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No Trend in the Intergenerational Transmission of Divorce

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ABSTRACT

This paper resolves apparent discrepancies concerning trends in the intergenerational transmission of divorce. Wolfinger (1999) reports that divorce transmission has weakened substantially, due perhaps to declining stigma or selectivity among those who divorce. While plausible, his findings contrast sharply with others (McLanahan and Bumpass 1988; Teachman 2002), who report no trend in divorce transmission. Our analyses of data from the National Survey of Families and Households, using life table methods and a Cox model stratified by marriage cohort, provide no support for a decline in divorce transmission. We note that the General Social Survey data used by Wolfinger lack information on marital duration, permitting analysis only for whether respondents have divorced by interview. Apparent declines in divorce transmission thus could be an artifact of longer exposures to risk, and hence higher probabilities of divorce by interview, for earlier marriage cohorts relative to later cohorts. Our analyses of the GSS and NSFH reveal that this artifact is indeed present. We conclude that there has been no trend in divorce transmission in recent decades in the United States.

INTRODUCTION

The intergenerational transmission of divorce has been well documented. Children of divorce are more likely to divorce than those who grew up in two-parent families (e.g., Amato 1996; Bumpass, Castro Martin, and Sweet 1991; Bumpass and Sweet 1972; Glenn and Kramer 1987; Keith and Finlay 1988; Kiernan and Cherlin 1999). There is, however, less consensus concerning *trends* in the intergenerational transmission of divorce. In a paper published in *Demography,* Wolfinger (1999) presents analyses that he interprets as suggesting that differences in divorce rates for children in divorced and intact families declined by almost 50 percent between 1973 and 1996. His findings stand in marked contrast to other studies that find little or no trend in the intergenerational transmission of divorce in the United States (McLanahan and Bumpass 1988; Teachman 2002).

In this paper, we resolve this apparent contradiction concerning trends in the intergenerational transmission of divorce by noting a fundamental flaw that afflicts Wolfinger's (1999) analyses. Wolfinger pools data from the 1973-96 General Social Surveys (GSS) to analyze trends in the probability that a respondent *ever* divorced by the time of GSS interview. Using this design, Wolfinger reports that the difference in the probability of ever divorcing for those raised in intact and nonintact families is higher for those who interviewed in earlier versus later periods. This decline in the probability of ever divorcing by survey year is then interpreted as evidence of a declining trend in the intergenerational transmission of divorce.

The key difficulty lies in use of Wolfinger's (1999) dependent variable—the *probability* that a respondent is *ever* divorced by GSS interview—to assess historical trends in demographic behaviors, typically and optimally measured in terms of *rates*. A classic demographic observation concerning exposure and risk is that differences in exposure can generate differences in the probability of ever divorcing even if divorce *rates* are identical. This problem is easiest to understand in terms of trends in divorce by marriage cohort: If divorce rates were identical across marriage cohorts, respondents in earlier marriage cohorts would nevertheless have higher probabilities of ever divorcing than respondents in later marriage cohorts simply by virtue of longer exposures to the risk of divorce. We present theoretical findings in Appendix 1 showing this observation holds not only for aggregate trends but also for trends in *differential* behaviors as in the intergenerational transmission of divorce. As a result, Wolfinger's empirical evidence concerning a historical decline in the differentials in the *probability* of ever divorcing for children of intact and divorced families could easily be an artifact reflecting no historical trend in divorce *rates*.

We present empirical analyses using the 1987-88 National Survey of Families and Households (NSFH) and 1973-1996 General Social Surveys, with the latter analyses used both to replicate Wolfinger's findings and to illustrate the artifact arising from the flawed design employed in his analyses. Our results suggest no evidence of a trend in the intergenerational transmission of divorce, a conclusion consistent with findings based on the first five waves of the National Survey of Family Growth (McLanahan and Bumpass 1988; Teachman 2002), but at odds with Wolfinger's (1999) analyses of the 1973-1996 General Social Surveys. Both our NSFH and GSS analyses reveal biases stemming from how one deals with variations in the exposure to the risk of divorce. We conclude that the association between a parental divorce and offspring's own marital success is still as powerful as in decades past, despite the increasing prevalence and acceptance of divorce.

THEORY

Background

Social scientists and the general public alike have long been interested in the consequences of family disruption on child well-being (Seltzer 1994), and in their implications for social inequality (Ellwood and Jencks 2004). Thus, it is perhaps not surprising that Wolfinger's (1999) conclusion that "adults raised by divorced mothers are far less likely to get divorced themselves than they were 20 years ago" (*USA Today*, August 11, 1999) was hailed as "Good News for the Children of Divorce" (*New York Times* August 17, 1999).[1]

The interest in studying the *trends* of consequences of divorce stems from what it might tell us about changes in the composition of those who marry and evolving norms and attitudes towards divorce. For example, it is plausible that the social stigma attached to divorce has declined over time as more individuals and families experience divorce and as public opinion has become more accepting of divorce (Thornton 1989). If true, less stigmatization against divorced parents and their children might lead to a decline in the intergenerational transmission of divorce (Diekmann and Engelhardt 1999; Glenn and Kramer 1987; Wolfinger 1999) with increasing rates of divorce and cohabitation acting to select out the least satisfactory marriages (Elwert 2005), thus raising the average marital quality for those who remained married (Glenn 1996). However, other studies suggest the opposite, with marriages contracted recently having lower quality than marriages

[1] Media reports confuse an apparent rising trend in *aggregate* divorce rates with the trend in *differential* divorce rates between adult children of divorce and those who raised by both biological parents. Trends in aggregate divorce rates in the past few decades suggests that both adults raised in both intact and divorce families face a higher risk of divorce than their counterparts several decades ago.

contracted in the past (Rogers and Amato 1997). Moreover, a decline in the intergenerational transmission of divorce may carry important policy implications, with divorce potentially of less concern if children of divorce were in fact less stigmatized, and themselves less likely to divorce, than in the past.

