An Empirical Test of Crisis, Social Selection, and Role Explanations of the Relationship Between Marital Disruption and Psychological Distress: A Pooled Time-Series Analysis of Four-Wave Panel Data

Although a higher level of psychological distress has been found in many studies of divorced compared with married individuals, explanations for this difference remain elusive. Three basic theoretical explanations have been proposed. Social role theory maintains that the role of being divorced is inherently more stressful than that of being married; crisis theory attributes the higher stress to role transitions and transient stressors of the disruption process, and social selection theory claims that the higher stress levels among the divorced result from the selection of people with poor mental health into divorce. Some empirical support is available for each of these approaches, but all three have not been tested simultaneously in a longitudinal study. This research empirically evaluates the efficacy of these theories in a pooled time-series analysis of a four-wave panel of married persons followed over 12 years. The pooled-time series random effects model was used to estimate the effects of social roles, crisis, and social selection. The results provide evidence that the higher stress levels of the divorced primarily reflect the effect of social role with selection and crisis effects making small contributions only.

A number of previous studies clearly show that the divorced report higher levels of psychological distress than do the married (Booth & Amato, 1991; Coombs, 1991; Gove, Hughes, & Briggs, 1983; Mastekaasa, 1994; Ross, 1995; Waite, 1995). Empirical studies have not been as definitive about the factors that account for this relationship. Each of the three explanations that have been proposed have some empirical support. A social selection explanation maintains that persons with high psychological distress and mental disorders are disproportionately selected into divorce and less likely to remarry, yielding higher distress scores among the currently divorced (Aseltine & Kessler, 1993). According to crisis theory, the disruption process and resultant role transitions temporarily elevate distress (Booth & Amato). Role theory attributes the greater psychological distress reported by the divorced to the more difficult life circumstances they experience (Ross).
An adequate simultaneous test of these competing explanations in a single study has been limited by small sample sizes, cross-sectional designs, limited duration of panel studies, lack of measures of distress pre- and postdisruption, and analysis design limitations. The purpose of this study is to overcome some of these limitations by testing these explanations with pooled-time series models in a large nationally representative four-wave panel sample extending over a 12-year period. Measures used include psychological distress both pre- and postmarital disruptions.

SOCIAL SELECTION

One explanation for the higher psychological distress of the divorced compared with the married is that married persons with preexisting mental health problems are often inadequate marital partners and thus are more likely to be selected into divorce than persons without these problems. This has been referred to as the social selection hypothesis (Avison, 1999). This theoretical perspective asserts that at least part of the relationship between distress and divorce is not causal and reflects the influence of stable individual personality and social characteristics on the odds of divorcing and on psychological distress. A number of research studies have examined the selection hypothesis (Aseltine & Kessler, 1993; Bloom, Niles, & Tatcher, 1985; Davies, Avison, & McApline, 1997; Kitson & Sussman, 1982; Mastekaasa, 1997; Robins & Regier, 1991), but the magnitude of the selection effect and the extent of its impact on the observed difference in distress levels between married and divorced persons are less certain. Studies involving small samples followed over time have not found significant differences in psychological disorders at the beginning of the time period between persons who do or do not eventually divorce (Doherty, 1983; Menaghan, 1985; Menaghan & Lieberman, 1986). These same studies did find a large difference in psychological disorders at the end of the time period between the two groups, suggesting that even if selection is at work, it does not account for much of the observed difference. Other studies with larger samples followed over time (Booth & Amato, 1991; Wertlieb, Budman, Demby, & Randall, 1984) found significantly higher psychological stress levels at the beginning of the time period for persons divorcing in the future compared with those who were not, although controlling for these did not eliminate significant differences in distress following divorce.

The problem with interpreting these findings as support for the selection explanation is that the actual breakup of the relationship is often the last point in a long-term process of dissolution (Kitson & Morgan, 1990), causing a predivorce rise in distress that may extend back several years. Whether the dissolution process is a cause or effect of predivorce distress levels would be difficult to separate in the short-term panel studies most common in the literature. Panels that measure distress 5 to 10 years predivorce would do a better job of separating these effects because of the greater temporal distance from the event. High levels of psychological stress found many years predivorce are more likely to reflect enduring personality factors or more serious mental disorders and not the results of stress from a dissolving relationship. The panel data used here contain measures of psychological distress obtained up to 11 years before a divorce occurred. Finding that persons who divorce, compared with those who do not, have significantly higher stress levels many years before a divorce would provide some support for the selection argument. Of course, some marriages may be habitually conflict ridden and unsatisfying, leading to high distress levels well before the start of a dissolution process. In these cases, high levels of distress many years before the marriage may signal a chronic relationship problem independent of individual predisposing factors (Johnson & Booth, 1998).

It is important to differentiate between the type of social selection process that accounts for why divorced persons have higher psychological stress levels than the married and the more general issue of personality and social factors that predict selection into divorce (Gottman, 1994). High levels of psychological distress occurring well before a divorce may reflect underlying stable personality disorders, but many personal characteristics may make people poor marital risks without elevating their levels of psychological distress.

