) and Sahm (2007) that women are significantly more risk averse than men.

Given the parameterization described by (2), we can construct a log-likelihood function that depends on the data (age, risk category) and the parameters $\beta, \overline{\alpha}, \sigma_\alpha$, and σ_u. We compute gender-specific maximum likelihood estimates of each parameter and use these estimates to calculate each individual's expected ρ_{it} at each age that her marriage is observed. The log-likelihood function, maximum likelihood estimates, and expression for the conditional expectation of ρ_{it} are provided in the appendix. These conditional expectations form the variable *Risk Tolerance,* which reflects each individual's coefficient of relative risk tolerance in each year.

To substantiate our claim that risk preferences change over time, in table 1 we summarize how individuals' first responses to the two income gamble questions compare to their second responses. Almost 90% of the individuals in our samples provide responses in both 1993 and 2002, but for this cross-tabulation we include a small number of 1993-2004 and 2002-2004 comparisons as well. Focusing first on women, table 1 reveals that roughly half the sample falls into risk category 1 (least tolerant) based on the first response, and that 68% of these women remain in category 1 when they answer the income gamble questions a second time. Among the women whose first response places them in category 4 (most tolerant), only 25% remain in the same category—that is, 75% of these women appear to become *less* risk tolerant over time, while only 32% of the women who are initially in category 1 appear to become *more* risk tolerant over time. Among women who are initially in category 2 (3), 53% (61%) report a *lower* risk tolerance the second time, while only 28% (17%) report a *higher* level. These patterns reveal why we model $\log \rho_{it}$ as a function of age: while much of the within-person variation in risk category can be attributed to reporting error, women appear to become less risk tolerant as they age.

While the patterns described in the preceding paragraph apply to men as well, table 1 reveals two salient differences between men and women: men are more risk tolerant than women, and are somewhat less likely to decrease their risk tolerance with age. For example, on the basis of their first responses, men are five percentage points less likely than women to fall into category 1 (44.7% versus 49.8%) and six percentage points more likely to fall into category 4 (26.1% versus 19.5%). Among men whose first response places them in category 1 or 2, 36% report a *higher* level of risk tolerance with their second response (versus only 28-32% of women). Among men whose first response places them in category 4, 72% move to a lower level of risk tolerance on the basis of their second response (versus 75% of women).

Table 2 summarizes the distribution of the variable *Risk Tolerance* among women and men in each self-reported risk category; for this table, we use the value of *Risk Tolerance* for the year corresponding to the individual's first response to the income gamble questions. While both the mean and median of *Risk Tolerance* increase monotonically with the risk category, as expected, there is considerable variation in *Risk Tolerance* within each category. For example, it ranges from 0.21 to 1.43 among women and from 0.23 to 1.80 among men whose income gamble responses place them in category 2, even though the upper and lower bounds for that category are 0.27 and 0.50 (footnote 9). This imperfect correspondence between individuals' categorical responses and our variable reflects the fact that we "smooth" over a considerable amount of reporting error in constructing *Risk Tolerance*. As shown in appendix table A, the two estimated error variances in our $\log \rho_{it}$ model are roughly equal in magnitude for men and women, which suggests that half the total variation is due to error.

C. Other covariates

In addition to our measure of risk tolerance, we use a large number of variables to control for the expected gains to marriage and divorce. These variables are intended to capture heterogeneity in match quality, marriage-specific capital, intra-household specialization, intra-household consumption smoothing, divorce costs, remarriage opportunities, and attitudes toward marriage and divorce.[10] To organize our discussion, we group the variables into three categories: economic measures, demographic and family background characteristics, and environmental factors. Table 3 contains summary statistics for each of these variables, along with brief definitions and an indication of whether the variable is time-varying within marriages.

Our economic variables include a measure of each individual's net family assets. We construct this variable by summing the values of homes, automobiles, other possessions, cash holdings, stocks, bonds, trusts, retirement accounts, and various other assets that are reported in each interview from 1985 onward, and subtracting the reported values of mortgages, business debts, and other debts.[11] We include net assets in our model to capture the value of marriage-specific capital and public goods (homes, automobiles, *etc*.) that increase the gains to marriage and lower the probability of divorce.

We include four income measures among our covariates. First, we control for the couple's total labor income, which is the sum of the individual's wage and salary income and his or her spouse's income in the last year.[12] This variable is intended to capture the "income effect" component of the gains to marriage (Moffitt 2000; Oppenheimer 1997)—*i.e.*, the fact that couples with higher income (as well as assets) can enjoy greater joint consumption. Second, we control for the share of total family income contributed by the individual. This variable reflects the degree to which a husband and wife exploit their comparative advantages in market and home production, thereby increasing the gains to marriage (Becker 1974; Becker *et al.* 1977; Oppenheimer 1997). In addition, the share of total income contributed by the woman represents her economic independence, which is a key component of her expected gains to divorce (Oppenheimer 1997). For

[10]See Becker *et al.* (1977), Lehrer (2003) and Weiss (1997) for discussions of each theoretical argument and empirical evidence of their effects on the probability of divorce.

[11]We impute values when a respondent says she has a particular asset or debt but does not report its value. If the item's value is reported in both an earlier *and* later interview, we use the closest-reported values to linearly interpolate the missing value. If multiple values are reported either before *or* after the missing year, we use estimated coefficients from a within-person regression of asset values on year to linearly extrapolate the missing value. This procedure is identical to the method used to create a total net worth variable planned for subsequent releases of the NLSY79, although we take the additional step of imputing values for 1979-84.