Prior Research on Trends in Divorce Transmission

Only a handful of studies have investigated whether the association between parental divorce and offspring's own divorce has shifted over time. McLanahan and Bumpass (1988), analyzing data from the 1982 National Survey of Family Growth, found that the intergenerational transmission of divorce for those married after 1970 was similar for all marriage cohorts in their sample, suggesting little or no trend by marriage cohort in divorce transmission. Teachman (2002) pooled data from five waves of the National Survey of Family Growth to investigate differentials in divorce rates for a number of observed characteristics of respondents. He found that the effect of parental divorce remained fairly constant for marriages contracted between 1950 and 1984, a finding again suggesting little or no trend by marriage cohort in the intergenerational transmission of divorce. In analyzing the German Youth Institute's 1988 Family Survey, Diekmann and Engelhardt (1999) report similar effects of family structure on offspring's divorce rates for those born before and after World War II, with parental marital dissolutions prior to 1946 primarily reflecting war casualties and those after 1946 more likely due to divorce. These results suggest little or no trend by birth cohort in the intergenerational transmission of divorce. Engelhardt and colleagues (2002) report a finding based on the German Life History Study that they interpret as consistent with a decline in the intergenerational transmission of divorce. However, because none of their coefficients for the trend are statistically significant, their results can also be interpreted as showing no discernible trend in the transmission of divorce risks. Thus, findings from two U.S. and one German studies examining six nationally representative surveys show little or no trend in the intergenerational transmission of divorce, with one German study being inconclusive.

An important characteristic of these studies is that they all employed hazard regression methods, with McLanahan and Bumpass (1988), Teachman (2002), and Engelhardt et al. (2002) reporting results based on Cox models and Diekmann and Engelhardt (1999) from a sickle model (Diekmann and Mitter 1984). In particular, all four studies estimated the duration-specific rate of divorce and thus controlled for differences in exposure to the risk of divorce as experienced by different birth or marriage cohorts. By contrast, Wolfinger's (1999) GSS analyses examined trends in the probability that a respondent ever divorced by the time of survey interview. The multiple cross-sectional GSS sample yields a design in which marriages contracted in recent years[2] are exposed to a shorter period of time at risk of divorce, compared to marriages contracted in earlier years. As noted above, the difficulty is that older marriage cohorts will have longer exposures to the risk of divorce; hence, even if divorce rates were identical across marriage cohorts, respondents from earlier marriage cohorts would nevertheless have a higher probability of ever divorcing than respondents in later marriage cohorts because of their longer exposure to the risk of divorce. The same difficulty arises in assessing historical trends in the *differential* divorce rates between

[2] Those who married in recent years were also included disproportionately in the recent GSS surveys. This may explain why the same trend holds for both marriage cohort and survey year.

individuals raised in intact families and in divorced families. According to results reported in Appendix 1, the same reason may lead to an artifactual *decline* in the *differential probability* of ever divorcing between the two groups even if the *differential risk* of divorce between the two groups has not changed.

DATA

We use data from the 1973-1996 General Social Survey used by Wolfinger and an additional data source, the 1987-88 National Survey of Families and Households. We omit a detailed description of the GSS data, which can be found in Wolfinger (1999). The National Survey of Families and Households consists of a national probability sample of persons ages 19 and over who resided in the United States in 1987 and 1988, with a main sample of 9,643 cases and an oversample of 3,374 cases of minorities, single-parent families, families with stepchildren, cohabiting couples, and recently married persons (Sweet, Bumpass, and Call 1988). The survey provides detailed information on marital histories. We contrast our NSFH results with our replication of Wolfinger's GSS results supplemented with additional analyses of the 1973-1996 GSS data. To ensure comparability, we follow Wolfinger's sample selection criteria save for minor details[3] and employ a comparable set of controls in both our NSFH and GSS

[3] The size of our GSS sample differs from Wolfinger's by a few thousand cases because he excludes respondents raised in all other family structures than "intact two-parent families and mother-only families resulting from divorce or separation, or mother/stepfather families resulting from divorce or separation" (Wolfinger 1999:416). We choose not to impose this sample restriction to ensure that any null findings are not an artifact of lack of statistical power. Nevertheless, our results are not sensitive to this choice (see the

analyses. Table 1 gives the descriptive statistics of our NSFH and GSS analytic samples. While the two samples are reasonably similar on most characteristics, the GSS sample contains a slightly wider range of marriage cohorts, containing first marriages contracted between 1901 and 1996, than the NSFH sample, with first marriages contracted between 1915 and 1988.

[Table 1 about here].

RESULTS

Trends in the Probability of Divorce Transmission by Survey Year

We begin by replicating Wolfinger's (1999) GSS analysis using (1) a logistic regression model for the probability of ever divorcing by the GSS interview, (2) survey year to proxy trends in divorce, (3) a linear spline of ($AGE - AGEWED$) as a proxy adjustment for "right censoring" (see Wolfinger, p. 417), and (4) the same set of control covariates used by Wolfinger. Comparing Wolfinger's coefficients in column 1 of Table 2 with our estimates in column 2 shows that the minor differences in sample size yield estimates that are slightly smaller in absolute values than Wolfinger's but otherwise very similar in sign, magnitude and statistical significance. In column 3 of Table 2, we drop the linear spline of ($AGE - AGEWED$) that Wolfinger uses as a rough proxy for right censoring; comparing results of this model to those in the first two columns shows that dropping the linear spline has almost no impact on the resulting coefficient estimates.

[Table 2 about here].

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results in Appendix 3). More importantly, we successfully replicate Wolfinger's findings without following this sample restriction (see Table 2).

Trends in the Probability of Divorce Transmission by Marriage Cohort

In Table 3, we show that Wolfinger's results are largely unchanged when modeling trends in the probability of divorce transmission using marriage cohort instead of survey year. The results from both the GSS and NSFH exhibit a decline in the probability of divorce transmission by marriage cohort. Our GSS results (columns 1 and 2) follow the same pattern as the GSS results in Table 2, with the linear spline of ($AGE - AGEWED$) yielding little difference in coefficient estimates. By contrast, our NSFH results exhibit a statistically significant declining trend in the probability of divorce transmission only when we do not employ the linear spline specification (column 3). Including the linear spline (column 4) reduces the magnitude and statistical significance of the coefficient for parental divorce, changes the sign and statistical significance of the coefficients for marriage cohort and the interaction of parental divorce and marriage cohort, and substantially drives up the standard error for marriage cohort. These results suggest that the NSFH sample is more sensitive to Wolfinger's linear spline ($AGE - AGEWED$) specification than is the GSS sample.