CRISIS THEORY

Crisis theory (Booth & Amato, 1991; Tschann, Johnston, & Wallerstein, 1989) views marital disruption as a life crisis that temporarily changes mental health. Marital disruption is considered a serious life challenge that stresses those involved in the process (Mastekaasa, 1992). A crisis involves a life events stressor that is a discrete, ob-
The role transition process leads to higher distress scores among divorced persons, we would expect psychological distress to increase until the actual divorce or permanent separation (Kitson & Morgan, 1990) and fall as the transition passes. Booth and Amato (1991) found evidence in panel data that psychological stress rose before the crisis and then returned to levels comparable to those of married individuals two years after the marital disruption. This supports the predictions of the stress model. Studies only examining the relationship between time since disruption and psychological symptoms (Amato & Partridge, 1987; Aseltine & Kessler, 1993; Brown, Felton, Whiteman, & Manela, 1980; Chiriboga, Roberts, & Stein, 1978; Plummer & Koch-Hattem, 1986), however, found no significant decline in distress as the distance from the event increased. These findings are not consistent with the crisis model to the extent that the distress is a consequence of transition process.

ROLE THEORY

Role theory focuses on the relatively constant and enduring, or chronic, stresses and strains of certain roles (Pearlin, 1999). Living as a divorced individual often involves social isolation, lack of social support, economic hardship, and added childcare responsibilities (for parents), or stresses associated with child responsibilities (custody arrangements, etc.; Booth & Amato, 1991; McLanahan, 1983; Ross, 1995; Waite, 1995). Not only are there likely to be more chronic stressors in the everyday life of the divorced compared with the married individual, but they may be less likely to effectively cope with these stressors because of less available social support (Avison, 1999).

According to role theory, the greater distress levels of divorced persons are a permanent feature of that state. Because no adjustment process is implied, length of time since the divorce should not be related to level of psychological distress. Re-marriage or cohabiting by divorced persons should reduce distress levels if this explanation is correct (Ross, 1995). Entering a new conjugal relationship would reduce the chronic stressors of single parenthood, increase the odds of greater economic stability, and increase the levels of available emotional and social support. Studies that found higher psychological stress levels among the divorced (Ross, 1995), and no decline in distress level over time from post-divorce (Amato & Partridge, 1987; McLanahan, 1983; Mastekaasa, 1994) provide evidence for the validity of the social role perspective.

DIFFERENTIAL EFFECTS

It has been proposed that the psychological consequences of divorce or permanent separation will depend on both pre- and postdivorce characteristics and conditions (Aseltine & Kessler, 1993; Booth & Amato, 1991). A particularly robust finding is the improved mental health status of persons leaving conflict-filled and low-quality marriages (Aseltine & Kessler; Booth & Amato; Gove, 1972; Menaghan & Lieberman, 1986; Tschann et al., 1989; Wheaton, 1990).

The resources and responsibilities available postdivorce also should condition the extent of the stress induced by this role. Previous studies have identified available income, education, number of close friends, residence, presence of children, and related factors that influence the course of postdivorce stress (Aseltine & Kessler, 1993; Berman & Turk, 1981; Booth & Amato; Doherty, Su, & Needle, 1989; Goetting, 1981; Hetherington, Cox, & Cox, 1978; Menaghan & Lieberman, 1986; Tschann et al., 1989). Many of these factors were tested in cross-sectional studies. Further tests of some of these effects are still needed in the context of a more complete longitudinal model of factors producing psychological distress among divorced persons.

Findings from previous studies lead us to expect that the pattern of psychological distress associated with marital dissolution may differ by gender (Horowitz, Raskin White, Howell-White, 1996). Divorced women compared with men are more likely to have declines in standard of living and are more likely to have custody of the children. Both of these may increase the chronic stress of the divorced role for women. Studies have found that single mothers (including divorced) report higher psychological distress than married mothers (Avison, 1999). On the other hand, sev-
eral studies have shown that in regard to health and psychological well-being men generally benefit more from a marital relationship than do women (Waite, 1995).

MODEL TESTED

In this article we develop and test a model that can simultaneously evaluate the extent to which these theoretical explanations account for the higher levels of psychological distress observed for divorced compared to married persons. The model is designed for multiple-wave panel data in which observations of psychological distress, current marital role, date of the occurrence of marital events (divorce, remarriage, cohabitation, etc.), and other background variables are measured at each wave.

The data to be analyzed by this model are organized as a pooled-time series. In such a data set, each wave of observations for each individual in the sample is represented by a separate record (Johnson, 1995) so the total sample size is the number of individuals times the number of waves. This data structure violates the assumption of independent observations because the same respondent contributes more than one record to the data set. In this situation, ordinary least squares regression analysis is inappropriate. Two different estimators—fixed effect and random effects—have been developed for data organized in this manner (Allison, 1994). The random effects model is a generalized least squares solution in which the model parameters are solved through a weighted combination of within- and between-individual covariances (Hsiao, 1986). This method allows variables to be included in the equation that do not vary over time for individuals (e.g., gender, race) and is normally more efficient than the fixed effects estimator (Allison).

The fixed effects estimator for pooled time-series models is based entirely on variation over time within individuals in the sample. Variables that are invariant over time must be excluded from the model. In addition, because information about variation between individuals is not used to estimate parameters, the method is less efficient and standard errors are higher.

The fixed effects estimator, which is limited to variables that vary over time within individuals, can produce estimates that are net of all observed or unobserved differences between individuals that are time-invariant. The estimates are not biased by cohort or selection effects or the effects of individual differences that are time-invariant (e.g., background variables, genetic tendencies, stable personality predispositions).

