DIVORCE AS RISKY BEHAVIOR*

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Given that divorce often represents a high-stakes income gamble, we ask how individual levels of risk tolerance affect the decision to divorce. We extend the orthodox divorce model by assuming that individuals are risk averse, that marriage is risky, and that divorce is even riskier. The model predicts that conditional on the expected gains to marriage and divorce, the probability of divorce increases with relative risk tolerance because risk averse individuals require compensation for the additional risk that is inherent in divorce. To implement the model empirically, we use data for first-married women and men from the 1979 National Longitudinal Survey of Youth to estimate a probit model of divorce in which a measure of risk tolerance is among the covariates. The estimates reveal that a 1-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 notion that divorce entails a greater income gamble for women than for men.

"Better the devil you know than the devil you don't."

—English proverb

For many individuals, divorce is a high-stakes gamble. The gains that they receive by remaining married are far from certain, given that future income, asset values, and non-pecuniary rewards (including love) are susceptible to random shocks. Nonetheless, the value of a current marriage can appear to be a "sure bet" compared with 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, growth of their own labor market earnings, and luck in the remarriage market—all of which are subject to considerable randomness. Although 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, Landes, and Michael (1977), we assume that 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 the risk premium. In fact, we assume that *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 instead of marriage increases monotonically in an Arrow-Pratt index of risk aversion. Simply put, a risk averse individual is predicted to be less likely to divorce than is a more risk tolerant counterpart.

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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 lifetime income gambles. We also control for an array of variables intended to measure the economic gains to marriage and divorce, including characteristics of marriage markets and U.S. state laws governing divorce.

Our estimates reveal that risk preference plays an important role in the decision to divorce, especially for women. For example, among representative women in their fourth year of marriage, a 1-point (1.8 standard deviation [SD]) increase in risk tolerance raises the predicted probability of divorce by 11.4%. Among representative men with the same marriage duration, an identical 1-point increase in risk tolerance (which equals 1 SD 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 and that divorce entails a greater income gamble for women than for men.

Although risk and uncertainty are central to many 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 (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 have considered how individual variation in risk aversion comes into play; they argued that risk sharing motives 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) argued 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 the only empirical studies to use a measure of individual risk preference as a determinant of marital transitions, Schmidt (2008) and Spivey (2010) found that the waiting time to marriage decreases with risk aversion, presumably because risk averse individuals attach less value to continued search and/or place more value on the risk pooling gains to marriage.

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 *all* respondents' risk preferences while also supporting detailed analyses of transitions into and out of marriage. During three of the 22 interviews conducted through 2006, NLSY79 respondents were asked whether they would accept two hypothetical, lifetime 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. Although identical income gamble questions were included in multiple rounds of the Health and Retirement Study, that survey's focus on older individuals makes it less appropriate for an analysis of divorce. The questions were also included in the Panel Study of Income Dynamics but were asked only of *employed* respondents in a *single* interview year. In 2004, the German Socio-Economic Panel Study

1. Details on the income gamble questions are provided in the data section. The design and validity of these questions—which originated in the Health and Retirement Study—are discussed in Barsky et al. (1997) and Kimball, Sahm, and Shapiro (2008).

(GSOEP) 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 GSOEP has followed a large, representative sample of individuals for more than 20 years—and collected detailed information on labor market activities and family formation—it is a viable non-U.S. alternative to the NLSY79 for an analysis of the effects of risk preference on marital dissolution.[2]

THE DECISION TO DIVORCE

We rely on the simple, canonical discrete choice model that is often used to identify determinants of divorce (Becker et al. 1977; Charles and Stephens 2004; Hoffman and Duncan 1995; Weiss and Willis 1997). The model is based on the assumption that couples marry because they expect marriage to bring them higher utility than the alternative states, and subsequently divorce when new information causes them to change their assessment of the relative gains to marriage. This view of the decision process leads to the estimation of a sequential, discrete choice model with proxies for the expected gains to marriage and divorce as regressors. Analysts have relied on a range of theoretical ideas (e.g., intrahousehold specialization, consumption smoothing, bargaining, and marital search) to justify their choice of covariates, but they have only considered the case where decision makers are risk neutral.[3]

In this section, we demonstrate how risk aversion is likely to affect the divorce decision. In the unrealistic case in which the value of continued marriage is known with certainty and only divorce entails a risk, well-known principles of utility theory apply: whereas risk neutral individuals divorce whenever the expected utility of divorce exceeds the known utility of continued marriage, risk averse individuals require the expected utility of divorce to exceed the utility of marriage by "enough" to compensate for the risk (Pratt 1964). We devote the first subsection to formalizing the choice model and showing that this well-known risk premium argument also applies to the case in which marriage is risky but divorce is riskier.

Although the risk premium argument provides an intuitive interpretation of the effect of risk tolerance on the decision to divorce, in the second subsection of this article, we consider alternative interpretations. Specifically, we discuss ways in which an individual's level of risk tolerance might be determined by, or otherwise related to, expected gains to marriage and divorce. If we fail to control fully for these gains, then a positive link between risk tolerance and divorce could reflect the fact that highly risk tolerant individuals gain relatively less from marriage than do their more risk averse counterparts.

Effects of Risk Tolerance on the Choice Between Two Risky Options

Let $M_{it} = M(X_{it}, X^h_{it}, X^c_{it}, \varphi^c_{it})$ be the lifetime consumption that individual i receives if she remains married from time t until 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^h_{it}); tangible factors, such as joint financial assets that characterize the couple (X^c_{it}); and intangible characteristics of the marriage, such as love (φ^c_{it}). The lifetime consumption that the woman receives if she instead divorces at time t is $D_{it} = D(X_{it}, X^h_{it}, X^c_{it}, Z_{it})$, where Z_{it} represents current and future divorce costs and characteristics of the marriage market. The value of divorce includes X^h_{it} and X^c_{it} insofar as financial components of these

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.

3. Becker et al. (1977:1143) implicitly acknowledged that marriage is risky when they claimed 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 did not explicitly consider the risky nature of divorce and, in fact, appeared to assume that agents are risk neutral.

vectors affect property settlements, alimony, and child support, while components such as children affect the indirect costs of divorce.

Each individual has an increasing, concave utility function $U(C_{it})$ defined over consumption that implies an Arrow-Pratt measure of relative risk tolerance, $\rho_{it} = -U' / C_{it}U''$. We assume that relative risk tolerance (the inverse of relative risk aversion) varies across individuals and ranges from zero to infinity. Although we assume that everyone is risk averse, the limiting case is an individual who is neutral toward risk and would need no premium to accept the riskier of two options with equal expected payoffs.

Before turning to the case in which both marriage and divorce involve risk, we consider the decision-rules that maximize (expected) utility in 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 $\pi_{it} \geq 0$ is the risk premium that 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) established that under this scenario, π_{it} decreases monotonically with the index of relative risk tolerance, 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}.