[Table 3 about here]

Sensitivity of Trends in the Probability of Divorce Transmission to Different Exposures to Risk

In Table 4, we assess the sensitivity of estimated trends in the probability of divorce transmission when individuals face different lengths of exposure to risk. As noted earlier, apparent declines in the probability of divorce transmission may reflect no trend in the intergenerational transmission of divorce risks because the logistic regressions in Wolfinger (1999) do not account for different exposures to the risk of divorce. In Table

4, we show the sensitivity of Wolfinger's results when GSS respondents face different exposures to the risk of divorce by gradually restricting the analysis to subsamples of increasingly smaller exposure years (column 2). The first line of Table 4 (Model 1) reproduces Wolfinger's results using the entire GSS period of observation in which GSS respondents face a maximum of 32 years of exposure. As before, the coefficient for the interaction of parental divorce and marriage cohort is -.024, that is, negative, and statistically significant, thus consistent with Wolfinger's assertion of a significant decline in the probability of divorce transmission. However, as we narrow the years of exposure for GSS respondents in subsequent models, the magnitude of the interaction coefficient shrinks towards zero and loses statistical significance. The results thus show that conclusions from a logistic regression framework for the probability of divorce transmission appear to vary with duration of exposure, with the sample comprising the longest average durations of exposure yielding coefficients consistent with a declining probability of divorce transmission and the sample with the most homogeneous exposures of exposure yielding coefficients that suggest no trend in the probability of divorce transmission.

[Table 4 about here]

Discrete-Time Logistic Hazard Regression Using the NSFH

The analyses in Table 4 using the GSS are relatively crude because assessing the sensitivity of results to different durations of exposure required relying on different GSS subsamples. In Table 5, we use data on the marriage duration in the NSFH to estimate a series of discrete-time logistic regressions for the duration-specific risk of divorce. Unlike the strategy in Table 4, we retain all of the NSFH sample but create different numbers of person-period records across models. In these models, the number of periods per person (and thus total number of person-period records in the data) regulate the degree to which exposure to the risk of divorce is adjusted. In particular, when the data are constructed with only one period per person, this yields estimates identical to the logistic regression model used by Wolfinger (1999). However, when we increase the number of periods per person (e.g., making essentially a "person-year" dataset in Model 10), this method reflects the typical application of a discrete-time hazard modeling strategy (Allison 1982). For formal details, see Appendix 2.

[Table 5 about here]

When we specify a model in which there is only one duration-period for all persons (Model 1), this specification yields coefficients identical to those in column 3 in Table 3. Thus, the GSS and NSFH analyses in Table 3, in which we replicate Wolfinger's (1999) finding of a declining trend in the probability of intergenerational divorce transmission, can be seen within a discrete-time hazard regression framework as resting on the assumption that variation by duration in the risk of divorce can be modeled with one parameter—an extremely strong and highly implausible assumption. As we relax this assumption by increasing the number of duration-year-intervals in subsequent models, the coefficient for the interaction of parental divorce and marriage cohort decreases in magnitude and becomes statistically insignificant for all models with 5 or more duration intervals. Note in particular that the lack of significance for the coefficient for the interaction of parental divorce and marriage cohort is not due to inflated standard errors; instead, the standard error for this coefficient is relatively stable across models and, indeed, reduces slightly with the number of duration-periods specified. These results, thus, suggest that the coefficient for the trend in the intergenerational transmission of divorce risks is highly sensitive to model assumptions about how the risk of divorce varies by duration, with models making weaker assumptions yielding coefficients that are not statistically significant and small in magnitude. We interpret these results as strong evidence in support of our assertion that Wolfinger's (1999) finding of a decline in the intergenerational transmission of divorce is a methodological artifact stemming from different exposures to the risk of divorce.

Nonparametric Life-Table Estimates

In Figure 1, using the NSFH we present smoothed nonparametric estimates of the logarithm of the duration-specific risk of divorce that make no statistical assumptions about the functional form of the underlying divorce rates (Wu 1989). We report these exploratory results in order to motivate specific elements of our final confirmatory analyses using a stratified Cox model reported in the next section.

[Figure 1 about here]

The smoothed nonparametric estimates in Figure 1 suggest that adult children from divorced families (the solid line in each panel) have consistently higher divorce rates than those growing up in intact families (the dashed line in each panel). The top panel consists of marriages contracted between 1935 and 1954, the middle panel between 1955 and 1974, and the bottom panel between 1975 and 1988. Within each panel, the pairs of lines are roughly parallel, indicating that the usual proportionality assumption appears to hold within a marriage cohort. However, note that the log risks across panels differ, with the two earliest marriage cohorts following a roughly monotonic pattern of duration dependence and the most recent marriage cohort following a unimodal pattern of duration dependence. These results suggest that specifying a single baseline hazard across marriage cohorts may yield a misspecified model of divorce risks. Finally, note that the differentials in divorce rates for adult children from divorce and non-divorced families are approximately the same across marriage cohorts (about 0.6 on the logged monthly rate scale). Thus, these nonparametric estimates suggest that adult children of divorce who contracted their marriages between 1935 and 1954 are about as likely to separate or divorce than those from intact families as those who contracted their marriages between 1955 and 1974 and between 1975 and 1988. Overall, these exploratory results, which control only for duration, parental divorce, and marriage cohort, are nevertheless suggestive of no trend in the intergenerational transmission of divorce.

Analyses Using a Stratified Cox Model

As noted above, the nonparametric results in Figure 1 are akin to bivariate correlations in that they lack the usual statistical controls for other covariates. In Table 6, we report results from multivariate analyses that control for other covariates. We use a Cox model specification stratified by the three 20-year marriage cohorts in Figure 1. This model specifies a common baseline hazard for each 20-year marriage cohort but allows a different and unspecified functional form for the baseline hazard across the three marriage cohorts. The effects of covariates are assumed to be proportional, and are constrained to be equal across strata (Therneau and Grambsch 2000). We report estimates from two models, a first in which we estimate the main effects of parental divorce and marriage cohort, and a second model in which we add the interaction of parental divorce and marriage cohort.

[Table 6 about here]

In Model 1, both coefficients for parental divorce and marriage cohort are in the expected directions and statistically significant, with the risk of divorce increasing with successive cohorts (see also Preston and McDonald 1979). In Model 2, the coefficient for the interaction of parental divorce and marriage cohort is small (-0.001) and not statistically significant (standard error five times the size of the estimated coefficient), showing no trend in the intergenerational transmission of divorce rates. A comparison of likelihoods for the two nested models show a negligible increment in fit. There appears to be substantial collinearity between the main effect of parental divorce and its interaction effect with marriage cohort, with the interaction term increasing the standard error for the coefficient for parental divorce by a factor of five; however, the standard error for the interaction term (.005) is smaller than the corresponding standard errors in both our replication of Wolfinger's GSS analysis (.006, see Table 3) and in Wolfinger's own logistic regression (.008, Wolfinger 1999, Table 1, Model 3, p. 418).

Results in Table 6 are also consistent with the patterns observed in the nonparametric estimates in Figure 1. Note also that the magnitude of the interaction terms relative to the main effect of parental divorce are larger in the logistic regressions in Table 3 than in the hazard regression results of Table 6 by a factor greater than 4. Put another way, the hazard regression point estimates suggest that if there were any decline in the trend for divorce transmission, it would require more than 100 years to reach the comparable trend in less than a quarter century (1973-1996) as implied by Wolfinger's estimates and those in the logistic regression models reported in Table 2 and Table 3.