For this analysis, the random effects model is the estimator of choice because one of the three explanations we want to investigate involves a between-individual effect—the effect of social selection. Because of the stronger inferences possible with the fixed effects solution, we also estimate models by fixed effects with the time invariant variables excluded. Doing this allows us to evaluate the possibility of specification errors in the random effects solution. Both models require at least two time periods for each individual in the sample but allow the number of time periods included to vary among individuals. This advantage is important in the analysis of multiple-wave survey panels in which attrition is common (Johnson, 1995).

The basic random-effects model for a pooled data set with i individuals and t time periods is as follows:

\[ P_i = u + b_1S_{i1} + b_2M_{i1} + b_3D_{i1} + b_4W_{i1} \\
+ b_5C_{i} + b_6T_{it} + b_7F_{it} + b_8X_{ki} \\
+ b_9Z_{jit} + e_{it}, \]

where \( u \) is a constant term, \( e_{it} \) an error term, and the bs are regression coefficients. The variables are as follows: \( P_i \) psychological distress; \( S_{i1} \) a measure of social selection (1 = individuals who ever divorced or permanently separated, 0 = all others); \( M_{i1} \) a measure of whether the respondent had been in at least one previous marriage before the first wave (1 = previously married, 0 = all other); \( D_{i1} \) current divorce status of individual i at time t (1 = divorced / permanently separated, 0 = all other); \( W_{i1} \) current widowed status (1 = widowed, 0 = all other); \( C_{i1} \) current cohabiting status (1 = cohabiting, 0 = all other); \( T_{it} \) a measure of the time to a marital disruption; \( F_{it} \) a measure of the time since a marital disruption; \( X_{ki} \) time-invariant control variables; and \( Z_{jit} \) time-varying control variables.

The \( T_{it} \) and \( F_{it} \) terms that measure proximity to marital disruption are used to estimate crisis effects. We hypothesize that if the crisis model is correct, we will find a significant increase in distress as the divorce approaches and a significant decrease following the disruption. In the models tested, these were included in reverse coded form so that the higher the score, the closer the disruption event. This coding also facilitates interpreta-
tion of the parameters associated with the social selection effect ($S_t$). If the social selection model is correct, we hypothesize that $S_t$ will have a significant positive effect on psychological distress net of other factors. The parameter $S_t$, coded as a dummy variable to indicate whether the person ever divorced or was permanently separated during the course of the study period, can be interpreted as the effect on distress well before or after a past or future disruption—not one close to the event. Current marital role is measured using three dummy variables ($D_{it}$, $W_{it}$, and $C_{it}$), with the reference (omitted) group being the currently married. If the social role explanation is correct, we hypothesize that the difference between the currently married and the currently divorced will be statistically significant not of other factors included in the model. Random and fixed effects estimates for this model were computed with the xtreg procedure in STATA (StataCorp, 1997).

Models were also developed that included interactions between a number of the variables to test for differential effects on distress. Differential effects by gender for a most of the variables in the model were tested by including sets of interaction terms. Marital happiness in the wave immediately preceding a disruption was introduced into the model to evaluate if the effects of a disruption on psychological distress differed for persons departing a relatively unhappy versus a relatively happy marriage.

**METHOD**

**The Sample**

This study uses four waves of panel data spanning 12 years. The first wave was collected over the telephone in 1980, making use of a random-digit dialing national probability sample. Completed interviews were obtained for 2,033 persons aged 19–55 in intact marriages. The refusal rate in 1980 was 18%, and the overall response rate compared favorably with those for other telephone surveys (Groves & Kahn, 1979). Follow-up interviews with all respondents who could be located and agreed to participate were conducted in 1983, 1988, and 1992. The percentage of the original sample reinterviewed in each of the subsequent waves was 78% in 1983, 66% in 1988, and 58% in 1992. Attrition was greater for younger respondents, men, renters, southern residents, African Americans, and those residing in metropolitan areas, but differences were small (Booth, Amato, Johnson, & Edwards, 1993). Logistic regression models with attrition over the four waves as the dependent variable found no significant effect of psychological distress measured in 1980 on panel attrition. Attrition also was not significantly related to several measures of marital quality (e.g., marital happiness, divorce proneness) in 1980.

The analysis was based on a data set that pooled the results from the four waves of the study. Records for respondents were included if they had answered the survey in two or more of the waves. The total sample size analyzed here was 5,676 records from 1,593 different individuals represented in the sample, with the average respondent contributing 3.56 records to the pooled data set. During the 12 years of the study, there were 243 divorces and permanent separations and 84 remarriages. In the pooled data set, 804 of the records were contributed by those who ever divorced or permanently separated during the 12-year period. The number of divorces and remarriages may appear small for this size sample given that around one half of all marriages are expected to end in divorce, with the majority of them remarrying (Cherlin, 1981). Because the 1980 sample was representative of persons in intact marriages, many were first interviewed after surviving the earlier years of marriage (which have much higher dissolution rates). Therefore, we would expect much lower rates of divorce than would have been found if the 1980 sample had been restricted to newlyweds. A survival analysis of divorce probability for the 385 persons in the sample who were married 3 or fewer years in 1980 revealed an expected failure rate of 35% after 12 years of marriage, a figure that approaches the rate of marital failures reported for this cohort (e.g., Cherlin).