[12]We impute wage and salary income for noninterview years using the same interpolation/extrapolation method that we use for the asset variable (footnote 11).

both reasons, an increase in the woman's share of income is predicted to increase the probability of divorce, holding total income constant. Following Hess (2004), we also control for the correlation coefficient between the individual's labor income and his or her spouse's labor income to measure the extent of intra-household income risk-sharing. A couple with negatively correlated incomes are best able to exploit the risk-sharing advantages of marriage, and are less likely to divorce as a result. Because income correlation cannot be computed for marriages that contribute only one observation, we also include a dummy variable indicating that the variable is missing; we set the income correlation to zero in such cases. [13]

Because individuals' labor income can be endogenously determined by their beliefs about a future divorce (Greene and Quester 1982; Johnson and Skinner 1986; Stevenson 2007), we also use model specifications that replace the income variable with predicted versions. We construct predicted total income and the predicted income share contributed by the individual from predicted values of each partner's annual income. To predict each sample member's income, we use age, age-squared, age-cubed, three dummy variables for schooling attainment, Armed Forces Qualifications Test (AFQT) scores, number of children, occupation dummies, state dummies, and the median income in the county of residence in the given calendar year.[14] To predict husbands' and wives' income, we omit AFQT scores (which are unavailable for spouses) and use the spouse's values for the remaining variables; race/ethnicity and state and county of residence are known only for respondents, so we assume husbands and wives are identical in these dimensions. To estimate the predicting equations, we use observations corresponding to first marriages for *all* male and female NLSY79 respondents, and estimate separate, sex-specific equations for blacks, Hispanics, and whites. Following Hess (2004), we predict

[13]We experimented with a number of additional income variables used by Hess (2004) and others, including the mean income gap between the husband and wife, the ratio of their within-marriage income variances, and the level of each partner's income variance. None had a statistically significant coefficient in our divorce models or, more importantly, a discernible effect on the estimated coefficients for our risk preference measure, so we do not include them.

[14]We use percentile AFQT scores constructed from scores on the Armed Services Vocational Aptitude Battery, which was administered to NLSY79 respondents in 1980. Because respondents were ages 15-22 when the test was taken, we first regress AFQT scores on year-of-birth dummies to compute age-adjusted values. Our measure of median county income comes from the City and County Data Books, which we describe in footnote 18.

the couple's income correlation directly from a race/ethnicity-specific regression of observed correlation on the same variables used to predict annual income.

The demographic controls used in our divorce model include the number of children in the household, dummy variables indicating whether any children are age six or younger or male, and a dummy variable indicating whether any children were born before the marriage began. These child-related variables are intended to capture a key component of marriage-specific capital (Becker 1974; Becker *et al.* 1977).[15] We control for whether a male child resides in the household because empirical evidence indicates that divorce is less likely and remarriage more likely for women with sons (Lundberg and Rose 2003; Morgan and Pollard 2002). The "premarriage children" variable indicates a lack of marital capital insofar as "existing" children may have biological parents outside the marriage, and it also measures match quality, given that marriages that are instigated by a pregnancy may be less strong than other marriages (Becker *et al*. 1977).

Other measures of match quality include the individual's age at marriage, the absolute value of the difference in the husband's and wife's age, the absolute value of the difference in their highest grade completed, and dummy variables indicating whether the individual cohabited with his or her current spouse immediately prior to the marriage and whether he or she instead cohabited with another partner. Individuals who marry at relatively later ages may have decreased search costs (and, therefore, higher quality marriages) as a result of prior matching experience (Becker *et al*. 1977), and positive assortative mating on age and schooling attainment are also expected to increase the gains to marriage (Becker 1974). While numerous empirical studies have shown that premarital cohabitation is associated with increased divorce (Axinn and Thornton 1992; Brien *et al* 2006; Lillard *et al.* 1995), the effect of cohabiting with one's current spouse is theoretically ambiguous. If cohabitation is used as a "testing ground," then couples who eventually choose to marry should be relatively well matched; however, divorce-prone couples may be more likely than others to self-select into premarital cohabitation.

Our demographic controls also include dummy variables indicating whether the

[15]Because marriage-specific investments are expected to increase over the course of a marriage, we also include dummy variables indicating current marriage duration to control for investments for which we lack observed proxies. Given that we also control for age at marriage, these duration dummies capture variation in current age as well.

individual is black or Hispanic (with white the omitted group), the religion in which the individual was raised (Baptist, Catholic, or other/none, with Protestant the omitted group), and whether the individual lived with her mother only, with her mother and a stepfather, or without her mother at age 14; individuals who lived with both biological parents form the omitted group. These controls are intended to capture widely-documented effects of religion, race/ethnicity, and family background on attitudes toward marriage, the characteristics of marriage markets, and other factors that influence entry into and exit from marriage (Bumpass *et al.* 1991; Lehrer and Chiswick 1993; Raley 2000). As an additional measure of marriage-related attitudes, we use responses to a question that asked NLSY79 respondents if they agree with the statement: "Women are much happier if they stay at home and take care of their children." We construct a dummy variable that equals one if the respondent agreed or strongly agreed with the statement, and zero if she disagreed or strongly disagreed.[16]

Our environmental variables include three measures of the legal climate governing divorce and the division of property in the individual's state of residence in the given calendar year. We use a dummy variable to indicate whether state law requires that only "no-fault" divorces be granted; the omitted category identifies states that either allow or require (in the given year) that "fault" be established as grounds for divorce. We also include a dummy variable indicating whether the state uses "no-fault" for property division and alimony decisions, and a variable that identifies the mandatory separation period required before a no-fault or unilateral divorce is granted; the separation duration variable equals zero if the state imposes no separation requirement.[17] Both the theoretical and empirical effects of "no fault" or unilateral divorce laws on divorce decisions have been debated in the literature for many years (Allen 1992; Becker *et al.* 1977; Friedberg 1998; Mechoulan 2006; Peters 1986; Stevenson 2007; Stevenson and Wolfers 2004), with recent findings (Wolfers 2006) suggesting that the liberalization of divorce law leads to increased divorce rates in the short-run.

We include four additional environmental variables that measure the

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[16]The question was asked in 1979, 1982, and 1987. We use the most recently-reported response for each person-year observation.