CONCLUSION

Our results, like numerous prior studies, provide continuing evidence consistent with a view that children who experience the divorce of their parents are themselves more likely to divorce. But in contrast to Wolfinger (1999), who asserts that the intergenerational transmission of divorce has declined by nearly 50%, we find no trend in the intergenerational transmission of divorce, a result mirroring findings in Diekmann and Engelhardt (1999), McLanahan and Bumpass (1988) and Teachman (2002). We provide evidence showing that the likely source of the discrepancy between Wolfinger's findings and those of others (including ours) is due to differential exposure to the risk of divorce by marriage cohort. We conclude that Wolfinger's findings are a methodological artifact stemming from the fact that those married earlier will have longer durations of exposure to divorce, while those married later will have shorter durations of exposure. As a result, Wolfinger's logistic regression estimates of trends in the *probability* of the intergenerational transmission of divorce lack appropriate controls for exposure, whereas hazard regression estimates of trends in the intergenerational transmission of *duration-specific divorce rates* will automatically adjust for differences in exposure to risk. Our analyses using both the NSFH data and Wolfinger's GSS data illustrate that the potential biases caused by omitting appropriate controls for duration of exposure to risk are not just a theoretical possibility but in fact generate Wolfinger's spurious finding of a declining trend in the intergenerational transmission of divorce.

Our conclusion that there is no trend in the intergenerational transmission of divorce holds for marriage cohorts in the period between 1915 and 1988. It is, of course, possible that trends may differ for cohorts of marriages in the 1990s and later, but our analyses also show, we believe, that for the cohorts of marriages analyzed by Wolfinger (1999), there is no trend in the intergenerational transmission of divorce.[4]

Whether there is no trend in the intergenerational transmission of divorce or a declining trend is not merely a methodological question, but one that speaks more generally to important and ongoing debates in the social sciences more generally. For example, Wolfinger argues that as divorce has become more widespread, children of divorce will be less highly selected. We agree that this is a highly plausible (and testable) hypothesis, but we also conclude that the empirical evidence is not consistent with it. More generally, the intergenerational transmission of family behaviors has been argued to be a central component in reproducing and maintaining inequality (Biblarz and Raftery 1999; McLanahan 1985). Other highly influential social scientists have argued that growing up in a nonintact family structure is not merely associated with social and economic disadvantage, but plausibly considered as a causal relationship via a variety of mechanisms (see, e.g., Cherlin 1999; McLanahan and Sandefur 1994). If true, the question of whether there has or has not been a decline in the intergenerational transmission of divorce would presumably carry important implications. Others have presented evidence suggesting that the intergenerational transmission of poverty and the intergenerational transmission of family structure follow independent pathways (Musick and Mare 2006); nevertheless, even for these authors, questions of intergenerational

[4] More precisely, our NSFH analyses provide us with data on cohorts of marriages between 1915 and 1988, while Wolfinger analyzes whether respondents have ever been divorced by GSS interview for the 1973 to 1996 GSS surveys. As a consequence, the marriage cohorts in the GSS data analyzed by Wolfinger overlap substantially with those in the NSFH; in particular, they contain at most six years of additional marriage cohorts in the post-1988 period than do the 1988 NSFH data.

transmission are of central substantive importance. As a consequence, we view the questions we raise about Wolfinger's assertion—that there has been a marked decline in the intergenerational transmission of divorce—to be of sufficient importance to bear careful empirical scrutiny.

APPENDIX 1

Formal results on differential probabilities and rates

In this appendix, we provide some formal results specific to inferences about trends in the intergenerational transmission of divorce. We begin with a review of Wolfinger's model for trends in the intergenerational transmission of divorce, and then show how differential exposure to risk can yield trends in the probability of ever divorcing even if divorce rates are identical for successive marriage cohorts.

Because the GSS did not collect data on the date of marital dissolution, Wolfinger (1999) restricts his analyses to the probability of ever divorcing:

$$\log\left(\frac{p}{1-p}\right) = \beta_0 + \beta_1 \cdot PaDiv + \beta_2 \cdot Yr + \beta_3 \cdot (PaDiv \times Yr) + \sum \boldsymbol{\beta}\mathbf{x} \qquad (1)$$

where *p* denotes the probability of ever divorcing by GSS interview, *PaDiv* is a dummy variable equal to 1 if a respondent's parents divorced by age 16, *Yr* denotes calendar year of GSS survey, $PaDiv \times Yr$ denotes the interaction of parental divorce and calendar year, and $\mathbf{x}$ is a vector of control variables, including linear spline of ($AGE - AGEWED$) for the difference between age at interview and age at first marriage. Wolfinger asserts that a spline specification for ($AGE - AGEWED$) adjusts for right censoring and interprets a negative coefficient for β_3, the interaction term of $(PaDiv \times Yr)$, as evidence for a declining trend in *rates* of the intergenerational transmission of divorce. Note, however, that β_3 is more correctly interpreted as the difference in the *log-odds* of ever divorcing by survey for adults from intact families and divorced families as interacted with calendar year.

Consider a thought experiment in which two marriage cohorts are composed of identical individuals except that in one cohort, first marriages are contracted in calendar year *A* while in the other, first marriages are contracted in calendar year *B*, with $A < B$. Then following Wolfinger's GSS design, let both cohorts be interviewed in the same calendar year τ, with the duration between marriage and survey given by is $\tau^A = (\tau - A)$ and $\tau^B = (\tau - B)$, with $\tau^A > \tau^B$. The key insight is that the probability of ever divorcing by interview at calendar year τ will differ even though, by assumption, individuals in the two cohorts are identical. More precisely, let S^A and S^B denote the survivor probability (i.e., the probability of not divorcing) by survey at τ. Then:

$$S^A = \exp\left(- \int_0^{\tau^A} r(u \mid \mathbf{x}) du \right) \tag{2a}$$

$$S^B = \exp\left(- \int_0^{\tau^B} r(u \mid \mathbf{x}) du \right) \tag{2b}$$

where $r(u \mid \mathbf{x})$ is the duration-specific risk of divorce at duration *u* conditioning on a vector of covariates **x**. By assumption, individuals in the two cohorts are behaviorally identical; hence, $r(u \mid \mathbf{x})$ is identical in (2a) and (2b). To simplify the exposition of ideas, we assume that the duration-specific risk of divorce follows a non-defective distribution. Then note that because $\tau^A > \tau^B$, $S^A < S^B$ in (2a) and (2b) despite identical divorce risks by virtue of different exposures to risk.