**Measures**

A scale created from five items included in all four waves was used to measure psychological distress. The first item was an indicator of mental distress derived from responses to the question: “Have you felt extremely unhappy, nervous, irritable, or depressed?” Those responding no were coded 0, yes in the past 3 years were coded 1, yes recently were coded 2. Those who responded “recently” were asked in a follow-up question whether this led them to cut down on activities for several days or more. Those responding yes were coded 3. The second item was the respondent’s self-rating of their health status: “In general, would you say your own health is (1) excel-
lent, (2) good, (3) fair, or (4) poor?” A third item tapped global happiness: “Taking all things together, how would you say you are these days? Would you say you are (1) very happy, (2) pretty happy, or (3) not too happy?” The last two items were perceptions related to life satisfaction: “Would you strongly agree, agree, disagree, or strongly agree that things are as interesting as they were?” and “Would you strongly agree, agree, disagree, or strongly agree that you have gotten what expected out of life?” These five items were coded so high scores indicated more stress, z scored (based on means and standard deviations from the 1983 wave), and combined in a scale that was the average z score of the items responded to coefficient alpha reliability for the scale was .65; none of the items could be excluded without reducing alpha.

A more complete scale of psychological distress was available in some of the waves, but we used the five-item version because it was the only one available in all four waves. Waves 2, 3, and 4 contained 24 items measuring psychological stress/demoralization, including 8 items from Langner’s 22-item index of psychological disorder (Langner, 1962). To assess the validity of the five-item measure, we compared it with more complete scales in the 1983 wave (which had the largest sample size with responses to all 24 items). We correlated the five-item scale with a summed scale of the eight Langner items and a 17-item scale created from factor and scale analyses of all 24 stress items. The five-item stress scale is correlated at .58 with the eight-item Langner scale (α = .69) and .87 with the more reliable (α = .86) 17-item stress scale, giving us confidence that our findings can be generalized to more complete measures of psychological symptoms.

Measures of social selection, social role, and crisis were created as discussed in the basic pooled-time series model presented above. Social selection was measured by two indicators. The first was a dummy variable coded 1 if during the 12 years of the panel study (1980–1992) respondents ever divorced or permanently separated and coded 0 otherwise. The second dummy variable indicated if the respondent had been in another marriage before 1980 (1 = yes; 0 = no). Both measures are time invariant.

The measures of social role are time varying and were created from the respondent’s current marital status at the time of the interview. Three dummy variables measured whether the respondent was currently divorced or separated, widowed, or cohabiting. The omitted (reference) group was the married.

To measure the effect of crisis, two variables were created to ascertain the proximity of the interview to and from a marital disruption. Respondents who were divorced or permanently separated were asked to report time since this disruption had occurred (in months). From this measure and the decimal year of the interview, we calculated the exact decimal date of the disruption. For respondents divorcing more than once during the panel period, the earliest divorce was selected and interview waves following the second divorce were excluded from the data set. The small number of persons (24) with missing information on the time of the disruption were assigned the mean number of months from the disruption reported by respondents with complete data in the same wave.

Use of decimal years in the coding allowed precise estimates of duration. Once the exact decimal date of the disruption was created, the distance from this event was calculated for each respondent for each wave by comparing the decimal year in which the interview took place with the decimal year of the disruption. Decimal years to and from a disruption were coded as separate variables. For example, if a divorce occurred in June 1987 (decimal year 87.5) and the respondent was interviewed in the first wave in August 1980 (decimal year 80.67) and in Wave 2 in May 1992 (decimal year 92.42), then the coding for years to the disruption would be 6.83 decimal years in Wave 1 and the coding for years from the disruption would be 4.92 decimal years in Wave 4. Persons not experiencing a disruption during the 12 years of the panel were coded 0 on both duration measures.

The actual measures of decimal years to and from the disruption used in the regression models were transformed to aid in the interpretation of the regression parameter estimates. The number of years to and from the disruption was subtracted from 5 so persons closest to the disruption had the highest scores. All persons 5 or more years from the disruption were assigned a score of 0. For example, for the years to a disruption measure, a person interviewed 1 year before their marriage disrupted would be coded 4 (5−1) for that wave, whereas a person 7 years from disruption would be coded 0 (5−7 = −2, with all negative numbers recoded to zero). This procedure allows us to interpret the regression coefficient associated with the social selection variable as the effect on distress of persons who are still five or more years
Marital Disruption and Psychological Distress