Having established the role of risk preference when divorce is the only risky option, we turn to the scenario that more accurately describes the divorce decision: marriage is risky, and divorce is even riskier. We assume that 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, but 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 it depends on which second husband (if any) is selected as well as realizations of his characteristics over time. Second, while X^h_{it} and X^c_{it} 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 divorce.

Women are typically more dependent than men on spousal income, alimony, child support, and other income sources that become more uncertain upon divorce (Bianchi, Subaiya, and Kahn 1999; Cancian, Danziger, and Gottschalk 1993; Light 2004). Thus, the arguments in the preceding paragraph imply that divorce entails a greater income gamble for women than for men.[4] Although it seems noncontroversial to assume that divorce is riskier for women than for men, on average, we expect the risk to differ among individuals of a given gender. For example, women with high earnings potential, no children, and/or explicit prenuptial agreements are likely to face relatively little uncertainty about their consumption if they divorce. Our empirical model identifies the effects of M_{it}, D_{it}, and ρ_{it} on the probability of divorce at the *mean* level of unobserved risk.

To demonstrate that the risk premium argument continues to apply 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 X_{it}, X^h_{it}, X^c_{it}, φ_{it}, and Z_{it} evolve over time, we simply assume 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

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

4. We do not attempt to measure the riskiness of divorce by comparing actual pre- and post-divorce income because risk is based on *ex ante* assessments and not *ex post* realizations. Stated differently, whether a woman ultimately "wins" or "loses" the income gamble inherent in divorce is not an indication of the risk she faced.

As Chateauneuf, Cohen, and Meilijson (2004) demonstrated, an alternative definition is that F_D single-crosses F_M such that the (negative) horizontal distance $F^{-1}_D(c) - F^{-1}_M(c)$ is nondecreasing in every interval below the crossing point.

Location-independent risk is the most general stochastic order to *guarantee* 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] Location-independence is a plausible distributional assumption, but it does not have to hold for the risk premium argument to apply. Alternative definitions of "riskier" might entail deviations from monotonicity yet still yield a negative correlation between π_{it} and ρ_{it}.

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 that the woman is willing to pay to draw $D_{it} - \pi_{it}$ from the less-risky distribution F_M rather than face the riskier divorce outcome. 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.

Effects of Risk Tolerance on the Gains to Marriage and Divorce

The preceding discussion provides a familiar 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. A woman's risk preference can also affect (or be correlated with) her search for a husband both before and after her current marriage, the extent to which she engages in within-household risk sharing, and her bargaining power. Matching, risk sharing, and bargaining contribute to the relative gain associated with marriage which, in turn, affects the probability of divorce. In this subsection, we consider how risk preference might affect the probability of divorce through these additional channels.

Consider a situation in which 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 originated in the job search literature (Pissarides 1974) and has been applied to marital search (Schmidt 2008; Spivey 2010), suggests that components of M_{it} increase in ρ_{it}. In contrast to the prediction emerging from the risk premium framework, we might find that the probability of divorce *decreases* in ρ_{it} to the extent that ρ_{it} is positively correlated with unmeasured components of M_{it}.

This naive prediction does not necessarily hold after we acknowledge that cohabitation is another option available to single women. As shown 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. Thus, 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 relaunch the search process after a divorce, any relationship between risk preference and M_{it} can also exist between risk preference and D_{it}. In short, search models suggest ways in which ρ_{it} might be correlated

5. Ross (1981) demonstrated 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 guarantees 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 definition of riskiness proposed by Bickel and Lehmann (1979).

with M_{it} and D_{it} but do not yield an unambiguous prediction about the effect of risk tolerance on divorce.

Risk sharing provides another mechanism by which a woman's risk preference can 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), one prediction is that a highly risk averse woman derives a higher level of expected utility from marriage than does a more risk tolerant woman (Schmidt 2008; Spivey 2010). However, Chiappori and Reny (2006) argued 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. Only to the extent that couples fail to sort on risk preference would we expect unobserved elements of M_{it} that represent intrahousehold risk sharing to be correlated with ρ_{it}.[6]

More generally, any factor that (1) affects the probability of divorce, (2) is left unmeasured in our empirical choice model, and (3) is correlated with ρ_{it} can affect our inferences about the relationship between risk tolerance and divorce decisions. Many "errors in variables" interpretations are immediately undermined by the fact that we estimate markedly different relationships for women than for men; although this fact is consistent with the risk premium argument, an alternative interpretation requires a gender difference in the omitted variable's effect on divorce decisions and/or correlation with risk tolerance. By invoking bargaining models of marriage (Lundberg and Pollak 1994, 1996; Manser and Brown 1980; McElroy and Horney 1981), for example, we could argue that highly risk tolerant women (but not men) succeed in allocating marital gains toward themselves. Although we have no *a priori* reason to believe that bargaining power is systematically related to risk preference, bargaining power is a prime example of a factor that affects divorce decisions and is controlled for imperfectly.

ESTIMATION OF THE DIVORCE MODEL

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

$$S_{it} = \beta_1\ \rho_{it} + \beta_2\ X_{it} + \beta_3\ X^h_{it} + \beta_4\ X^c_{it} + \beta_5\ \varphi^c_{it} + \beta_3\ Z_{it} + \varepsilon_{it}, \tag{2}$$

where ρ_{it} continues to represent the Arrow-Pratt coefficient of relative risk tolerance, and X_{it}, X^h_{it}, X^c_{it}, φ^c_{it}, and Z_{it} represent the factors described earlier. The risk premium argument 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, correlations between ρ_{it} and unmeasured components of M_{it} and D_{it} can also affect our estimate of β_1. We do not expect these indirect effects to be systematically positive or negative, but they exist if our detailed set of covariates fails to control adequately for the gains to marriage.

In Eq. (2), ε_{it} represents unobserved factors that influence the probability of divorce. We assume that ε_{it} is a normally distributed random variable and that, conditional on the control variables, ε_{it} has a zero mean and constant variance. Given these assumptions, 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.

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—but this factor is readily measured.

DATA

Sample Selection

Our primary data source is the 1979 National Longitudinal Survey of Youth (NLSY79). The original NLSY79 sample consists of a nationally representative subsample of 6,111 individuals born between 1957 and 1964; an oversample of 5,295 blacks, Hispanics, and disadvantaged nonblacks/non-Hispanics born between 1957 and 1964; and a sample of 1,280 individuals born between 1957 and 1961 who enlisted in the military. All 12,686 sample members were interviewed in 1979, and subsequent interviews were conducted every year through 1994 and biennially thereafter. We use data from survey years 1979–2004. The military subsample was dropped from the survey in 1985, and the disadvantaged oversample was 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.