The above is a classic demographic example, but it does not address whether a similar result holds for trends in the intergenerational transmission of divorce. That is, the example given above concerns two identical cohorts under different exposures, whereas questions involving trends in the intergenerational transmission of divorce involve differentials between two non-identical groups with different exposures.

The short answer is that the same result holds. To see this, consider a second thought experiment in which (i) children of divorce have higher divorce risks than their counterparts whose parents did not divorce, and (ii) that the difference in (i) follows a proportional hazard specification by a constant e^k that does not vary with marital duration. Under these assumptions, the duration-specific risk of divorce $r_1(u \mid \mathbf{x})$ for children of divorce will be higher than $r_0(u \mid \mathbf{x})$, the risk for their counterparts who grew up with both biological parents.

Let $P_1(t)$ denote the probability of divorcing by marital duration *t* for those whose parents were divorced by age 16, $S_1(t) = 1 - P_1(t)$ denote the probability of not yet having divorced by marital duration *t*; and let $P_0(t)$, and $S_0(t) = 1 - P_0(t)$ denote the equivalent quantities for those who did not experience a parental divorce. Then

$$S_0(t) = \exp\left(-\int_0^t r_0(u \mid \mathbf{x})du\right) \qquad \text{if } PaDiv = 0 \qquad (3a)$$

$$S_1(t) = \exp\left(-e^k \cdot \int_0^t r_0(u \mid \mathbf{x})du\right) \qquad \text{if } PaDiv = 1 \qquad (3b).$$

Under assumption (i), children of divorce have higher divorce risks, even after conditioning on the **x**'s, than their counterparts whose parents did not divorce. This implies, under proportionality assumption (ii) with $k > 0$ and $e^k > 1$, that $P_0(t) > P_1(t)$ and $S_1(t) > S_0(t)$. Similarly, by assuming a non-defective distribution, we write:

$$P_0(t) = \exp\left(-\int_t^\infty r_0(u \mid \mathbf{x})du\right) \qquad \text{if } PaDiv = 0 \qquad (4a)$$

$$P_1(t) = \exp\left(-e^k \cdot \int_t^{\infty} r_0(u \mid \mathbf{x})du\right) \qquad \text{if } PaDiv = 1 \qquad (4b).$$

The coefficient for *PaDiv* is thus the difference in log-odds between $P_1(t)$ and $P_0(t)$:

$$\log\left(\frac{P_1(t)}{1-P_1(t)}\right) - \log\left(\frac{P_0(t)}{1-P_0(t)}\right) = \log\left(\frac{P_1(t)}{S_1(t)}\right) - \log\left(\frac{P_0(t)}{S_0(t)}\right)$$

$$= \{\log(P_1(t)) - \log(S_1(t))\} - \{\log(P_0(t)) - \log(S_0(t))\}$$

$$= \{\log(P_1(t)) - \log(P_0(t))\} - \{\log(S_1(t)) - \log(S_0(t))\} \qquad (5).$$

From equations (3a) and (3b), we get:

$$\log(S_1(t)) - \log(S_0(t)) = e^k \cdot \left(-\int_0^t r_0(u \mid \mathbf{x})du\right) - \left(-\int_0^t r_0(u \mid \mathbf{x})du\right)$$

$$= (1-e^k) \cdot \left(\int_0^t r_0(u \mid \mathbf{x})du\right) \qquad (6).$$

Similarly, from Equations (4a) and (4b), we get:

$$\log(P_1(t)) - \log(P_0(t)) = e^k \cdot \left(-\int_t^{\infty} r_0(u \mid \mathbf{x})du\right) - \left(-\int_t^{\infty} r_0(u \mid \mathbf{x})du\right)$$

$$= (1-e^k) \cdot \left(\int_t^{\infty} r_0(u \mid \mathbf{x})du\right) \qquad (7).$$

Hence, the coefficient for *PaDiv* expressed in Equation (5) becomes $(7)-(6)$:

$$\log\left(\frac{P_1(t)}{1-P_1(t)}\right) - \log\left(\frac{P_0(t)}{1-P_0(t)}\right)$$

$$= (1-e^k) \cdot \left(\int_t^{\infty} r_0(u \mid \mathbf{x})du\right) - (1-e^k) \cdot \left(\int_0^t r_0(u \mid \mathbf{x})du\right)$$

$$= (1 - e^k) \cdot \left\{ \left(\int_t^\infty r_0(u \mid \mathbf{x}) du \right) - \left(\int_0^t r_0(u \mid \mathbf{x}) du \right) \right\} \quad (8).$$

For simplicity, let

$$H = \int_0^t r_0(u \mid \mathbf{x}) du \quad (9a),$$

and $$H' = \int_t^\infty r_0(u \mid \mathbf{x}) du \quad (9b).$$

Thus,

$$\log\left(\frac{P_1(t)}{1 - P_1(t)} \right) - \log\left(\frac{P_0(t)}{1 - P_0(t)} \right) = (1 - e^k) \cdot (H' - H) \quad (10).$$

Now consider the respective logit coefficients for two cohorts *A* and *B*:

$$\delta^A = \log\left(\frac{P_1^A(t)}{1 - P_1^A(t)} \right) - \log\left(\frac{P_0^A(t)}{1 - P_0^A(t)} \right)$$

$$= (1 - e^k) \cdot \left\{ \left(\int_{\tau^A}^\infty r_0(u \mid \mathbf{x}) du \right) - \left(\int_0^{\tau^A} r_0(u \mid \mathbf{x}) du \right) \right\}$$

$$= (1 - e^k) \cdot (H^{A}{}' - H^A) \quad (11a),$$

and $$\delta^B = \log\left(\frac{P_1^B(t)}{1 - P_1^B(t)} \right) - \log\left(\frac{P_0^B(t)}{1 - P_0^B(t)} \right)$$

$$= (1 - e^k) \cdot \left\{ \left(\int_{\tau^B}^\infty r_0(u \mid \mathbf{x}) du \right) - \left(\int_0^{\tau^B} r_0(u \mid \mathbf{x}) du \right) \right\}$$

$$= (1 - e^k) \cdot (H^{B}{}' - H^B) \quad (11b).$$

The difference between the logit coefficients of the two cohorts is

$$\delta^A - \delta^B = (1 - e^k) \cdot \left\{ (H^{A}{}' - H^A) - (H^{B}{}' - H^B) \right\}$$

$$= (1 - e^k) \cdot \left\{ (H^{A}{}' - H^{B}{}') - (H^A - H^B) \right\} \quad (12).$$

Because $\tau^A > \tau^B$ and from (9a) and (9b), we have $H^A > H^B$ and $H^{A}{}' < H^{B}{}'$. Thus, $H^{A}{}' - H^{B}{}' < 0$ and $-(H^A - H^B) < 0$. Also because $k > 0$, $e^k > 1$, and $1 - e^k < 0$, the entire right-hand side of Equation 12 is positive. Therefore, we have established that $\delta^A > \delta^B$.