[17]Our data are taken from Ellman and Lohr (1998), Mechoulan (2006) and tables available at http://www.law.cornell.edu/topics/Table_Divorce.htm.

characteristics of the individual's county of residence for the given year.[18] These variables include the county-and year-specific unemployment rate and divorce rate, and the percent of the county population with the same race/ethnicity (black, white, or Hispanic) as the individual, and the percent of the county population that is male. Similar variables have been used by Lichter *et al.* (2002) and Gould and Paserman (2003) as controls for economic opportunities and marriage market characteristics.

IV. Findings

Tables 4A-B contain gender-specific probit estimates for three alternative specifications of our divorce model. Specification 1 includes all the controls described in section III.C (with the income variables based on actual values rather than predictions), along with our measure of relative risk tolerance. As seen in table 4A, the estimated coefficient for *Risk Tolerance* in specification 1 is large, positive, and precisely estimated for women; its estimated marginal effect is 0.0037 at the sample means. Table 4B reveals a much smaller and less precisely parameter estimate for men, with an estimated marginal effect of 0.0013. Given the unconditional, 12-month divorce rate of 0.04 (table 3), this means a one-point increase in the Arrow-Pratt index of relative risk tolerance is predicted to increase the probability of divorce by 9.25% for women and 3.25% for men.[19] The model includes a broad array of variables representing the gains associated with both marriage and divorce—including controls for risk sharing (within-couple income correlation) and match quality (*e.g.,* age at marriage, premarital cohabitation) that may themselves be affected by an individual's level of risk tolerance. Thus, we believe *Risk Tolerance* represents the risk premium each person needs as compensation for the additional risk associated with divorce. The probability of divorce increases in risk tolerance because the more risk tolerant (less risk averse) an individual is, the smaller is his or her risk premium. The estimated effect is considerably larger for women than for men because women are more risk averse than men, and are likely to face far more

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[18]These data come from various issues of the City and County Data Book, which is produced by the U.S. Census Bureau. This data source does not provide annual observations for every variable, so we use the closest available year.

[19]A one-point increase in *Risk Tolerance* corresponds to 1.8 standard deviations for women and 1.0 standard deviations for men (table 3). Alternatively, a one-point increase is slightly greater than a move from the 10^{th} percentile to the 90^{th} percentile in the distribution for women (table 5A), and somewhat less than a move from the median to the 90^{th} percentile for men (table 5B).

income risk upon divorcing (Bianchi *et al*. 1999; Burkhauser *et al*. 1991; Duncan and Hoffman 1985; Smock *et al.* 1999).

In specification 2 in table 4, we replace total income, the individual's share of total income, and the within-couple income correlation with predicted values, given that these variables are likely to be endogenous to the expected probability of divorce. The use of predicted income significantly alters our inferences about the effects of income-related factors on divorce, but has only a minor effect on the estimated coefficients for *Risk Tolerance:* for women, the specification 2 estimate (0.052) is only 4% larger than the estimate in specification 1, while for men the specification 2 estimate (0.023) is 21% larger than the corresponding specification 1 estimate. In neither case is the cross-model difference in estimates statistically different than zero at conventional significance levels. In specification 3, we revert to the "actual" income variables but omit *Risk Tolerance* from the model to determine whether its inclusion influences the estimated effects of other determinants of divorce. Surprisingly, we find that the estimated coefficients for all variables—including assets, income correlation, the individuals's income share, age at marriage, and other factors that we expect to be correlated with risk preference—are virtually invariant to the inclusion or exclusion of *Risk Tolerance*. Risk preference has a nontrivial effect on the probability of divorce, but it appears not to have a significant effect on (measurable) marital sorting or contemporaneous income and asset levels.

To assess how the estimated effect of risk preference compares in magnitude to the effects of other key determinants of divorce, we focus on the specification 1 estimates in tables 4A-B. For women, a one-point (1.8 standard deviation) increase in *Risk Tolerance* is associated with a 9.25% increase in the predicted probability of divorce. An identical 9.25% marginal effect is generated by a \$119,000 (0.5 standard deviation) loss of assets, an 18 percentage point (0.5 standard deviation) increase in the woman's income share, a 1.5 year (0.4 standard deviation) reduction in her age at marriage, or a 12 month (one standard deviation) increase in the state's mandatory separation requirement.[20] For men, a one-point (1.0 standard deviation) increase in *Risk Tolerance* generates a 3.25% increase in the predicted probability of divorce, as does a \$93,000 (0.4

[20]The predicated probability of divorce increases with the mandatory separation duration because states that do not allow unilateral divorce do not have separation requirements.

standard deviation) decrease in assets, a 13 percentage point (0.6 standard deviation) decrease in the man's income share, or a 0.6 year (0.2 standard deviation) decrease in the age at marriage. For both women and men, the estimated effect of a one-point increase in relative risk tolerance is comparable in magnitude to the estimated effects of modest changes in assets, income, and age at marriage.

In contrast, the estimated effects of risk tolerance are dominated in magnitude by the impact of a number of background and taste factors. Women without male children are predicted to be 16% (0.0064/0.04) more likely to divorce than are women with boys, the presence of children born prior to marriage is predicted to raise women's divorce probabilities by 49%, women who lived apart from their mothers at age 14 are predicted to be 24% more likely to divorce than are women who lived with both biological parents, and women without "traditional views" (*i.e.,* who disagree that "women are much happier if they stay at home and take care of their children") are predicted to be 16% more likely to divorce than are their "traditional" counterparts. While "traditional views" proves to have an unimportant effect on divorce probabilities for men, the absence of pre-school children (ages 0-6) is predicted to raise a man's divorce probability by 45% while the presence of "premarriage" children is predicted to raise the same probability by 25%; men who lived apart from their mothers at age 14 are predicted to be 13% more likely to divorce than are men who lived with both biological parents.