We construct separate samples of men and women. Of the 6,283 female and 6,403 male respondents in the NLSY79, we omit 1,093 women and 512 men because they marry prior to their first interview in 1979. We impose this selection rule because a subset of covariates cannot be identified for in-progress marriages. We eliminate 1,108 (1,704) of the remaining women (men) because they remained "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.[7] We omit 850 of these women and 866 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 left the survey prior to 1993. Finally, we omit 18 women and 23 men because their marriage began the same year as their last interview, which prevents us from observing the marriage over at least one 12-month interval. These selection rules leave us 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 women and 37,662 person-year observations for men. 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 (e.g., 1995, 1997) when NLSY79 respondents were not interviewed by imputing values for select time-varying covariates from information reported during adjacent interviews. Each marriage contributes between 1 and 24 observations to the sample, with a mean of 12.1 observations per marriage (SD = 6.9).

Measuring Risk Tolerance

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

7. We focus on first marriages to simplify the analysis. Individuals in later marriages differ systematically from the "first married" in terms of age, the presence of children, asset levels, and levels of relative risk tolerance. It is likely that they also differ in important unobserved dimensions, such as match quality. We would not want to ignore these differences by combining all marriages into a single sample, yet exploring the differences goes beyond the scope of the current study.

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%.

We form a four-way, ordinal ranking based on individuals' direct responses to the 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-half 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 *RT* is a cardinal measure of risk that can be compared in a meaningful fashion across individuals, and is inversely related to the risk premium described earlier. (We also use the four categorical variables as an alternative measure of risk tolerance.) To compute *RT*, we modify the estimation procedure proposed by Barsky et al. (1997) to incorporate the multiple responses to the income gamble questions 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. Ahn (2010), Kimball et al. (2008), and Sahm (2007) described variants of the computational method.

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

$$U(C_i) = \frac{C_i^{1-1/\rho_{it}}}{1-1/\rho_{it}}, \tag{3}$$

where ρ_{it} is the coefficient of relative risk tolerance for individual i at time t. We can infer lower and upper bounds for each individual's ρ_{it} from categorical responses to the income gamble questions. For example, if a respondent accepts the first gamble (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:

$$\frac{1}{2}U(2I)+\frac{1}{2}U\left(\frac{2}{3}I\right)\geq U(I) \text{ and } \frac{1}{2}U(2I)+\frac{1}{2}U\left(\frac{1}{2}I\right) < U(I). \tag{4}$$

Given our parameterization of the utility function, we infer that this individual's ρ_{it} lies between 0.5 and 1.0.[8]

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

$$\log \rho_{it} = \beta \text{AGE}_{it} + \alpha_i + u_{it}, \tag{5}$$

where $\alpha_i \sim N(\overline{\alpha}, \sigma_u^2)$ and $u_{it} \sim N(0, \sigma_u^2)$. We allow variation in $\log \rho_{it}$ to depend on age and unobserved random effects that we decompose into person-specific factors (α_i) and time-varying factors such as measurement error (u_{it}). Although risk preference is often viewed as an innate, time-invariant characteristic, we include age in our model in light of evidence presented in Ahn (2010) and Sahm (2007) that individuals tend to grow more risk averse with age; we present our own evidence of this pattern in Table 2. Moreover, we estimate Eq. (5) separately for men and women in light of the evidence that women are more risk averse than men (Sahm 2007)

8. 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,∞), respectively.

Table 1. Maximum Likelihood Estimates of Risk Preference Parameters

	Women		Men	
Parameter	Estimate	SE	Estimate	SE
β	–0.035	0.003	–0.047	0.003
$\bar{\alpha}$	–0.178	0.104	0.525	0.114
σ_α	1.277	0.029	1.390	0.032
σ_u	1.131	0.019	1.222	0.021
Log-Likelihood	–13,746.41		–13,599.71	
Number of Individuals	4,618		4,577	
Number of Observations	12,481		11,903	

Note: Parameters are for Eq. (5).

Given the parameterization shown in Eq. (5), we construct a log-likelihood function that depends on the data (age, risk category) and parameters β, $\bar{\alpha}$, σ_α, and σ_u. We compute gender-specific maximum likelihood estimates of each parameter—using data for *all* men and women in the NLSY79 who answer the income gamble questions, regardless of whether they appear in our sample—and use these estimates to calculate each individual's expected ρ_{it} at each age that her marriage is observed. The maximum likelihood estimates appear in Table 1; details on the procedure appear in the appendix. These conditional expectations form the variable *RT,* which represents each individual's coefficient of relative risk tolerance in each year.[9]

To substantiate our claim that risk preferences change over time, in Table 2 we compare individuals' first and second responses to the income gamble questions. Almost 90% of the individuals in our samples provided 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. Table 2 reveals that roughly one-half of the women in our sample fall 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, 53% reported a lower risk tolerance the second time, while only 28% reported a higher level; for those initially in Category 3, 61% reported a lower level and 17% reported a higher level the second time. These patterns reveal why we model log ρ_{it} as a function of age: although much of the within-person variation in risk category can be attributed to reporting error, women also become less risk tolerant as they age.

Table 2 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. On the basis of their first responses, men are 5 percentage points less likely than women to fall into Category 1 (44.7% versus 49.8%) and 6 percentage points more likely to fall into Category 4 (26.1% versus 19.5%). Among men whose first response places

9. We do not correct the standard errors in our probit model for sampling variation in *RT* because to our knowledge, analytic methods for computing standard errors for two-step models (e.g., Murphy and Topel 1985) cannot be extended to our model. Using bootstrap methods for a similar model, Kimball et al. (2008) demonstrated that sampling variance in the computed regressor has little effect on estimated standard errors and does not alter their inferences regarding statistical significance.

Table 2. Distribution of Risk Category Based on Second Response by Risk Category Based on First Response

Risk Category, First Response	Women: Risk Category, Second Response					Men: 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]
Number of Individuals	1,686	334	435	422	2,887	1,573	316	482	547	2,918

Notes: 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. Categories 1 and 4 consist, respectively, of the least and most risk tolerant individuals. Numbers in brackets are percentages of column totals, and all other numbers are percentages of row totals.

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 3 summarizes the distribution of the variable *RT* among women and men in each self-reported risk category; for this table, we use the value of *RT* for the year corresponding to the individual's first response to the income gamble questions. Although both the mean and median of *RT* increase monotonically with the risk category, as expected, considerable variation exists in *RT* 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 (see footnote 8). 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 *RT*. As shown in Table 1, the two estimated error variances in our log ρ_{it} model are roughly equal in magnitude for men and women, which suggests that one-half of the total variation is attributable to error.

Other Covariates

We use many variables to control for the expected gains to marriage and divorce. These variables are intended to capture heterogeneity in match quality, marriage-specific capital, intrahousehold specialization, intrahousehold consumption smoothing, bargaining power, divorce costs, remarriage opportunities, and attitudes toward marriage and divorce. To organize our discussion, we group the variables into economic measures, demographic and family background characteristics, and environmental factors. Summary statistics are in Table 4.