TABLE 1. DESCRIPTIVE STATISTICS FOR VARIABLES IN THE POOLED DATASET

<table>
<thead>
<tr>
<th>Variables</th>
<th>$M$</th>
<th>$SD$</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>Psychological stress</td>
<td>−.0635</td>
<td>.6551</td>
<td>−1.742</td>
<td>2.984</td>
</tr>
<tr>
<td>Social selection</td>
<td>.1350</td>
<td>.3417</td>
<td>.000</td>
<td>1.000</td>
</tr>
<tr>
<td>Role (divorced/separated = 1)</td>
<td>.0463</td>
<td>.2102</td>
<td>.000</td>
<td>1.000</td>
</tr>
<tr>
<td>Widowed (yes = 1)</td>
<td>.0106</td>
<td>.1023</td>
<td>.000</td>
<td>1.000</td>
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<td>Cohabiting (yes = 1)</td>
<td>.0076</td>
<td>.0867</td>
<td>.000</td>
<td>1.000</td>
</tr>
<tr>
<td>Decimal years from divorce</td>
<td>.2294</td>
<td>1.1600</td>
<td>.000</td>
<td>13.477</td>
</tr>
<tr>
<td>Decimal years to divorce</td>
<td>.3484</td>
<td>1.4960</td>
<td>.000</td>
<td>11.555</td>
</tr>
<tr>
<td>Divorced before 1980</td>
<td>.0937</td>
<td>.2915</td>
<td>.000</td>
<td>1.000</td>
</tr>
<tr>
<td>Respondent’s age</td>
<td>40.7012</td>
<td>10.0776</td>
<td>17.000</td>
<td>68.000</td>
</tr>
<tr>
<td>Gender of respondent ($F = 1$)</td>
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<td>.4878</td>
<td>.000</td>
<td>1.000</td>
</tr>
<tr>
<td>Years of education</td>
<td>13.8365</td>
<td>2.6667</td>
<td>.000</td>
<td>30.000</td>
</tr>
<tr>
<td>No. of children in household</td>
<td>1.2736</td>
<td>1.2092</td>
<td>.000</td>
<td>7.000</td>
</tr>
<tr>
<td>Presence of children (yes = 1)</td>
<td>.6422</td>
<td>.4794</td>
<td>.000</td>
<td>1.000</td>
</tr>
<tr>
<td>Total family income (in $1000)</td>
<td>36.3150</td>
<td>16.4742</td>
<td>2.500</td>
<td>65.000</td>
</tr>
</tbody>
</table>

Note: $N = 5,676$.

from the marital disruption. We chose 5 years as the point beyond which the measure would be coded as 0 based on an exploratory analysis of the scatterplots of years to and from disruption on psychological distress. Lowess (Fox, 1991) nonparametric smoothed regression lines were fitted to the scatterplot of psychological distress by years to and from disruption. These curves showed that the regression line further than 4 to 4 years predisruption was flat. Closer to the disruption, the lowess analysis showed a linear increase in psychological distress. The curve of psychological distress by time from disruption was basically flat in both the lowess and regression models. The same coding was used for both in the interest of consistency, however. Alternative models for the effect of distance from disruption using different thresholds also were tested and are considered in the discussion section.

Demographic and background measures were included in the analysis as statistical controls. These were age of the respondent (in years), gender, years of schooling, number of children under 18 residing in the household, total household income at the time of the interview, and an indicator of whether the household income information was missing (less than 2% was missing). Marital happiness reported in the wave immediately preceding marital disruption was also used in the analysis of differential effects. The measure of marital happiness was an 11-item summed scale of items measuring happiness with various aspects of the marriage. It has been used in a number of previous studies (c.f., Johnson, White, Edwards, & Booth, 1986). Descriptive information for the pooled-time series data set on each of these variables is presented in Table 1.

RESULTS

The first regression model simultaneously estimates, with the random effects estimator, the influence of selection, crisis, and social role while controlling for basic background variables. It is presented as Random Effects Model 1 in Table 2. We first examine the evidence for the crisis explanation. The years from disruption were in the predicted direction (higher stress levels closer to the disruption), but the coefficient was close to zero and not statistically significant. Years to a disruption was significant ($p < .01$). Psychological distress scores predicted by the model for persons immediately before the disruption (who would have a score of approximately 5 on this measure) were around one third a standard deviation above persons 5 years from a disruption (who score 0 on this measure). This finding suggests that going through marital dissolution before the actual disruption is stressful and becomes more so as individuals move toward actual breakup of the relationship. Because scores did not decline following disruption, the crisis explanation appears inadequate.

The social selection variable in this equation measures the effect on psychological stress for persons who experienced a disruption but were 5 or more years away from it. The positive sign of the regression coefficient indicated slightly higher distress scores for persons 5 or more years pre- or postdisruption and was significant at the .05 level.
### Table 2. Regression Estimates for Pooled Time-Series Models with Psychological Distress as the Dependent Variable