To judge these magnitudes further, we consider the predicted, one-year divorce probabilities of representative women and men at various current marriage durations and with varying levels of risk tolerance. In the top panels of tables 5A-B, we consider representative individuals for whom all variables other than *Risk Tolerance* and duration equal the mean (if continuous) or mode (if discrete) for a subsample of individuals with the *same gender* and marriage duration. We assign each representative man and woman a current marriage duration ranging from one year to six years, and a level of risk tolerance equal to different points in his or her *gender-specific* risk preference distribution. Table 5A reveals that in the fourth year of marriage, a representative woman with risk tolerance at the 10^{th} percentile has a 3.4% chance of divorcing in the next year, while an otherwise identical woman with risk tolerance at the 90^{th} percentile has a predicted probability that is 11.8% higher. At all marriage durations except 1-12 months, where both predicted and actual divorce probabilities are extremely low, we find that an increase in risk tolerance from the 10th percentile to the 90th percentile is associated with a 10-12% increase in the woman's predicted probability of divorce. When we move from the median level of risk tolerance to the 90th percentile, the predicted probability of divorce increases by 8-9% at most durations. At any marriage duration beyond 12 months, the most risk tolerant woman in our sample is roughly 80% more likely to divorce than is her counterpart with the minimum level of risk tolerance. The top panel of table 5B reveals that the corresponding estimates are considerably smaller for men. Focusing on current durations of 37-48 months, the predicted probability of divorce increases by 8.3% when we move from the 10th percentile to the 90th percentile, by 4.3% when we move from the median to the 90th percentile, and by 54.5% when we move from the minimum to the maximum level of risk tolerance. The corresponding numbers for women are 11.8%, 8.6%, and 76.5%, despite the fact that each movement within the gender-specific risk distribution represents a smaller *absolute* change in risk tolerance for women than for men.

While there is little question that the relationship between risk tolerance and divorce probabilities is more pronounced for women than for men, we have thus far defined representative individuals with respect to their own gender, and assigned these individuals gender-specific values of *Risk Tolerance*. In the bottom panel of tables 5A-B, we consider identical levels of risk preference for both men and women, based on a pooled distribution of *Risk Tolerance.* We compute one set of predicted probabilities for a representative woman and a representative man who continue to possess gender-specific mean and modal characteristics, and another set for an identical woman and man, both of whom are assigned mean and modal characteristics based on a pooled sample of men and women. Focusing on the (roughly) one-point increase in *Risk Tolerance* that corresponds to a movement from the median to the 90th percentile within the pooled distribution, we find that a representative woman with the higher level of risk tolerance has a 3.9% predicted probability of divorce in the next year, which is 11.4% higher than the predicted probability for her counterpart with a median level of risk tolerance. Among men, the corresponding change in predicted divorce probabilities is 4.3%. When we instead consider identical changes in risk tolerance for *identical* men and women, we find that a one-point increase in *Risk Tolerance* (from the median to the 90th percentile) is associated with a 16.7% increase in the predicted probability of divorce for women, and a 4.0% increase for men. In short, we find that women are more responsive than observationally equivalent men to a given change in risk preference. This finding is consistent with the notion that divorce is riskier for women than for men.

V. Concluding Comments

Beginning with Becker *et al*. (1977), researchers analyzing the decision to divorce have invoked a model in which risk neutral agents choose the state (marriage or divorce) that maximizes expected utility. To implement this model empirically, researchers invariably estimate a discrete choice model for the probability of divorce in which the covariates are determinants of the consumption associated with marriage (*M*) and divorce (*D*). In this study, we extend the orthodox divorce model by assuming that agents are risk averse, remaining married is a risky option, and divorce is an even riskier option. As long as divorce is *location-independent* riskier than marriage, our model yields the familiar result that the risk premium that compensates agents for the greater risk associated with divorce decreases monotonically in the Arrow-Pratt index of risk tolerance. In short, we predict that, conditional on *M* and *D*, the probability of divorce increases with the level of individual risk tolerance.

We assess this prediction by using samples of "first married" individuals from the 1979 National Longitudinal Survey of Youth to estimate gender-specific probit models of divorce. We control for a rich array of determinants of *M* and *D*, and we include a measure of the Arrow-Pratt coefficient of relative risk tolerance derived from responses to questions on the willingness to make hypothetical, large-stakes gambles with permanent income. We find that relative risk tolerance is an important determinant of divorce. With current marriage duration equal to 37-48 months and all other characteristics equal to gender-specific sample means and modes, the most risk tolerant woman in the sample is 76% more likely to divorce than is the least risk tolerant woman. A woman with risk tolerance equal to the 90th percentile is 11.8% more likely to divorce than is a woman whose risk tolerance equals the 10th percentile, and 8.6% more likely than a woman with the median level of risk tolerance. Among men, for whom divorce entails much less of an income gamble, a corresponding move from the median to the

90th percentile (in the male-specific distribution) is associated with a 4.3% decrease in the predicted probability of divorce.

Our evidence indicates that holding constant all other factors—from initial "match quality" to the level of marriage-specific investments to the couple's joint financial well-being to the individual's degree of financial independence—highly risk averse individuals are more likely to remain married, while their more risk tolerant counterparts are more likely to divorce. Our interpretation of this finding is that highly risk averse individuals require a relatively larger gain in expected consumption before they are willing to accept the greater risk associated with divorce. If this interpretation is correct, then existing policies designed to raise the welfare of divorced women by improving the enforcement of child support agreements and providing income assistance to low-income, unmarried mothers may serve the additional purpose of enticing relatively risk averse women to end their marriages.

Appendix: Additional Details on Computing a Measure of Relative Risk Tolerance

Equation (2) in section III.B describes how we model each individual's true relative risk tolerance parameter (ρ_{it}) as a gender-specific function of age, individual unobserved factors that are fixed over time, and idiosyncratic shocks due to reporting error. We use that model, along with data on age and categorical responses to the income gamble questions, to compute gender-specific maximum likelihood estimates for the parameters $\beta, \overline{\alpha}, \sigma_\alpha$, and σ_u, which we then use to calculate each individual's expected ρ_{it} at every age. In this appendix, we describe the sample log-likelihood function, provide the maximum likelihood estimates, and give expressions for the conditional expectation of ρ_{it}.