Our economic variables include a measure of each individual's net family assets. We construct this variable by summing the values of homes, automobiles, 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.[10] We include net assets in our model to capture the value of

10. We impute values when a respondent says that she has a particular asset or debt but does not report its value. If the item's value is reported in an earlier and later interview, we use the closest-reported values to linearly

Table 3. Summary Statistics for Risk Tolerance Variable by Risk Category Based on First Response

Risk Category, First Response[a]	Risk Tolerance (RT)[b] Mean	SD	Minimum	Median	Maximum	Number
Women						
1	0.25	0.18	0.11	0.20	1.06	1,601
2	0.46	0.25	0.21	0.42	1.43	430
3	0.60	0.30	0.26	0.55	1.76	556
4	1.15	0.85	0.38	0.99	3.67	628
All	0.51	0.56	0.11	0.38	3.67	3,214
Men						
1	0.32	0.26	0.11	0.26	1.31	1,480
2	0.59	0.37	0.23	0.49	1.80	366
3	0.74	0.44	0.29	0.60	2.25	582
4	1.65	1.30	0.43	1.28	5.42	870
All	0.77	0.90	0.11	0.47	5.42	3,298

[a]See note to Table 2 for variable definition.

[b]Computed relative risk tolerance for the same year as the first response to the income gamble questions.

marriage-specific capital and public goods that increase the gains to marriage and lower the probability of divorce.

Our covariates include four income measures. First, we control for the couple's total labor income, which is the sum of the partners' wage and salary income in the last year. This variable is intended to capture such gains to marriage as the joint consumption associated with income (Moffitt 2000). 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). 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 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 spouses' labor income to measure the extent of intrahousehold income risk sharing. A couple with negatively correlated incomes are best able to exploit the risk sharing advantages of marriage and, as a result, are less likely to divorce. 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.[11]

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 the total net worth variable in the NLSY79, although we take the additional step of imputing values for 1979–1984.

11. We experimented with 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 or a discernible effect on the estimated coefficient for RT.

Table 4. Definitions and Summary Statistics for Variables Used in Divorce Model

Variable	Definition	Women Mean	Women SD	Men Mean	Men SD
Divorce	1 if divorces during interval	0.04		0.04	
Risk Tolerance	Arrow-Pratt coefficient of relative risk tolerance	0.51	0.56	0.75	0.95
Economic 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,b]	0.08	0.54	0.06	0.58
No correlation	1 if income correlation is missing[b]	0.01		0.01	
Predicted total income	Predicted total income[a]	53.19	24.94	49.25	24.32
Predicted income share	Predicted own share of total income[a]	34.98	17.17	71.30	17.99
Predicted income correlation	Predicted income correlation[a,b]	0.09	0.17	0.06	0.18
Demographic Variables					
Number of children	Number of children in household	1.42	1.19	1.35	1.20
Children aged 0–6	1 if any children age 6 or younger	0.50		0.48	
Male children	1 if any male children	0.53		0.49	
Premarriage children	1 if any children born before marriage[b]	0.16		0.06	
Age at marriage	Age at marriage[b]	23.77	4.35	25.02	4.21
Age gap	Difference in spouses' ages[b]	3.43	3.61	2.78	2.77
Schooling gap	Difference in spouses' years of school[b]	1.42	1.62	1.31	1.52
Cohabited with spouse	1 if cohabited with spouse[b]	0.31		0.35	
Cohabited with other	1 if cohabited with other partner[b]	0.02		0.02	
Black	1 if black[b]	0.24		0.23	
Hispanic	1 if Hispanic[b]	0.19		0.20	
Baptist	1 if religion is Baptist[b]	0.26		0.25	
Catholic	1 if religion is Catholic[b]	0.39		0.38	
Other religion	1 if no or other religion[b]	0.10		0.11	
Lived with mother	1 if lived with mother only, age 14[b]	0.14		0.14	
Lived with mother/stepfather	Lived with mother/stepfather, age 14[b]	0.06		0.06	
Lived without mother	No mother, age 14[b]	0.07		0.07	
Traditional views	1 if agrees women are happier at home[b]	0.27		0.37	

(continued)

(Table 4, continued)

Variable	Definition	Women Mean	Women SD	Men Mean	Men SD
Environmental Variables					
No-fault	1 if no-fault law for divorce[c]	0.35		0.35	
Property no-fault	1 if no-fault law for property settlement[c]	0.42		0.42	
Separation duration	Minimum required separation (months)[c]	10.59	11.53	10.54	11.50
County unemployment rate	County unemployment rate	6.90	3.16	6.77	3.04
County divorce rate	County divorce rate	4.91	1.96	4.93	1.89
County race	Percentage of county population same race	59.63	31.19	59.21	32.46
County male	Percentage of county population male	48.82	1.32	48.86	1.22
Number of Person-Year Observations		38,733		37,662	
Number of Individuals		3,214		3,298	

Note: The model also includes dummy variables identifying current marriage duration.

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

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

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

Because an individual's labor income can be endogenously determined by her beliefs about a future divorce (Greene and Quester 1982; Johnson and Skinner 1986; Stevenson 2007), we also use a specification that replaces the income variables with predicted versions. Replacing actual labor income with its predicted value is also in the spirit of Pollak's (2005) argument that wage rates are preferred to actual earnings as measures of bargaining power. We predict respondents' income with a cubic in age, three dummy variables for schooling attainment, age-adjusted Armed Forces Qualifications Test (AFQT) scores, number of children, occupation dummy variables, state dummy variables, and the median income in the county of residence in the given calendar year. To predict spousal income, we omit AFQT scores and assume that husbands and wives share race/ethnicity and state and county of residence to skirt the fact that this information is known only for respondents. We estimate separate, gender-specific predicting equations for blacks, Hispanics, and whites using observations for first marriages for all NLSY79 respondents, regardless of whether they appear in our sample. Following Hess (2004), we predict 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 6 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). We control for whether a male child resides in the household in light of empirical evidence that divorce is less likely and remarriage is more likely for women with sons (Lundberg and Rose 2003). The "premarriage children" variable indicates a lack of marital capital insofar as these 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 and 10 dummy variables that identify current marriage duration. These variables control for variation in current age as well; the addition of direct controls for age proved to have insignificant effects on the estimates. We also control for the absolute value of the difference in the husband's and wife's age, the absolute value of the difference in their highest education grade completed, and dummy variables identifying premarital cohabitation. Individuals who marry at relatively later ages may have decreased search costs (and, therefore, may have 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). Although numerous empirical studies have shown that premarital cohabitation is associated with increased divorce (Axinn and Thornton 1992; Brien, Lillard, and Stern 2006; Lillard, Brien, and Waite 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 individual is black or Hispanic, the religion in which the individual was raised (Baptist, Catholic, or other/none, with Protestant as the omitted group), and family composition at age 14. 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, Martin, and Sweet 1991; Lehrer and Chiswick 1993). As an additional measure of marriage-related attitudes, we use responses to a question that asked NLSY79 respondents whether they agree with the following statement: "Women are much happier if they stay at home and take care of their children." We construct a dummy variable that equals 1 if the respondent agreed or strongly agreed with the statement and 0 if she disagreed or strongly disagreed.