The above derivations thus show that for two behaviorally identical cohorts, *A* and *B*, the logit coefficient for *PaDiv*, which Wolfinger interprets as evidence for the intergenerational transmission of divorce, will be larger for cohort *A* than *B* as a consequence of cohort A's longer exposure to risk. This extends the classic insight concerning exposure and probability of ever experiencing an event to *differentials* in the risk of an event.

APPENDIX 2

Discrete-time Logistic Regression Analyses

In the discrete-time hazard logistic regression analyses using the NSFH data, we specify a set of discrete-time logistic regression models as:

$$\log\left(\frac{q_t}{1-q_t}\right) = \sum_{t=1}^{T} \alpha_t D_t + \beta_1 PaDiv + \beta_2 Marcoh + \beta_3 (PaDiv \times Marcoh) + \boldsymbol{\beta}\mathbf{x}$$

where q_t is the predicted probability of divorce within the marital duration denoted by period *t* conditional on a marriage survived up to period $t-1$, and D_t is a set of dummy variables indicating period *t*. This model is known to approximate a continuous-time proportional hazard model in which the discrete-time baseline hazard is specified via $\sum_t \alpha_t D_t$ and the effects of covariates are multiplicative for each *t* (Allison 1982). When the length of each period *t* shrinks and the number of periods *T* increases, the true continuous-time baseline hazard is better approximated by $\sum_t \alpha_t D_t$ and the logistic coefficient estimates of β's in the discrete-time hazard specification and their counterparts in a continuous-time hazard specification will converge to one another. At the other extreme, when all the person-period records are collapsed into one and thus there is only one period (i.e., $T = 1$) per person, the baseline hazards of $\sum_t \alpha_t D_t$ then reduce to a single intercept β_0.

APPENDIX 3

Results for Replications Using the Same Samples Restrictions as in Wolfinger (1999)

Table A-1: Descriptive Statistics of NSFH and GSS samples

	NSFH1				1973-96 GSS			
Variable	Mean	s.d..	Min	Max	Mean	s.d.	Min	Max
Age at 1st marriage	22.43	4.74	9	72	22.01	4.42	12	73
Black	.14		0	1	.09		0	1
Catholic	.25		0	1	.25		0	1
Marriage cohort (yr.)	63.72	16.77	15	88	59.41	17.47	1	96
AGE-AGEWED	23.36	16.76	.1	72	25.37	16.35	0	74
Divorce	.36		0	1	.31		0	1
Edu (< hs)	.23		0	1	.25		0	1
Edu (some college)	.20		0	1	.04		0	1
Edu (bachelor)	.12		0	1	.13		0	1
Edu (post graduate)	.06		0	1	.06		0	1
Agewed missing	.01		0	1	.13		0	1
Male	.39		0	1	.42		0	1
Pa Edu missing	.12		0	1	.04		0	1
Prestg missing	.38		0	1	.62		0	1
Onlykid	.05		0	1	.05		0	1
Parental divorce	.09		0	1	.09		0	1
Pa Edu (< hs)	.41		0	1	.44		0	1
Pa Edu (some college)	.08		0	1	.02		0	1
Pa Edu (bachelor)	.05		0	1	.07		0	1
Pa Edu (post graduate)	.02		0	1	.04		0	1
Occu. Prestige	37.60	14.86	14	90	43.90	8.24	17	86
Rural	.29		0	1	.63		0	1
Duration (in month)	200.37	184.61	1	828				
Survey year					84.78	7.25	73	96
N (unweighted)	9,214				21,858			

TABLE A-2: Trends in the Probability of Divorce Transmission by Survey Year. GSS Sample.

	Wolfinger Model 3, Table 1	GSS replication with spline	GSS replication without spline
Parental divorce	2.56 ** (.68)	2.69 ** (.66)	2.68 ** (.63)
Survey year	.05 ** (.004)	.04 ** (.004)	.04 ** (.004)
Parental divorce × survey year	-.02 ** (.008)	-.02 ** (.008)	-.03 ** (.007)
N	21,963	21,858	21,858

* p < .05
** p < .01

TABLE A-3: Trends in the Probability of Divorce Transmission by Marriage Cohort. GSS and NSFH Samples.

	GSS	GSS+spline	NSFH	NSFH+spline
Parental divorce	2.12**	1.60**	1.41**	.43
	(.23)	(.24)	(.44)	(.48)
Marriage cohort	.004**	.038**	-.003	.009
	(.001)	(.004)	(.002)	(.076)
Parental. divorce ×	-.025**	-.016**	-.015*	.001
Marriage cohort	(.003)	(.004)	(.006)	(.007)
N	21,858	21,858	9,214	9,214

* p < .05
** p < .01

TABLE A-4: Trends in the Probability of Divorce Transmission by Marriage Cohort under Different Exposures to Risk.

Model	Exposure Time(yr)	Parental Divorce		Marriage Cohort		**Parental Divorce × Marriage Cohort**		N
1	0-32	2.51	(.38)**	-.024	(.002)**	**-.029**	**(.005)****	14,703
2	2-32	2.36	(.38)**	-.021	(.002)**	**-.027**	**(.005)****	14,361
3	4-32	2.12	(.40)**	-.011	(.002)**	**-.024**	**(.006)****	13,424
4	6-32	2.32	(.42)**	-.002	(.002)	**-.027**	**(.006)****	12,428
5	8-32	2.12	(.44)**	.005	(.003)*	**-.024**	**(.006)****	11,400
6	10-32	1.95	(.48)**	.011	(.003)**	**-.021**	**(.007)***	10,392
7	12-32	2.33	(.52)**	.017	(.003)**	**-.028**	**(.008)****	9,312
8	14-32	2.15	(.58)**	.022	(.004)**	**-.025**	**(.009)****	8,311
9	16-32	1.98	(.64)*	.030	(.004)**	**-.022**	**(.010)***	7,288
10	18-32	2.04	(.71)	.039	(.005)**	**-.023**	**(.011)***	6,222
11	20-32	1.36	(.80)	.044	(.006)**	-.010	(.013)	5,271
12	22-32	1.33	(.89)	.047	(.007)**	-.009	(.015)	4,365
13	24-32	1.65	(1.00)	.055	(.008)**	-.017	(.017)	3,540
14	26-32	1.04	(1.15)	.055	(.010)**	-.006	(.020)	2,700
15	28-32	1.81	(1.37)	.064	(.012)**	-.018	(.024)	1,880
16	30-32	1.69	(1.78)	.093	(.017)**	-.019	(.032)	1,114

* p < .05
** p < .01

Note: Standard errors in parentheses.
All models include the same set of control variables as in Wolfinger (1999).