We form the sample log-likelihood functions for men and women by summing each individual's log-likelihood function. Each individual's contribution takes a different form depending on whether she provides one, two, or three responses to the income gamble questions asked in 1993, 2002, and 2004. If individual *i* answers the income gamble questions in a single year *t*, the probability that her risk category (c_{it}) is *j* (*j*=1,2,3,4) is:

$$
\begin{aligned}
\mathrm{P}(c_{it} = j \mid \mathrm{Age}_{it}) &= \mathrm{P}(\log \underline{\rho}_j < \log \rho_{it} < \log \overline{\rho}_j) \\
&= \Phi\left(\frac{\log \overline{\rho}_j - \overline{\alpha} - \beta AGE_{it}}{\sqrt{\sigma_\alpha^2 + \sigma_u^2}}\right) - \Phi\left(\frac{\log \underline{\rho}_j - \overline{\alpha} - \beta AGE_{it}}{\sqrt{\sigma_\alpha^2 + \sigma_u^2}}\right)
\end{aligned}
$$

where $\underline{\rho}_j$ and $\overline{\rho}_j$ are the lower and upper bounds associated with risk category *j* (see footnote 9) and Φ is the univariate normal cumulative distribution function. If individual *i* answers the income gamble questions in years *t* and *s*, the probability that her risk category is *j* in year *t* and *k* in year *s* is:

$$
\begin{aligned}
\mathrm{P}(c_{it} = j, c_{is} = k \mid \mathrm{Age}_{it}, \mathrm{Age}_{is}) &= \mathrm{P}(\log \underline{\rho}_j < \log \rho_{it} < \log \overline{\rho}_j, \log \underline{\rho}_k < \log \rho_{is} < \log \overline{\rho}_k) \\
&= \int \Bigg[\left\{ \Phi\left(\frac{\log \overline{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u}\right) \right\} \times \\
&\qquad \left\{ \Phi\left(\frac{\log \overline{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u}\right) \right\} \Bigg] dF(\alpha_i).
\end{aligned}
$$

If an individual answers the income gamble questions in all three years, we extend the preceding expression to define the probability that her risk category is *j* in year *t*, *k* in year *s,* and *l* in year *r:*

$$
\begin{aligned}
&\mathrm{P}(c_{it} = j, c_{is} = k, c_{ir} = l \mid \mathrm{Age}_{it}, \mathrm{Age}_{is}, \mathrm{Age}_{ir}) \\
&\quad = \mathrm{P}(\log \underline{\rho}_j < \log \rho_{it} < \log \overline{\rho}_j, \log \underline{\rho}_k < \log \rho_{is} < \log \overline{\rho}_k, \log \underline{\rho}_l < \log \rho_{ir} < \log \overline{\rho}_l) \\
&\quad = \int \left[\left\{ \Phi\left(\frac{\log \overline{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u} \right) - \Phi\left(\frac{\log \underline{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u} \right) \right\} \times \right. \\
&\qquad \left\{ \Phi\left(\frac{\log \overline{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u} \right) - \Phi\left(\frac{\log \underline{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u} \right) \right\} \times \\
&\qquad \left. \left\{ \Phi\left(\frac{\log \overline{\rho}_l - \alpha_i - \beta AGE_{ir}}{\sigma_u} \right) - \Phi\left(\frac{\log \underline{\rho}_l - \alpha_i - \beta AGE_{ir}}{\sigma_u} \right) \right\} \right] dF(\alpha_i).
\end{aligned}
$$

The maximum likelihood estimates that we compute are given in table A:

Table A: Maximum Likelihood Estimates
of Risk Preference Parameters

	Women		Men	
Parameter	Estimate	S.E.	Estimate	S.E.
β	-.035	.003	-.047	.003
$\overline{\alpha}$	-.178	.104	.525	.114
σ_α	1.277	.029	1.390	.032
σ_u	1.131	.019	1.222	.021
Log likelihood	-13,746.41		-13,599.71	
No. individuals	4,618		4,577	
No. observations	12,481		11,903	

We use the estimates in Table A to compute the expected coefficient of relative risk tolerance for every individual *i* at *any* time τ, conditional on her categorical response to the income gamble questions, her age at the time(s) those questions were answered, and her age at τ. The computation of these conditional expectations depends on whether the individual responds to the income gamble questions once, twice, or all three times.

For an individual i whose response to the income gamble questions in year t (and no other year) places her in risk category j, her expected relative risk tolerance at time τ is:

$$E(\rho_{it} \mid c_{it} = j, AGE_{it}, AGE_{i\tau}) = \exp(\hat{\bar{\alpha}} + \hat{\beta}\text{AGE}_{i\tau} + \frac{1}{2}\hat{\sigma}_{\alpha}^2)\frac{\text{P}(\log\underline{\rho}_j < \log\rho_{it} + \hat{\sigma}_{\alpha}^2 < \log\overline{\rho}_j)}{\text{P}(\log\underline{\rho}_j < \log\rho_{it} < \log\overline{\rho}_j)}.$$

If an individual answers the income gamble questions in years t and s, the expectation is:

$$E(\rho_{it} \mid c_{it} = j, c_{is} = k, AGE_{it}, AGE_{is}, AGE_{i\tau}) =$$

$$\exp(\hat{\bar{\alpha}} + \hat{\beta}\text{AGE}_{i\tau} + \frac{1}{2}\hat{\sigma}_{\alpha}^2)\frac{\text{P}(\log\underline{\rho}_j < \log\rho_{it} + \hat{\sigma}_{\alpha}^2 < \log\overline{\rho}_j, \log\underline{\rho}_k < \log\rho_{is} + \hat{\sigma}_{\alpha}^2 < \log\overline{\rho}_k)}{\text{P}(\log\underline{\rho}_j < \log\rho_{it} < \log\overline{\rho}_j, \log\underline{\rho}_k < \log\rho_{is} < \log\overline{\rho}_k)}.$$

The preceding expression is extended accordingly for an individual who responds in all three years.