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 in the given year; the omitted category identifies states that either allow or require 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 0 if the state imposes no separation requirement.[12] 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 (Becker et al. 1977; Friedberg 1998; Mechoulan 2006; Peters 1986; Stevenson 2007), 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 characteristics of the individual's county of residence for the given year. These variables—from various issues of the City and County Data Book—include the county- and year-specific unemployment rate and divorce rate, the percentage of the county population with the same race/ethnicity (black, white, or Hispanic) as the individual, and the percentage of the county population that is male. Similar variables have been used by Lichter, McLaughlin, and Ribar (2002) as controls for economic opportunities and marriage market characteristics.

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

FINDINGS

Estimated Effects of Risk Tolerance

In Tables 5 and 6, we present estimates for women and men, respectively, for three specifications of our divorce model. Specification 1 includes all controls described earlier (with the income variables based on actual values rather than predictions), along with *RT*. Specification 2 uses predicted values for the four income variables but is identical to Specification 1 in all other respects. Specification 3 reverts to actual values of the income variables but omits *RT*.

Table 5 reveals that the estimated coefficient for *RT* in Specification 1 is large, positive, and precisely estimated for women; its estimated marginal effect is 0.0037 at the sample means. Table 6 reveals a much smaller and less precisely estimated effect for men, with an estimated marginal effect of 0.0013. Given the unconditional, 12-month divorce rate of 0.04, this means a 1-point increase in relative risk tolerance—which corresponds to 1.8 and 1.0 standard deviations for women and men, respectively—is predicted to increase the probability of divorce by 9.25% for women and 3.25% for men. The model includes a broad array of variables representing the gains associated with both marriage and divorce, including controls for risk sharing and match quality that may themselves be affected by an individual's level of risk tolerance. Thus, we believe a plausible interpretation of the estimated effect of *RT* is that it represents the risk premium associated with divorce. The predicted probability of divorce increases in *RT* because the more risk tolerant a person is, the smaller is her risk premium—that is, smaller is the amount by which the expected gains to divorce must exceed the expected gains to marriage to compensate her for the additional risk inherent in divorce. The fact that the estimated effect is considerably larger for women than for men is consistent with the view that women tend to face greater income risk than men upon divorcing because of women's reliance on income sources (husband's earnings, alimony, child support) that are highly unpredictable.

In Specification 2 in Tables 5 and 6, 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 *RT:* for women, the Specification 2 estimate (0.052) is 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 from zero at conventional significance levels.

Specification 3 in Tables 5 and 6 is identical to Specification 1 except that *RT* is excluded. We find that the estimated coefficients for all variables—including assets, income correlation, the individual'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 *RT*. Risk preference has a nontrivial effect on the probability of divorce, yet omitting *RT* from the model does not cause other observed factors to "absorb" its effect.[13]

To compare the estimated effect of risk preference to the effects of other key determinants of divorce, we focus on the Specification 1 estimates in Tables 5 and 6. For women, a 1-point (1.8 SD) increase in *RT* 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 SD) loss of assets, an 18–percentage-point (0.5 SD) increase in the woman's income share, a 1.5-year (0.4 SD) reduction in her age at marriage, or a 12-month (1.0 SD) increase in the

13. Moreover, a likelihood ratio test reveals that the exclusion of *RT* does not cause a substantial decrease in explanatory power. The same is true when we exclude other covariates (e.g., assets and premarital cohabitation) with statistically significant coefficients.

Table 5. Probit Estimates of Effects of Variables on Probability of Divorce: Women

Variable	Specification 1		Specification 2[a]		Specification 3	
	Coefficient	Marginal Effects	Coefficient	Marginal Effects	Coefficient	Marginal Effects
Risk Tolerance	0.050 (0.023)	0.0037	0.052 (0.022)	0.0037	—	—
Economic Variables						
Assets / 100	–0.042 (0.014)	–0.0031	–0.030 (0.012)	–0.0021	–0.042 (0.014)	–0.0031
Total income / 100	0.006 (0.050)	0.0005	–0.504 (0.091)	–0.0354	–0.007 (0.050)	–0.0005
Income share	0.003 (0.001)	0.0002	0.009 (0.001)	0.0006	0.003 (0.001)	0.0002
Income correlation	–0.035 (0.027)	–0.0026	0.189 (0.092)	0.0131	–0.033 (0.027)	–0.0024
No correlation	1.519 (0.174)	0.3362	1.486 (0.172)	0.3171	1.524 (0.174)	0.3383
Demographic Variables						
Number of children	–0.077 (0.020)	–0.0056	–0.031 (0.020)	–0.0021	–0.076 (0.020)	–0.0056
Children aged 0–6	–0.073 (0.034)	–0.0054	–0.046 (0.035)	–0.0032	–0.074 (0.034)	–0.0054
Male children	–0.087 (0.035)	–0.0064	–0.081 (0.035)	–0.0057	–0.089 (0.035)	–0.0066
Premarriage children	0.231 (0.046)	0.0195	0.199 (0.046)	0.0158	0.232 (0.046)	0.0196
Age at marriage	–0.033 (0.004)	–0.0024	–0.023 (0.004)	–0.0016	–0.033 (0.004)	–0.0024
Age gap	0.005 (0.005)	0.0004	0.017 (0.006)	0.0012	0.005 (0.005)	0.0004
Schooling gap	–0.014 (0.009)	–0.0011	–0.011 (0.009)	–0.0008	–0.014 (0.009)	–0.0010
Cohabited with spouse	0.033 (0.033)	0.0025	0.040 (0.033)	0.0029	0.033 (0.033)	0.0025
Cohabited with other	0.229 (0.098)	0.0206	0.191 (0.102)	0.0159	0.234 (0.097)	0.0211
Black	–0.055 (0.065)	–0.0039	–0.159 (0.068)	–0.0103	–0.052 (0.065)	–0.0037
Hispanic	0.028 (0.063)	0.0021	–0.108 (0.064)	–0.0072	0.031 (0.063)	0.0023
Baptist	0.035 (0.043)	0.0026	–0.001 (0.042)	–0.0001	0.034 (0.043)	0.0025
Catholic	–0.047 (0.039)	–0.0034	–0.038 (0.039)	–0.0026	–0.048 (0.039)	–0.0035
Other religion	0.058 (0.049)	0.0044	0.036 (0.051)	0.0026	0.058 (0.049)	0.0045
Lived with mother	0.058 (0.044)	0.0044	0.060 (0.044)	0.0044	0.059 (0.044)	0.0045

(continued)

(Table 5, continued)