TABLE A-5: Trends in the Risk of Divorce Transmission by Marriage Cohort: Discrete-Time Duration-Specific Risk of Divorce under Different Numbers of Discrete Duration Periods.

Model	Max. Number of periods/person	Parental divorce		Marriage cohort		Parental divorce × Marriage cohort		Number of person-period records
1	1	1.41	(.44)**	-.003	(.002)	**-.015**	**(.006)***	9,214
2	2	1.38	(.41)**	-.002	(.002)	**-.014**	**(.006)***	10,705
3	3	1.31	(.37)**	.002	(.002)	**-.013**	**(.005)***	12,144
4	5	1.13	(.34)**	.010	(.002)**	**-.011**	**(.005)***	16,382
5	7	.98	(.33)**	.015	(.002)**	-.009	(.005)	20,547
6	14	.82	(.33)*	.022	(.002)**	-.007	(.005)	35,645
7	18	.74	(.34)*	.023	(.002)**	-.005	(.005)	43,332
8	24	.72	(.34)*	.025	(.002)**	-.005	(.005)	56,093
9	35	.69	(.34)*	.027	(.002)**	-.005	(.005)	81,510
10	70	.66	(.35)	.028	(.002)**	-.004	(.005)	153,448

Note: Robust standard errors in parentheses.

* p < .05
** p < .01

TABLE A-6: Estimates of Trends in the Intergenerational Transmission of Duration-Specific Divorce Risks. Stratified Cox model and the NSFH Sample.

	Model 1	Model 2
Parental divorce	.36**	.45
	(.08)	(.40)
Marriage cohort	.030**	.031**
	(.005)	(.005)
Parental. divorce × Marriage cohort		-.001
		(.006)
Log likelihood	-18,543.304	-18,543.276

Note: Robust standard errors in parentheses.
All models include the same set of control variables as Wolfinger (1999); N = 9,214.

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Figure 1: Nonparametric Estimates of the Logarithm of Duration-Specific Divorce Risks by Parental Divorce and Marriage Cohort.