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Table 1: Distribution of Risk Category Based on Second Response
by Risk Category Based on First Response

Risk Category (1^{st} response)	**Women**					**Men**				
	Risk Category (second response)					Risk Category (second response)				
	1	2	3	4	All	1	2	3	4	All
1	67.5	10.1	11.7	10.6	[49.8]	64.1	8.8	12.6	14.6	[44.7]
2	53.0	18.9	16.5	11.6	[13.4]	47.3	16.9	17.8	18.1	[12.4]
3	49.5	11.7	21.5	17.3	[17.2]	48.6	13.4	21.5	16.5	[17.9]
4	46.6	11.9	17.0	24.5	[19.5]	43.2	9.8	19.3	27.7	[26.1]
All	58.4	11.9	15.1	14.6	[100.0]	53.9	10.8	16.5	18.7	[100.0]
No. individuals	1,686	334	435	422	2,887	1,573	316	482	547	2,918

Note: Risk categories are based on responses to income gamble questions asked in 1993, 2002 and 2004. Samples exclude 327 women and 380 men who respond only once. Among the remaining individuals, 95% provide their first response in 1993 and their second response in 2002. Category 1 (4) consists of the least (most) risk tolerant individuals. Numbers in brackets are percents of column totals, and all other numbers are percents of row totals

Table 2: Summary Statistics for Risk Tolerance Variable
by Risk Category Based on First Response

Risk category (1st response)[a]	Risk Tolerance[b]					
	Mean	S.D.	Min.	Median	Max.	N
	Women					
1	.25	.18	.11	.20	1.06	1,601
2	.46	.25	.21	.42	1.43	430
3	.60	.30	.26	.55	1.76	556
4	1.15	.85	.38	.99	3.67	628
All	.51	.56	.11	.38	3.67	3,214
	Men					
1	.32	.26	.11	.26	1.31	1,480
2	.59	.37	.23	.49	1.80	366
3	.74	.44	.29	.60	2.25	582
4	1.65	1.30	.43	1.28	5.42	870
All	.77	.90	.11	.47	5.42	3,298

[a]See note to table 1 for variable definition.
[b]Computed relative risk tolerancefor the same year as the "first response" to the income gamble questions. See section III.B and the appendix for details.

Table 3: Definitions and Summary Statistics for Variables Used in Divorce Model

Variable	Definition	Women Mean	Women S.D.	Men Mean	Men S.D.
Divorce	1 if divorces during interval	.04		.04	
Risk variable					
Risk Tolerance	Arrow-Pratt coeff. of relative risk tolerance	.51	.56	.75	.95
Income variables					
Assets	Total net family assets[a]	102.37	256.73	107.97	263.13
Total income	Sum of spouses' labor incomes[a]	56.03	41.53	54.95	40.46
Income share	Own share of total income[a]	33.95	22.24	65.76	22.65
Income correlation	Correlation between spouses' incomes[a†]	.08	.54	.06	.58
No correlation	1 if income correlation is missing[†]	.01		.01	
Pred. total income	Predicted total income[b]	53.19	24.94	49.25	24.32
Pred. income share	Predicted own share of total income[b]	34.98	17.17	71.30	17.99
Pred. income corr.	Predicted income correlation[b†]	.09	.17	.06	.18
Demographic variables					
Number of kids	Number of children in household	1.42	1.19	1.35	1.20
Kids age 0-6	1 if any children age six or younger	.50		.48	
Male kids	male children	.53		.49	
Premarriage kids	children born before marriage[†]	.16		.06	
Age at marriage	Age at marriage[†]	23.77	4.35	25.02	4.21
Age gap	Difference in spouses' ages[†]	3.43	3.61	2.78	2.77
School gap	Difference in spouses' years of school[†]	1.42	1.62	1.31	1.52
Cohab. w/ spouse	1 if cohabited with husband[†]	.31		.35	
Cohab. with other	other partner[†]	.02		.02	
Black	1 if woman is black[†]	.24		.23	
Hispanic	Hispanic[†]	.19		.20	
Baptist	1 if religion is Baptist[†]	.26		.25	
Catholic	Catholic[†]	.39		.38	
Other religion	Other or no religion[†]	.10		.11	
Live with mom	1 if lived with mother only, age 14[†]	.14		.14	
Live w/ mom/step	mother/stepfather, age 14[†]	.06		.06	
Live without mom	no mother, age 14[†]	.07		.07	
Traditional views	1 if agrees women are happier at home[†]	.27		.37	
Environmental variables					
No fault	1 if no-fault law for divorce[c]	.35		.35	
Property no fault	property settlement[c]	.42		.42	
Separation dur.	Minimum required separation (months)[c]	10.59	11.53	10.54	11.50
Cty unemp. rate	County unemployment rate	6.90	3.16	6.77	3.04
Cty divorce rate	County divorce rate	4.91	1.96	4.93	1.89
County race	Percent of county population same race	59.63	31.19	59.21	32.46
County male	Percent of county population male	48.82	1.32	48.86	1.22
Number of person-year observations		38,733		37,662	
Number of individuals (number of first marriages)		3,214		3,298	

Continued on next page.

Table 3: Footnotes

[†]Variable does not change value over the duration of the marriage.

[a]Income levels are deflated by the CPI-U and expressed in thousands of 2000 dollars.

[b]Predicted total income is the sum of the individual's and his/her spouse's predicted annual income. Predicted income share is the individual's predicted annual income divided by predicted total income. Predicted income correlation is predicted directly from the same variables (described in section III.C) used to predict annual income.