<table>
<thead>
<tr>
<th>Variable</th>
<th>Random Effects Model 1</th>
<th>Random Effects Model 2</th>
<th>Random Effects Model 3</th>
<th>Random Effects Model 4</th>
<th>Random Effects Model 5</th>
</tr>
</thead>
<tbody>
<tr>
<td>Time from disruption</td>
<td>.0068</td>
<td>.0077</td>
<td>.0076</td>
<td>.0063</td>
<td>.0301</td>
</tr>
<tr>
<td>Time to disruption</td>
<td>.0635**</td>
<td>.0699**</td>
<td>.0620**</td>
<td>.0628**</td>
<td>.0991**</td>
</tr>
<tr>
<td>Social selection</td>
<td>.0981*</td>
<td>.1044*</td>
<td>.1044*</td>
<td>.2545**</td>
<td>.2545**</td>
</tr>
<tr>
<td>Previous divorce</td>
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<td>.1367**</td>
<td>.2145**</td>
<td>.2118**</td>
<td>.2118**</td>
</tr>
<tr>
<td>Divorced/permantly separated</td>
<td>.3022**</td>
<td>.2821**</td>
<td>.2870**</td>
<td>.2912**</td>
<td>.3228**</td>
</tr>
<tr>
<td>Widowed</td>
<td>.1359</td>
<td>.1076</td>
<td>.1226</td>
<td>.1180</td>
<td>.1228</td>
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<tr>
<td>Cohabiting</td>
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<td>−.2986**</td>
<td>−.2962**</td>
<td>−.4604**</td>
<td>−.4535**</td>
</tr>
<tr>
<td>Age</td>
<td>.0034**</td>
<td>.0091**</td>
<td>.0058**</td>
<td>.057**</td>
<td>.0053**</td>
</tr>
<tr>
<td>Gender</td>
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<td>−.0203</td>
<td>−.0219</td>
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<tr>
<td>Years of schooling</td>
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<td>−.0267**</td>
<td>−.0267**</td>
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<tr>
<td>Number of children in household</td>
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<td>.0174*</td>
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<td>Total family income</td>
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<td>−.0021**</td>
<td>−.0021**</td>
<td></td>
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<td>Missing income information</td>
<td>−.0874</td>
<td>−.0887</td>
<td>−.0900</td>
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<tr>
<td>Child present × previous divorce</td>
<td>−.1298*</td>
<td>−.1296*</td>
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</tr>
<tr>
<td>Child present × cohabiting</td>
<td>.3356*</td>
<td>.3537**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Lag happy × time from disruption</td>
<td>.0948**</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Lag happy × time to disruption</td>
<td>−.0736*</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Lag happy × divorced or permanently separated</td>
<td>−.3168**</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>.1992</td>
<td>−.4535</td>
<td>.0970</td>
<td>.0984</td>
<td>.1208</td>
</tr>
<tr>
<td>$R^2$</td>
<td>.044**</td>
<td>.232**</td>
<td>.051**</td>
<td>.052**</td>
<td>.063**</td>
</tr>
</tbody>
</table>

Note: Number of records = 5,676; number of individuals = 1,593.

* $p < .05$, **$p < .01$.

This provides some support for the social selection explanation. Nonetheless, the size of the effect is small (only about one third that of the divorce effect), suggesting that much of the difference is not accounted for in social selection.

The model estimates do provide substantial support for the social role explanation. The effect of divorce or separation was strong and significant. Persons divorced at the time of the interview had a psychological distress score around one third of a standard deviation higher than did married persons (the omitted group). Widows had higher distress scores as well, but the effect was not significant. A third indicator of social role is whether a divorced person is cohabiting. Cohabiting status had a strong effect on distress reduction, effectively eliminating the negative consequences of divorce. Because all cohabiters in this sample also are divorced, the net effect of cohabiting is to reduce the stressful effect of the divorce role. To compare the stress of living in a cohabiting relationship with the stress of living in a married relationship, we added the divorced and cohabiting coefficients, which yielded an effect size close to zero (.3022 − .2923 = +.0099). Divorced cohabiters had about the same stress levels as married persons.

Another significant effect noted in Model 1 was that persons previously married in 1980 had significantly higher stress scores than those in their first marriages in 1980. This could reflect either a social selection or role effect. We explore these alternative explanations in a later model.

Model 2 in Table 2 solves for the coefficients with the fixed effect estimator. This model controls for the additive effects of all measured and unmeasured differences between respondents, including selection and cohort effects because the coefficients are estimated using information on variation within individuals. Because measures that do not vary within individuals must be excluded, we drop social selection, previous divorce, and gender from the model. We also drop years of schooling because, although it does vary slightly within individuals, most of these changes appear to be reporting errors. The fixed effects estimates confirm the findings from the random models analysis, with the significant coefficients similar in magnitude and in the same direction.

Model 3 in Table 2 adds two variables—number of children in the household and total family income—that might play an intervening role in the effects of the divorce role on psychological stress. Because both the number of children pre-
sent in the household and income often changes following a divorce, we would expect the effects of the social role variables to be attenuated when introduced as controls. Although both number of children and income are significantly related to stress levels, including them in the equation has a minimal effect on any of the role or crisis coefficients. We cannot account for the greater stress of the divorced role by changes in income or number of children in the household.

The implications of the effects found in the model can best be visualized by plotting the changes in stress expected in the model as individuals age, dissolve their marriages, and remarry. In Figure 1 we plot the curves predicted by the random effects from model 1 for three hypothetical persons from ages 30–46. These trajectories are indicated by the solid lines in the figure. One person is in his or her first marriage and remains married over the age range (labeled continuously married in the figure). The second person also remains married throughout the period but had been married previously (labeled remarried). The third person divorces for the first time at age 38 and remarries at age 42 (labeled divorced and remarried). All other variables in the equation are held constant at their means during the period. The trajectories for predicted distress scores for the married and previously married are basically straight. There is little change in distress levels with age, and the previously married had slightly higher stress levels than the married. The person who divorces at age 38 begins at about the same level as the previously married person, due to the small but significant social selection effect observed. Five years before the disruption a sharp increase in distress level occurs. At the disruption, distress increases slightly and remains relatively high and at a nearly constant level. Although the figure shows a gradual decline following divorce and before remarriage, this decline was not statistically significant in the model. When they remarry, their psychological distress dropped to almost exactly the same levels predicted 5 or more years before a disruption. Because their stress levels also are approximately the same as those found for previously married persons when the study began, pairing these findings together supports the operation of a small but consistent selection effect.