Variable	Specification 1 Coefficient	Specification 1 Marginal Effects	Specification 2[a] Coefficient	Specification 2[a] Marginal Effects	Specification 3 Coefficient	Specification 3 Marginal Effects
Demographic Variables (cont.)						
Lived with mother/ stepfather	0.187 (0.051)	0.0159	0.192 (0.052)	0.0157	0.185 (0.051)	0.0158
Lived without mother	0.117 (0.055)	0.0094	0.096 (0.054)	0.0072	0.119 (0.055)	0.0096
Traditional views	–0.091 (0.033)	–0.0064	–0.089 (0.033)	–0.0060	–0.090 (0.033)	–0.0064
Environmental Variables						
No-fault	0.126 (0.047)	0.0096	0.106 (0.046)	0.0076	0.126 (0.047)	0.0096
Property no-fault	0.026 (0.038)	0.0019	0.073 (0.039)	0.0052	0.026 (0.038)	0.0019
Separation duration	0.005 (0.002)	0.0003	0.005 (0.002)	0.0004	0.005 (0.002)	0.0003
County unemployment rate	–0.004 (0.005)	–0.0003	–0.011 (0.005)	–0.0007	–0.005 (0.005)	–0.0003
County divorce rate	0.008 (0.008)	0.0006	0.009 (0.008)	0.0006	0.009 (0.008)	0.0007
County race	–0.000 (0.001)	–0.0000	–0.001 (0.001)	–0.0001	–0.000 (0.001)	–0.0000
County male	–0.019 (0.012)	–0.0014	–0.020 (0.012)	–0.0014	–0.018 (0.012)	–0.0013
Log-Likelihood	–6,065.282		–5,944.188		–6,068.289	

Notes: The sample consists of 38,733 observations for 3,214 women. Standard errors, shown in parentheses, account for nonindependence across observations for a given woman. Marginal effects are computed at the gender-specific sample means. Each specification also includes controls for current marital duration.

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

state's mandatory separation requirement. For men, a 1-point (1.0 SD) increase in *RT* generates a 3.25% increase in the predicted probability of divorce, as does a \$93,000 (0.4 SD) decrease in assets, a 13–percentage-point (0.6 SD) decrease in the man's income share, or a 0.6-year (0.2 SD) decrease in the age at marriage. For both women and men, the estimated effect of a 1-point increase in *RT* 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 several background 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. Although "traditional views" proves to have an unimportant effect on divorce probabilities for men, the absence of preschool children (ages 0–6) is

Table 6. Probit Estimates of Effects of Variables on Probability of Divorce: Men

	Specification 1		Specification 2[a]		Specification 3	
Variable	Coefficient	Marginal Effects	Coefficient	Marginal Effects	Coefficient	Marginal Effects
Risk Tolerance	0.019 (0.015)	0.0013	0.023 (0.015)	0.0015	—	—
Economic Variables						
Assets / 100	–0.022 (0.008)	–0.0014	–0.008 (0.008)	–0.0005	–0.022 (0.008)	–0.0014
Total income / 100	–0.028 (0.052)	–0.0018	–0.785 (0.102)	–0.0501	–0.028 (0.052)	–0.0018
Income share	–0.002 (0.001)	–0.0001	0.004 (0.001)	0.0002	–0.002 (0.001)	–0.0001
Income correlation	–0.022 (0.027)	–0.0015	0.098 (0.083)	0.0063	–0.022 (0.027)	–0.0015
No correlation	2.514 (0.367)	0.6989	2.487 (0.357)	0.6845	2.520 (0.367)	0.7010
Demographic Variables						
Number of children	–0.140 (0.023)	–0.0093	–0.147 (0.024)	–0.0094	–0.141 (0.023)	–0.0093
Children aged 0–6	–0.270 (0.038)	–0.0179	–0.305 (0.039)	–0.0195	–0.269 (0.038)	–0.0178
Male children	–0.067 (0.043)	–0.0044	–0.062 (0.043)	–0.0039	–0.067 (0.043)	–0.0045
Premarriage children	0.133 (0.071)	0.0099	0.089 (0.071)	0.0061	0.132 (0.071)	0.0098
Age at marriage	–0.031 (0.005)	–0.0021	–0.013 (0.005)	–0.0008	–0.032 (0.005)	–0.0021
Age gap	0.015 (0.005)	0.0010	0.006 (0.005)	0.0004	0.015 (0.005)	0.0010
Schooling gap	–0.021 (0.010)	–0.0014	–0.017 (0.010)	–0.0011	–0.021 (0.010)	–0.0014
Cohabited with spouse	0.011 (0.033)	0.0008	0.021 (0.034)	0.0014	0.012 (0.033)	0.0008
Cohabited with other	0.163 (0.096)	0.0125	0.113 (0.097)	0.0080	0.164 (0.096)	0.0126
Black	0.108 (0.068)	0.0076	–0.014 (0.072)	–0.0009	0.107 (0.068)	0.0075
Hispanic	–0.049 (0.067)	–0.0032	–0.173 (0.070)	–0.0100	–0.051 (0.067)	–0.0033
Baptist	0.093 (0.046)	0.0064	0.056 (0.046)	0.0037	0.094 (0.046)	0.0065
Catholic	0.084 (0.043)	0.0057	0.092 (0.044)	0.0060	0.087 (0.043)	0.0059
Other religion	0.083 (0.055)	0.0059	0.067 (0.056)	0.0045	0.085 (0.055)	0.0060
Lived with mother	0.057 (0.042)	0.0039	0.045 (0.042)	0.0030	0.055 (0.042)	0.0038

(continued)

(Table 6, continued)

Variable	Specification 1 Coefficient	Specification 1 Marginal Effects	Specification 2[a] Coefficient	Specification 2[a] Marginal Effects	Specification 3 Coefficient	Specification 3 Marginal Effects
Demographic Variables (cont.)						
Lived with mother/ stepfather	0.134 (0.056)	0.0100	0.118 (0.056)	0.0083	0.135 (0.056)	0.0101
Lived without mother	0.075 (0.058)	0.0053	0.056 (0.059)	0.0037	0.073 (0.058)	0.0052
Traditional views	0.001 (0.033)	0.0001	–0.041 (0.033)	–0.0026	0.003 (0.032)	0.0002
Environmental Variables						
No-fault	0.071 (0.046)	0.0048	0.072 (0.046)	0.0047	0.073 (0.046)	0.0049
Property no-fault	0.038 (0.039)	0.0025	0.053 (0.040)	0.0034	0.039 (0.039)	0.0026
Separation duration	0.000 (0.002)	0.0000	0.001 (0.002)	0.0001	0.000 (0.002)	0.0000
County unemployment rate	–0.000 (0.005)	–0.0000	–0.009 (0.005)	–0.0006	–0.000 (0.005)	–0.0000
County divorce rate	0.033 (0.008)	0.0022	0.035 (0.009)	0.0022	0.033 (0.009)	0.0022
County race	0.000 (0.001)	0.0000	–0.001 (0.001)	–0.0001	0.000 (0.001)	0.0000
County male	–0.035 (0.017)	–0.0023	–0.041 (0.017)	–0.0026	–0.035 (0.017)	–0.0023
Log-Likelihood	–5,613.018		–5,543.940		–5,614.212	

Notes: The sample consists of 37,662 observations for 3,298 men. Standard errors, shown in parentheses, account for nonindependence across observations for a given man. Marginal effects are computed at the gender-specific sample means. Each specification also includes controls for current marital duration.