[c]Based on state-specific divorce laws for the given calendar year.

Note: The model also includes dummy variables identifying current marriage duration which, given the control for age at marriage, also identify the individual's age.

Table 4A: Probit Estimates of Effects of Variables on Probability of Divorce (Women)

Variable	Specification 1			Specification 2[a]			Specification 3		
	Coeff.	S.E.	Marg.	Coeff.	S.E.	Marg.	Coeff.	S.E.	Marg.
Risk Tolerance	.050	.023	.0037	.052	.022	.0037	—	—	—
Income variables									
Assets/100	-.042	.014	-.0031	-.030	.012	-.0021	-.042	.014	-.0031
Total income/100	.006	.050	.0005	-.504	.091	-.0354	-.007	.050	-.0005
Income share	.003	.001	.0002	.009	.001	.0006	.003	.001	.0002
Income correlation	-.035	.027	-.0026	.189	.092	.0131	-.033	.027	-.0024
No correlation	1.519	.174	.3362	1.486	.172	.3171	1.524	.174	.3383
Demographics									
Number of kids	-.077	.020	-.0056	-.031	.020	-.0021	-.076	.020	-.0056
Kids age 0-6	-.073	.034	-.0054	-.046	.035	-.0032	-.074	.034	-.0054
Male kids	-.087	.035	-.0064	-.081	.035	-.0057	-.089	.035	-.0066
Premarriage kids	.231	.046	.0195	.199	.046	.0158	.232	.046	.0196
Age at marriage	-.033	.004	-.0024	-.023	.004	-.0016	-.033	.004	-.0024
Age gap	.005	.005	.0004	.017	.006	.0012	.005	.005	.0004
School gap	-.014	.009	-.0011	-.011	.009	-.0008	-.014	.009	-.0010
Cohab. w/ spouse	.033	.033	.0025	.040	.033	.0029	.033	.033	.0025
Cohab. with other	.229	.098	.0206	.191	.102	.0159	.234	.097	.0211
Black	-.055	.065	-.0039	-.159	.068	-.0103	-.052	.065	-.0037
Hispanic	.028	.063	.0021	-.108	.064	-.0072	.031	.063	.0023
Baptist	.035	.043	.0026	-.001	.042	-.0001	.034	.043	.0025
Catholic	-.047	.039	-.0034	-.038	.039	-.0026	-.048	.039	-.0035
Other religion	.058	.049	.0044	.036	.051	.0026	.058	.049	.0045
Live with mom	.058	.044	.0044	.060	.044	.0044	.059	.044	.0045
Live w/ mom/step	.187	.051	.0159	.192	.052	.0157	.185	.051	.0158
Live without mom	.117	.055	.0094	.096	.054	.0072	.119	.055	.0096
Traditional views	-.091	.033	-.0064	-.089	.033	-.0060	-.090	.033	-.0064
Environmental									
No fault	.126	.047	.0096	.106	.046	.0076	.126	.047	.0096
Property no fault	.026	.038	.0019	.073	.039	.0052	.026	.038	.0019
Separation dur.	.005	.002	.0003	.005	.002	.0004	.005	.002	.0003
Cty unemp. rate	-.004	.005	-.0003	-.011	.005	-.0007	-.005	.005	-.0003
Cty divorce rate	.008	.008	.0006	.009	.008	.0006	.009	.008	.0007
County race	-.000	.001	-.0000	-.001	.001	-.0001	-.000	.001	-.0000
County male	-.019	.012	-.0014	-.020	.012	-.0014	-.018	.012	-.0013
Other variables[b]									
Constant	-.230	.600	—	-.320	.584	—	-.240	.596	—
Log likelihood	-6065.282			-5944.188			-6068.289		

[a]Total income, income share, and income correlation are based on predicted income rather than actual income.

[b]Each specification also includes controls for current marital duration (age).

Note: The sample consists of 38,733 observations for 3,214 women. Standard errors account for nonindependence across observations for a given woman. Marginal effects are computed at the gender-specific sample means.

Table 4B: Probit Estimates of Effects of Variables on Probability of Divorce (Men)

Variable	Specification 1			Specification 2[a]			Specification 3		
	Coeff.	S.E.	Marg.	Coeff.	S.E.	Marg.	Coeff.	S.E.	Marg.
Risk Tolerance	.019	.015	.0013	.023	.015	.0015	—	—	—
Income variables									
Assets/100	-.022	.008	-.0014	-.008	.008	-.0005	-.022	.008	-.0014
Total income/100	-.028	.052	-.0018	-.785	.102	-.0501	-.028	.052	-.0018
Income share	-.002	.001	-.0001	.004	.001	.0002	-.002	.001	-.0001
Income correlation	-.022	.027	-.0015	.098	.083	.0063	-.022	.027	-.0015
No correlation	2.514	.367	.6989	2.487	.357	.6845	2.520	.3670	.7010
Demographics									
Number of kids	-.140	.023	-.0093	-.147	.024	-.0094	-.141	.023	-.0093
Kids age 0-6	-.270	.038	-.0179	-.305	.039	-.0195	-.269	.038	-.0178
Male kids	-.067	.043	-.0044	-.062	.043	-.0039	-.067	.043	-.0045
Premarriage kids	.133	.071	.0099	.089	.071	.0061	.132	.071	.0098
Age at marriage	-.031	.005	-.0021	-.013	.005	-.0008	-.032	.005	-.0021
Age gap	.015	.005	.0010	.006	.005	.0004	.015	.005	.0010
School gap	-.021	.010	-.0014	-.017	.010	-.0011	-.021	.010	-.0014
Cohab. w/ spouse	.011	.033	.0008	.021	.034	.0014	.012	.033	.0008
Cohab. w/ other	.163	.096	.0125	.113	.097	.0080	.164	.096	.0126
Black	.108	.068	.0076	-.014	.072	-.0009	.107	.068	.0075
Hispanic	-.049	.067	-.0032	-.173	.070	-.0100	-.051	.067	-.0033
Baptist	.093	.046	.0064	.056	.046	.0037	.094	.046	.0065
Catholic	.084	.043	.0057	.092	.044	.0060	.087	.043	.0059
Other religion	.083	.055	.0059	.067	.056	.0045	.085	.055	.0060
Live with mom	.057	.042	.0039	.045	.042	.0030	.055	.042	.0038
Live w/ mom/step	.134	.056	.0100	.118	.056	.0083	.135	.056	.0101
Live without mom	.075	.058	.0053	.056	.059	.0037	.073	.058	.0052
Traditional views	.001	.033	.0001	-.041	.033	-.0026	.003	.032	.0002
Environmental									
No fault	.071	.046	.0048	.072	.046	.0047	.073	.046	.0049
Property no fault	.038	.039	.0025	.053	.040	.0034	.039	.039	.0026
Separation dur.	.000	.002	.0000	.001	.002	.0001	.000	.002	.0000
Cty unemp. rate	-.000	.005	-.0000	-.009	.005	-.0006	-.000	.005	-.0000
Cty divorce rate	.033	.008	.0022	.035	.009	.0022	.033	.009	.0022
County race	.000	.001	.0000	-.001	.001	-.0001	.000	.001	.0000
County male	-.035	.017	-.0023	-.041	.017	-.0026	-.035	.017	-.0023
Other variables[b]									
Constant	.686	.807	—	.780	.822	—	.734	.810	—
Log likelihood	-5613.018			-5543.940			-5614.212		