The next set of models introduce interaction effects into the basic model to test for differential effect on the stress process. A number of interactions were tested but most were not statistically significant. Because previous studies have found differences between men and women in the effects of marital disruption on stress, we examined a number of interactions with gender. For example, the presence of children might make disruptions more stressful for women than men because women usually retain custody of children. We also might expect disruptions including children to be more stressful.

To fully evaluate the effects of gender on all the coefficients in Models 1 and 3, we compared the fit of a model, which allowed separate coefficients for men and women for all variables in the model with the fit of a model in which men and women had the same coefficients. There was no significant difference in the $R^2$ of these models, and only two of the coefficients had significantly different gender effects. These were the effects of education and the effects of income on distress. The effect of education was significantly larger for women, and the effect of income only held for men. We concluded that the effects of the divorce process on distress estimated in these models did not differ by gender.

Interactions between presence of children and other variables in the model (including the first seven independent variables listed in Table 2, Time to Disruption to Cohabiting) led to only two statistically significant effects, both involving number of children. The significant interaction effects are found in Model 4 in Table 2. Presence of children interacted significantly with the effects of previous divorce and cohabiting on distress. Number of children was dichotomized into presence (1) or absence (0) of children for the purposes of creating the interaction terms. When this interaction term is included in the equation, the additive term for previous divorce measures the
effect for people without children. Results show that those previously divorced without children report higher stress levels than do those previously divorced with children, although the distress levels for both were higher than for persons who had not been previously divorced. Previous divorce status had negative consequences, but the effect was stronger for those without children. This might reflect a selection effect—the category of persons with multiple marriages but no children may be more likely to include persons with serious preexisting mental disorders who are poor marriage risk.

According to the second significant interaction term, cohabiting following a divorce had a stronger effect on reduced reported distress (−.4604) among those without than it did for those with children present (−.4904−.3356 = −.1148). Although cohabiting with children present may increase social support and reduce stresses for the parent, the strains to parent-child relationships brought about by the introduction of another person to the household may mitigate some of these positive effects on the parent’s psychological well-being.

In Equation 5 of Table 2, we test for the effect noted in previous studies: Persons leaving stressful or unhappy marriages often exhibit improved psychological well-being. To test for this effect among those who divorced or permanently separated during the observed period, we divided the scores on the 11-item scale for marital happiness obtained in the interview immediately preceding the disruption into two groups (1 = relatively happy, 0 = relatively unhappy) by cutting the scale at the median score. We then multiplied this dichotomized lagged marital happiness indicator by decimal years to and from marital disruption, and added them to the equation. The results are presented in Model 5 of Table 2. A model was also tested that did not dichotomize lagged marital happiness and yielded similar findings. All three of the interaction terms were statistically significant. The implications of these interaction effects for predicted distress scores for persons relatively happy and relatively unhappy with their marriage in the pre-disruption wave are plotted as the dashed lines in Figure 1. We plotted expected scores for persons with relatively high (“Happy before divorce”) and relatively low marital happiness predisruption (“Unhappy before divorce”). The patterns of psychological distress predicted by the equation differ dramatically. For persons in troubled marriages (low happiness), the predicted level of stress increases steeply to the disruption then declines sharply following it. Distress level increased again (although this effect is not statistically significant) and then dropped following remarriage. For persons in a relatively happy relationship before the disruption, only a slight increase occurred up to the disruption, but the disruption itself produced a sharp gain in stress scores (although they do not reach the level predicted for a person leaving a more troubled relationship). Following the disruption, the stress levels declined with increased time from the disruption, indicating some support for a crisis model in this specific group. Remarriage did little to reduce distress in this group.

**Discussion**

Our findings clearly support a social role explanation for the higher stress levels of the divorced, provide only limited support for a stress model (and only for persons who were relatively happy before they divorced), and found evidence of a small social selection effect. Our findings appear to be somewhat at odds with those reported by Booth and Amato (1991), who analyzed the first three waves of data used here. It is possible that some of the assumptions about the models we tested, such as a 5-year threshold level for the effects of stresses caused by the dissolution process, might be producing erroneous results. Also, the relatively limited measure of psychological stress might be producing biased results. Because of these concerns, we evaluated the robustness of our findings in several different ways.

Two differences between findings by Booth and Amato (1991) and this study need to be reconciled. First, Booth and Amato concluded that the increased distress before divorce and decline observed afterward is consistent with a stress model. The increase before divorce is consistent in both studies, except that we map this more precisely by including a variable measuring time in decimal years to divorce, whereas they only use the wave intervals (3 or more years) to time the events. We only found a decline in distress following disruptions of relatively happy marriages—the effect was not significant for the total sample. This difference could reflect that we included a control for the effect of reforming unions after divorce, either marriages or cohabiting unions, and they did not. When Model 1 was re-estimated removing the social role variables (divorced, cohabited, and widowed) from the equa-
tion, the effect of years from disruption became statistically significant (p < .01) for the entire sample.

Our finding that persons leaving a troubled marriage immediately tend to reduce their stress levels is consistent with Booth and Amato's (1991) finding, but we believe the results support a social role rather than a crisis model. The psychological health of divorced persons tends to improve only when they move back into marriage or cohabiting relationships, that is, undergo another role transition. There is no overall evidence in our findings that remaining divorced becomes less stressful with time, a prediction of the stress model. Improved psychological health while still divorced was only found for those who were relatively happy with their marriage predisruption.