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

predicted to raise a man's divorce probability by 45%, and 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 compare predicted divorce probabilities for representative women and men at various marriage durations and with various levels of risk tolerance. In the top rows of Table 7, we consider representative women for whom all variables other than *RT* and duration equal the mean (if continuous) or mode (if discrete) based on subsamples of women with the same marriage duration. Table 7 reveals that in the fourth year of marriage, a representative woman with risk tolerance at the 10th percentile (based on the *RT* distribution for all women) has a .034 probability of divorcing in the next year, and an otherwise identical woman with risk tolerance at the 90th percentile has a predicted probability (.038) that is 11.8% higher. At all marriage durations shown in Table 7, 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 any marriage duration beyond 12 months, the most risk tolerant

Table 7. Predicted Divorce Probabilities for Representative Women and Men

Sample Used[a]	Marriage Duration (months)	Level of Risk Tolerance (RT)[b]				
		Min. = .08	p_{10} = .11	Med. = .34	p_{90} = 1.03	Max. = 5.50
Women	13–24	.033 (.004)	.033 (.005)	.034 (.005)	.037 (.005)	.059 (.015)
Women	37–48	.034 (.005)	.034 (.005)	.035 (.005)	.038 (.005)	.060 (.016)
Women	61–72	.025 (.004)	.025 (.004)	.026 (.004)	.028 (.004)	.046 (.013)
		Min. = .07	p_{10} = .13	Med. = .44	p_{90} = 1.64	Max. = 9.23
Men	13–24	.036 (.005)	.036 (.005)	.037 (.005)	.039 (.006)	.053 (.016)
Men	37–48	.022 (.004)	.022 (.004)	.023 (.004)	.024 (.004)	.034 (.011)
Men	61–72	.022 (.004)	.022 (.004)	.023 (.004)	.024 (.004)	.034 (.011)
		Min. = .07	p_{10} = .12	Med. = .39	p_{90} = 1.32	Max. = 9.23
Women	37–48	.034 (.005)	.034 (.005)	.035 (.005)	.039 (.005)	.086 (.033)
Women + Men	37–48	.036 (.005)	.036 (.005)	.038 (.005)	.042 (.006)	.084 (.033)
Men	37–48	.022 (.004)	.022 (.004)	.023 (.004)	.024 (.004)	.034 (.011)
Women + Men	37–48	.025 (.004)	.025 (.004)	.025 (.004)	.026 (.005)	.037 (.013)

Note: Standard errors are in parentheses.

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

[b]The representative woman or man under consideration is assigned a level of *RT* equal to the minimum, 10th percentile, median, 90th percentile, and maximum values for the total sample of women (top rows), men (middle rows), or women plus men (bottom rows).

woman in our sample is roughly 80% more likely to divorce than her counterpart with the minimum level of risk tolerance. The middle rows of Table 7 reveal that the corresponding estimates are considerably smaller for men. Focusing on current durations of 37–48 months, the predicted probability of divorce increases by 9.1% (.022 to .024) 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. These are *smaller* than the corresponding numbers for women despite the fact that each movement within the gender-specific risk distribution represents a *larger* absolute change in risk tolerance for men than for women.

We have thus far assigned our representative individuals gender-specific values of each covariate, including *RT*. In the bottom rows of Table 7, we consider identical levels of risk preference for both men and women. We compute one set of predicted probabilities for a representative woman and a representative man who continue to possess gender-specific mean/modal characteristics, and another set for an identical woman and man, both of whom are assigned mean/modal characteristics based on a pooled sample of men and women. Focusing on the (roughly) 1-point increase in *RT* 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 1-point increase in *RT* (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, 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.

Robustness Checks

Although our risk premium interpretation of the role of *RT* appears plausible, the estimated effect can also reflect correlation between *RT* and unobserved components of the expected gains to marriage and divorce. Despite our inclusion of a rich set of proxies for the expected gains to marriage and divorce, components of these gains invariably go unmeasured. We cannot separate the "risk premium" effect of *RT* from the "errors in variables" effect, but in this subsection, we experiment with *measureable* factors to demonstrate that our findings are not influenced by correctable forms of model misspecification.

The invariance of each covariate's estimated coefficient to the exclusion of *RT* (seen by comparing Specifications 1 and 3 in Tables 5 and 6) suggests that the estimated effects of risk preference in Specifications 1 and 2 are unlikely to reflect simple forms of misspecification, such as the omission of higher-order terms in income or assets. To substantiate this, in Table 8 we show the estimated coefficients for *RT* for a range of alternative specifications; the first row of Table 8 provides the Specification 1 estimates from Tables 5 and 6 to facilitate comparison.

Specification 4 in Table 8 is identical to Specification 1 except for the addition of $Assets^2$, $Assets^3$, and $Assets^4$. Any suspicion that the estimated effect of *RT* reflects misspecification of the asset effect is dispelled by the finding that Specifications 1 and 4 yield virtually identical estimated coefficients for *RT.* The same is true when we instead introduce a quartic function of total income in Specification 5. When we introduce more flexibility in the effect of schooling attainment—either by distinguishing whether the individual's highest grade completed is greater than, equal to, or less than the spouse's (Specification 6) or by adding four dummy variables to identify the individual's highest grade completed (Specification 7)—the estimated coefficient for *RT* increases slightly for men. However, as with Specifications 4 and 5, we fail to reject the null hypothesis that the difference in parameter estimates is zero. These findings confirm that the estimated effect of *RT* does not reflect an overrestrictive parameterization of our divorce model.

We also ask whether the estimated effect of *RT* is sensitive to the presence of children in the household, given that levels of risk tolerance and the riskiness of divorce might be closely tied to child-rearing. To explore this issue, in Specification 8, we restrict the samples to childless individuals. For women, the estimated coefficient for *RT* in specification 8 is 0.050, which is identical to the Specification 1 estimate. For men, the estimated coefficient decreases from 0.019 to 0.004 (although the difference in point estimates is statistically insignificant), which is consistent with the notion that divorce is relatively low-risk for childless men.

In the final alternative specification presented in Table 8, we replace *RT* with the direct, categorical responses to the income gamble questions described earlier in the Data section. Although we prefer the computed variable *RT* to the raw responses for

Table 8. Probit Estimates of Effects of Risk Tolerance on Probability of Divorce, Based on Alternative Specifications

	Women			Men		
Specification	Coefficient	SE	Marginal Effect	Coefficient	SE	Marginal Effect
1. From Tables 5 and 6	0.050	0.023	0.0037	0.019	0.015	0.0013
4. Add Quartic in Assets[a]	0.049	0.023	0.0036	0.019	0.015	0.0013
5. Add Quartic in Total Income[a]	0.049	0.023	0.0036	0.019	0.015	0.0013
6. Add Dummy Variables for Schooling Gap[b]	0.049	0.023	0.0036	0.020	0.015	0.0013
7. Add Own Schooling Dummy Variables[c]	0.048	0.023	0.0036	0.023	0.015	0.0015
8. Childless Individuals Only[d]	0.050	0.043	0.0042	0.004	0.020	0.0005
9. Categorical Risk Variables[e]						
Risk Category 2	0.027	0.032	0.0024	0.028	0.036	0.0019
Risk Category 3	0.064	0.038	0.0048	–0.013	0.042	–0.0009
Risk Category 4	0.139	0.068	0.0114	0.124	0.052	0.0091

Note: The table shows estimated coefficients for *RT* in Specifications 1 and 4–8, and estimated coefficients for categorical risk variables in Specification 9.