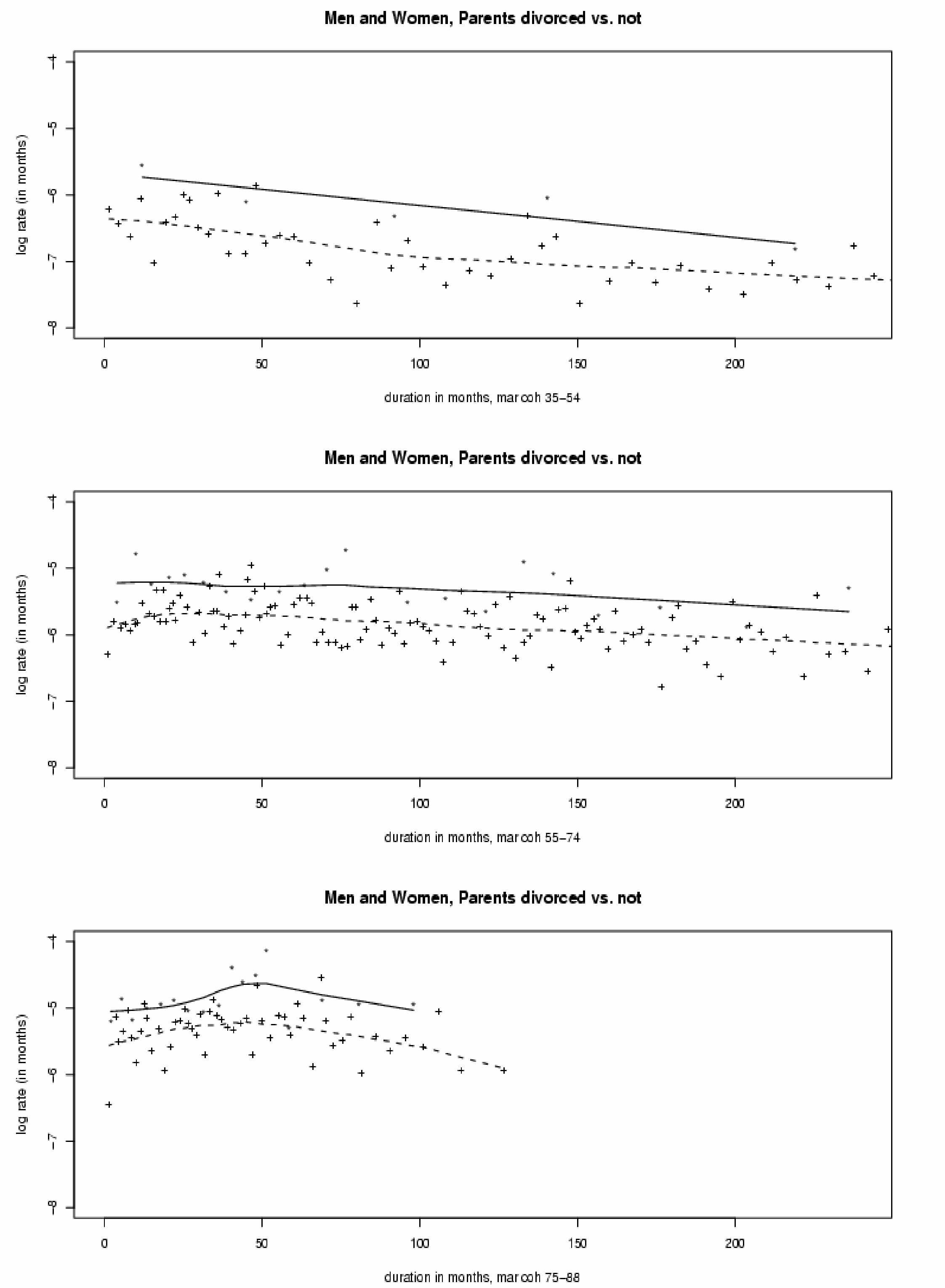

Table 1: Descriptive Statistics of NSFH and GSS samples

	NSFH1				1973-96 GSS			
Variable	Mean	Std. Dev.	Min	Max	Mean	Std. Dev.	Min	Max
Age at 1st marriage	22.61	4.77	7	72	21.97	4.42	12	73
Black	0.09		0	1	0.10		0	1
Catholic	0.26		0	1	0.25		0	1
Marriage cohort (yr.)	61.81	16.40	15	88	59.28	17.06	1	96
AGE-AGEWED	25.28	16.38	0.1	72	25.21	15.76	0	74
Divorce	0.27		0	1	0.28		0	1
Edu (< hs)	0.23		0	1	0.27		0	1
Edu (some college)	0.19		0	1	0.04		0	1
Edu (bachelor)	0.12		0	1	0.12		0	1
Edu (post graduate)	0.07		0	1	0.06		0	1
Agewed missing	0.01		0	1	0.13		0	1
Male	0.45		0	1	0.44		0	1
Pa Edu missing	0.13		0	1	0.05		0	1
Prestg missing	0.39		0	1	0.64		0	1
Onlykid	0.06		0	1	0.05		0	1
Parental divorce	0.09		0	1	0.09		0	1
Pa Edu (< hs)	0.42		0	1	0.46		0	1
Pa Edu (some college)	0.07		0	1	0.02		0	1
Pa Edu (bachelor)	0.05		0	1	0.07		0	1
Pa Edu (post graduate)	0.02		0	1	0.04		0	1
Occu. Prestige	38.00	14.92	14	90	43.52	8.02	17	86
Rural	0.27		0	1	0.37		0	1
Duration (in month)	234.07	190.23	1	828				
Survey year					84.48	7.31	73	96
N (unweighted)		10,216				25,923		

TABLE 2: Trends in the Probability of Divorce Transmission by Survey Year. GSS Sample.

	Wolfinger Model 3, Table 1	GSS replication with spline	GSS replication without spline
Parental divorce	2.56 ** (.68)	2.20 ** (.55)	2.12 ** (.53)
Survey year	.05 ** (.004)	.04 ** (.004)	.04 ** (.004)
Parental divorce × survey year	-.02 ** (.008)	-.02 ** (.006)	-.02 ** (.006)
N	21,963	25,923	25,923

* p < .05
** p < .01

TABLE 3: Trends in the Probability of Divorce Transmission by Marriage Cohort. GSS and NSFH Samples.

	GSS	GSS+spline	NSFH	NSFH+spline
Parental divorce	1.94**	1.41**	1.27**	.43
	(.19)	(.20)	(.38)	(.39)
Marriage cohort	.004**	.038**	-.002	.023
	(.001)	(.004)	(.002)	(.062)
Parental. divorce ×	-.022**	-.013**	-.012*	.001
Marriage cohort	(.003)	(.003)	(.005)	(.006)
N	25,923	25,923	10,216	10,216

* p < .05
** p < .01

TABLE 4: Trends in the Probability of Divorce Transmission by Marriage Cohort under Different Exposures to Risk.

Model	Exposure Time(yr)	Parental Divorce		Marriage Cohort		**Parental Divorce × Marriage Cohort**		N
1	0-32	2.09	(.33)**	-.024	(.002)**	**-.024**	**(.005)****	16,999
2	2-32	1.93	(.33)**	-.020	(.002)**	**-.022**	**(.005)****	16,615
3	4-32	1.64	(.35)**	-.011	(.002)**	**-.017**	**(.005)****	15,552
4	6-32	1.83	(.37)**	-.002	(.002)	**-.020**	**(.005)****	14,427
5	8-32	1.61	(.39)**	.005	(.002)*	**-.017**	**(.006)****	13,255
6	10-32	1.44	(.42)**	.011	(.003)**	**-.015**	**(.006)***	12,118
7	12-32	1.72	(.46)**	.017	(.003)**	**-.020**	**(.007)****	10,882
8	14-32	1.55	(.51)**	.022	(.003)**	**-.017**	**(.008)***	9,706
9	16-32	1.25	(.56)*	.029	(.004)**	-.012	(.009)	8,524
10	18-32	1.17	(.62)	.038	(.004)**	-.010	(.010)	7,295
11	20-32	.60	(.69)	.043	(.005)**	.001	(.011)	6,202
12	22-32	.64	(.77)	.047	(.006)**	.001	(.013)	5,163
13	24-32	1.08	(.86)	.055	(.007)**	-.010	(.015)	4,174
14	26-32	.37	(1.00)	.054	(.009)**	.003	(.017)	3,193
15	28-32	.83	(1.17)	.061	(.011)**	-.005	(.021)	2,232
16	30-32	1.21	(1.56)	.088	(.015)**	-.013	(.028)	1,312

* p < .05
** p < .01

Note: Standard errors in parentheses.
All models include the same set of control variables as in Wolfinger (1999).

TABLE 5: Trends in the Risk of Divorce Transmission by Marriage Cohort: Discrete-Time Duration-Specific Risk of Divorce under Different Numbers of Discrete Duration Periods.

Model	Max. Number of periods/person	Parental divorce		Marriage cohort		Parental divorce × Marriage cohort		Number of person-period records
1	1	1.27	(.38)**	-.002	(.002)	**-.012**	**(.005)***	10,216
2	2	1.22	(.35)**	-.001	(.002)	**-.012**	**(.005)***	11,912
3	3	1.18	(.33)**	.003	(.002)	**-.011**	**(.005)***	13,515
4	5	.99	(.30)**	.011	(.002)**	-.008	(.004)	18,255
5	7	.86	(.30)**	.016	(.002)**	-.006	(.004)	22,911
6	14	.72	(.30)*	.022	(.002)**	-.004	(.004)	39,784
7	18	.66	(.20)*	.024	(.002)**	-.003	(.004)	48,352
8	24	.66	(.30)*	.025	(.002)**	-.004	(.004)	62,621
9	35	.62	(.30)*	.027	(.002)**	-.003	(.004)	90,979
10	70	.57	(.31)	.029	(.002)**	-.002	(.004)	171,139

Note: Robust standard errors in parentheses.

* p < .05
** p < .01

TABLE 6: Estimates of Trends in the Intergenerational Transmission of Duration-Specific Divorce Risks. Stratified Cox model and the NSFH Sample.

	Model 1	Model 2
Parental divorce	.41**	.45
	(.07)	(.34)
Marriage cohort	.030**	.031**
	(.005)	(.005)
Parental. divorce × Marriage cohort		-.001
		(.005)
Log likelihood	-21,175.735	-21,175.726

Note: Robust standard errors in parentheses.
All models include the same set of control variables as Wolfinger (1999); N = 10,126.

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