Aseltine and Kessler (1993) argued that controlling for remarriage, which we do here, may introduce a selection bias to the extent that people with psychological disorders have more difficulty remarrying, leading to higher distressed scores for those remaining divorced. If the selection explanation were correct, we would not expect moving from divorced to remarried status to show declines in stress scores, because this view would argue that less distressed persons remarry. In the basic pooled-time series model, the effect of remarriage is estimated by moving from divorced to married status, which is represented by the same coefficient (b for divorced or separated) as moving form married to divorce status (but having the opposite sign). We test if these were equal effects by including a term in the equation for remarriage (coded 1 if the married status represents a remarriage, 0 otherwise). This effect was not statistically significant. Because it measured the difference between moving from married to divorced and divorced to married, we conclude that the effect of moving from divorced status to marriage is as strong as moving from married to divorce status. These findings cast doubt on a selection explanation for the lower distress levels observed for the remarried.

If the social selection effect is primarily a function of a greater divorce risk among persons with relatively serious personality disorders who thus exit marriage quite quickly, it is possible that this sample may underestimate this effect. Persons in relatively short-term marriages are underrepresented in the 1980 sample of the currently married. Although we believe this effect is likely to be quite small, additional research with different samples is needed to confirm this. The findings provide support for the explanation that the causal direction is stronger from divorce to psychological distress rather than from psychological distress to divorce but do not rule out the likelihood that enduring personality characteristics may both increase the risk of divorce and the susceptibility of persons to developing psychological distress post-divorce (Gottman, 1994). Studies including personality measures predivorce would be needed to evaluate such an effect.

The analyses reported here set 5 years as a reasonable threshold, or outer limit, for the beginning effects of psychological stress on a marital dissolution process. The choice of 5 years was based on extensive exploratory analysis using lowess regression smoothing techniques. To evaluate the sensitivity of this choice of thresholds, however, we repeated the analyses with shorter and longer ones. The 5-year thresholds produced the best fitting model (measured by \( R^2 \)). The results did not differ in a substantively important way by threshold level used, except that with a 10-year threshold the social selection effect was in the same direction but no longer significant. We gather that our basic conclusions of a strong social role effect and weak selection and crisis effects are robust with respect to the choice of thresholds.

Although the high correlation between the five-item psychological stress scale we used and the more complete 17-item scale suggest that our findings can be generalized to other distress measures, we performed several analyses to evaluate the robustness of the findings by changing the distress measure. In Waves 2, 3, and 4, we were able to augment the scale we used with three of the more physiological items from the Langner scale that were available for both married and divorced respondents in these waves. Excluding the first wave from the analysis and using the augmented scale produced results supporting the same conclusions reached with the four waves and the five-item scale. The five-item measure includes a measure of the recency of psychological distress and of overall happiness—components sometimes separated in other studies (Booth & Amato, 1991). Because the measure for recency of stress asks about distressful periods within the last 3 years, as well as recently, it is possible that the failure to find declines in distress with increased duration from the disruption may represent reporting on earlier events. To test for this possible bias, we repeated the analysis, dichotomizing this measure into persons who did and did not experienced stress recently. A random effects probit model for
though the purpose of this research was not to identify the factors in divorced roles that account for higher stress levels, the effect clearly was not due to changes in economic conditions or a result of the presence of children. Other evidence supporting the social role explanation comes from analyzing the differential effect on distress by how happy individuals were in their marriage predisruption. Those in the most troublesome marriages (least happy) actually improved their psychological health postdisruption, suggesting that the effect of leaving a stressful role was more important than the crisis of the disruption.

The crisis model predicted an increase in distress levels premarital disruption and a drop following it. Overall, we only found a rise in psychological distress up to the divorce—no fall postdisruption. Our findings did not support the view that role transitions and other changes occurring at the time of the divorce produce temporary stresses that dissipate after a period of adjustment to new life conditions. The rise predisruption could be equally well explained by social role theory. As a marriage begins to dissolve, living in one headed toward disruption becomes more stressful. Living in an unhappy and conflict-filled marriage has been shown to be highly stressful (Ross, 1995), a finding we confirmed here. Controlling for remarriage and cohabitation, there was no evidence of an overall significant decline in distress scores postdisruption. The only support for the crisis model was found among respondents who left marriages in which they were reasonably happy in the wave immediately preceding divorce. Among those unhappy with their marriages predisruption, psychological health appeared to decline rather than improve commensurate with increase in time postdisruption.

There was some support for a social selection effect, although this effect only accounted for a small amount of the total effect of divorce status. Because we could only measure social selection indirectly, based on the level of distress in the marriage more than 5 years before the divorce occurred, further research is needed before we can conclude that selection factors only have a small effect. Other studies that measure personality and social factors before the marriage occurred would provide a better test. What we do know from this study is that elevated distress levels of the divorced primarily occurred in the few years preceding the divorce and were not the result of a more chronic distress condition.

Although women are more likely to have cus-
tody of the children and report more decline in standard of living following divorce than do men, there was no evidence in this study that the stress process with divorce differed by gender. Men and women may respond to stress in different ways (Horowitz et al., 1996), so research studies with measures of multiple stress outcomes would be needed to further explore the gender effect.

Because we found that divorced persons’ psychological health improved when they formed a new relationship (either cohabiting or remarried), the social support and attachment benefits of an intimate living relationship (Ross, 1995) compared with divorced status may be the key explanatory factor in the higher distress levels among the divorced.

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