[a]Specifications 4 and 5 add three higher-order terms in *Assets* and *Total income*, respectively.

[b]Replaces *Schooling gap* with three dummy variables indicating whether the individual's highest grade completed is less than, equal to, or greater than the spouse's.

[c]Adds four dummy variables indicating whether the individual's highest grade completed is 0–11, 12, 13–15, or 16 or more.

[d]Reduces samples to childless individuals; samples sizes are 10,813 for women and 11,971 for men.

[e]Replaces *RT* with dummy variables indicating whether average response to income gamble questions is Category 2, 3, or 4 (most risk tolerant); Category 1 (least risk tolerant) is the omitted group.

the reasons discussed in that section, it is important to establish that the positive relationship between *RT* and divorce probabilities does not arise solely from the procedure used to compute the variable.[14]

Table 8 reveals that for women, the estimated effects of risk tolerance are 0.027, 0.064, and 0.139 for Categories 2, 3, and 4, respectively. Although the estimated coefficient for Category 4 (which represents the highest level of risk tolerance) is the only one that is statistically different than zero at conventional significance levels, this specification reveals a positive relationship between categorical levels of risk tolerance and divorce probabilities that is qualitatively similar to what is found for specifications that include the computed variable *RT*. For men, we obtain small and imprecisely estimated coefficients for Categories 2 and 3, but a precisely estimated coefficient of 0.124 for Category 4. The estimates for Specification 9 support our conclusion that the relationship between risk tolerance and divorce probabilities is positive for everyone, but more pronounced and more precisely estimated for women than for men.

CONCLUSION

Beginning with Becker et al. (1977), researchers analyzing the decision to divorce have used a decision-making framework in which risk neutral agents choose the state (marriage

14. The level of each individual's computed *RT* depends on her responses to the income gamble questions in all three years (assuming the individual answered the questions in 1993, 2002, and 2004). For comparability, the categorical variables used in Specification 9 represent the individual's (rounded) average response over all three years. For example, a woman whose responses place her in Categories 4, 3, and 3 is assigned Category 3.

or divorce) that maximizes expected utility. Most researchers have estimated a discrete choice model for the probability of divorce in which the covariates are proxies for the expected gains to marriage and divorce. In this study, we extend the orthodox divorce model by assuming that (1) agents are risk averse, (2) remaining married is a risky option, and (3) divorce is an even riskier option. We demonstrate that the familiar risk premium argument can be applied to this scenario: conditional on the expected gains to marriage and divorce, the probability of divorce increases in risk tolerance because highly risk averse individuals require a relatively large premium in order to accept the greater risk associated with divorce.

We assess the empirical relationship between risk tolerance and divorce decisions by using NLSY79 data to estimate gender-specific probit models of divorce. We control for a rich array of proxies for the gains to marriage and divorce, as well as a measure of the Arrow-Pratt coefficient of relative risk tolerance derived from responses to questions on the willingness to make hypothetical, large-stakes income gambles. We find that 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 is associated with a 4.3% decrease in the predicted probability of divorce.

Our findings could reflect effects of unobserved components of the gains to marriage and divorce that are correlated with individual levels of risk tolerance. However, they are consistent with the risk premium interpretation and with the notion that divorce entails a greater income gamble for women than for men. If this risk premium 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 (by reducing the riskiness of the gamble) of enticing relatively risk averse women to end their marriages.

APPENDIX: COMPUTING RELATIVE RISK TOLERANCE

We use the model given by Eq. (5) in the Data section, along with data on age and categorical responses to the income gamble questions, to compute gender-specific maximum likelihood estimates for the parameters β, $\overline{\alpha}$, σ_α, 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 responses to the income gamble question in one, two, or all three survey years (1993, 2002, and 2004). All NLSY79 respondents who answer the questions contribute to the likelihood function regardless of whether they appear in our "first marriage" samples. 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

$$P(c_{it} = j \mid Age_{it}) = P\left(\log \underline{\rho}_j < \log \rho_{it} < \log \overline{\rho}_j\right)$$
$$= \Phi\left(\frac{\log \overline{\rho}_j - \overline{\alpha} - \beta AGE_{it}}{\sqrt{\sigma_\alpha^2 + \sigma_u^2}}\right) - \Phi\left(\frac{\log \underline{\rho}_j - \overline{\alpha} - \beta AGE_{it}}{\sqrt{\sigma_\alpha^2 + \sigma_u^2}}\right),$$

where $\underline{\rho}_j$ and $\bar{\rho}_j$ are the lower and upper bounds associated with risk category j (see footnote 8) 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

$$P(c_{it} = j, c_{is} = k \mid \text{Age}_{it}, \text{Age}_{is}) = P\left(\log \underline{\rho}_j < \log \rho_{it} < \log \bar{\rho}_j, \log \underline{\rho}_k < \log \rho_{is} < \log \bar{\rho}_k\right)$$

$$= \int \left[\left\{\Phi\left(\frac{\log \bar{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u}\right)\right.\right.$$

$$\left.\left.\times \left\{\Phi\left(\frac{\log \bar{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u}\right)\right\}\right\}\right] dF(\alpha_i).$$

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:

$$P(c_{it} = j, c_{is} = k, c_{ir} = l \mid AGE_{it}, AGE_{is}, AGE_{ir})$$

$$= P\left(\log \underline{\rho}_j < \log \rho_{it} < \log \bar{\rho}_j, \log \underline{\rho}_k < \log \rho_{is} < \log \bar{\rho}_k, \log \underline{\rho}_l < \log \underline{\rho}_{ir} < \log \bar{\rho}_l\right)$$

$$= \int \left[\left\{\Phi\left(\frac{\log \bar{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_j - \alpha_i - \beta AGE_{it}}{\sigma_u}\right)\right\}\right.$$

$$\times \left\{\Phi\left(\frac{\log \bar{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_k - \alpha_i - \beta AGE_{is}}{\sigma_u}\right)\right\}$$

$$\left.\times \left\{\Phi\left(\frac{\log \bar{\rho}_l - \alpha_i - \beta AGE_{ir}}{\sigma_u}\right) - \Phi\left(\frac{\log \underline{\rho}_l - \alpha_i - \beta AGE_{ir}}{\sigma_u}\right)\right\}\right] dF(\alpha_i).$$

The maximum likelihood estimates that we compute are given in Table 1. We use these estimates to compute the expected 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

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

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

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

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

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

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