[a]Total income, income share, and income correlation are based on predicted income rather than actual income.

[b]Each specification also includes controls for current marital duration (age).

Note: The sample consists of 37,662 observations for 3,298 men. Standard errors account for nonindependence across observations for a given man. Marginal effects are computed at the gender-specific sample means.

Table 5A: Predicted Divorce Probabilities for Representative Women
With Varying Levels of Risk Tolerance (based on specification 1 in table 4A)

Sample used[a]	Current duration of marriage	Level of Risk Tolerance				
		Min: 0.08[b]	p_{10}: 0.11[b]	Med: 0.34[b]	p_{90}: 1.03[b]	Max: 5.50[b]
Women	0-12 months	.014 (.002)	.014 (.002)	.014 (.002)	.016 (.003)	.027 (.008)
Women	13-24 months	.033 (.004)	.033 (.005)	.034 (.005)	.037 (.005)	.059 (.015)
Women	25-36 months	.040 (.005)	.040 (.005)	.041 (.005)	.044 (.006)	.069 (.017)
Women	37-48 months	.034 (.005)	.034 (.005)	.035 (.005)	.038 (.005)	.060 (.016)
Women	49-60 months	.031 (.005)	.031 (.004)	.032 (.005)	.034 (.005)	.055 (.015)
Women	61-72 months	.025 (.004)	.025 (.004)	.026 (.004)	.028 (.004)	.046 (.013)
		Min: .07[c]	p_{10}: 0.12[c]	Med: 0.39[c]	p_{90}: 1.32[c]	Max:9.23[c]
Women	37-48 months	.034 (.005)	.034 (.005)	.035 (.005)	.039 (.005)	.086 (.033)
Women+Men	37-48 months	.036 (.005)	.036 (.005)	.038 (.005)	.042 (.006)	.084 (.033)

[a]A representative woman is assigned mean values for all continuous variables (except risk tolerance) and modal values for all discrete variables (except duration), where means and modes are based on a sample of *women* or *women and men* in the given duration category.

[b]Levels of risk tolerance are the minimum, 10th percentile, median, 90th percentile and maximum values for the total sample of women (38,733 observations).

[c]Levels of risk tolerance are the minimum, 10th percentile, median, 90th percentile and maximum values for the pooled sample of women and men (76,395 observations).

Note: Standard errors are in parentheses.

Table 5B: Predicted Divorce Probabilities for Representative Men
With Varying Levels of Risk Tolerance (based on specification 1 in table 4B)

Sample used[a]	Current duration of marriage	Level of Risk Tolerance				
		Min: 0.07[b]	p_{10}: 0.13[b]	Med: 0.44[b]	p_{90}: 1.64[b]	Max: 9.23[b]
Men	0-12 months	.008 (.002)	.008 (.002)	.008 (.002)	.009 (.002)	.013 (.005)
Men	13-24 months	.036 (.005)	.036 (.005)	.037 (.005)	.039 (0.06)	.053 (.016)
Men	25-36 months	.035 (.005)	.035 (.005)	.036 (.005)	.038 (.006)	.052 (.016)
Men	37-48 months	.022 (.004)	.022 (.004)	.023 (.004)	.024 (.004)	.034 (.011)
Men	49-60 months	.021 (.004)	.021 (.004)	.022 (.004)	.023 (.004)	.032 (.011)
Men	61-72 months	.022 (.004)	.022 (.004)	.023 (.004)	.024 (.004)	.034 (.011)
		Min: .07[c]	p_{10}: 0.12[c]	Med: 0.39[c]	p_{90}: 1.32[c]	Max:9.23[c]
Men	37-48 months	.022 (.004)	.022 (.004)	.023 (.004)	.024 (.004)	.034 (.011)
Women+Men	37-48 months	.025 (.004)	.025 (.004)	.025 (.004)	.026 (.005)	.037 (.013)

[a]A representative an is assigned mean values for all continuous variables (except risk tolerance) and modal values for all discrete variables (except duration), where means and modes are based on a sample of *men* or *women and men* in the given duration category.

[b]Levels of risk tolerance are the minimum, 10th percentile, median, 90th percentile and maximum values for the total sample of men (37,662 observations).

[c]Levels of risk tolerance are the minimum, 10th percentile, median, 90th percentile and maximum values for the pooled sample of women and men (76,395 observations).

Note: Standard errors are in parentheses.