

**EMPIRICAL ESSAYS ON DIVORCE
AND CHILD INVESTMENT**

By
John M. Nunley

A Dissertation Submitted to the Graduate School at Middle
Tennessee State University in Partial Fulfillment of the
Requirement for the Degree

Doctor of Philosophy/Economics

Murfreesboro, TN

August 2008

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TO MY PARENTS

RONALD AND MARY LYNN NUNLEY

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ABSTRACT

This dissertation consists of four chapters on economic, legal, and demographic determinants of divorce rates and child investment. The first chapter, “The Effects of Household Income Volatility on Divorce,” examines whether fluctuations in household income affect individual-level divorce propensities, finding that household income volatility plays a significant role in determining marriage outcomes. I find statistical evidence indicating that positive and negative household income volatility increases the probability of divorce for men and women. By contrast, positive shocks to household income lower the risk of divorce for lower-household income individuals, and increase the divorce risk for those with higher levels of household income. Negative shocks to household income raise the probability of divorce regardless of the level of household income. The second chapter, “Inflation and Other Aggregate Determinants of the Trend in U.S. Divorce Rates since the 1960s,” focuses on whether increases in the inflation rate in the 1960s and 1970s contributed to the sharp rise in the divorce rate. Inflation is found to have substantial, positive, persistent effects on the divorce rate. The third chapter (co-authored with Joachim Zietz), “Explaining the Evolution of the U.S. Divorce Rate,” extends research on determinants of the divorce rate by considering whether increased access to oral contraception contributed to the sharp rise in the divorce rate. We also explicitly take into account the potential impact of the Vietnam War on the divorce rate. Our econometric evidence supports the idea that increased access to oral contraception and the Vietnam War shifted the divorce rate to a new, higher level. Opposite to previous work, we find a negative relationship between the divorce rate and the rising economic

independence of women, for which their participation in higher education proxies. The fourth chapter (co-authored with Alan Seals), “Child-Custody Reform and Marriage-Specific Investment in Children,” considers whether the post-divorce allocation of children affects how married couples invest in their children, measured as children’s private school attendance. The econometric evidence indicates that the post-divorce allocation of children has negative consequences for children living in in-tact households, with the negative effects becoming larger in states that have property-division laws that favor one spouse over another.

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CHAPTER 1

INTRODUCTION

Economic analysis of family behavior has grown substantially since the seminal work of Gary Becker in the 1960s and 1970s. Economists who conduct research on family behavior have examined marriage, divorce, fertility decisions, investment in children, household labor-supply decisions, consumption patterns, and bargaining power within households. This dissertation, which is composed of four essays, encompasses and extends existing research on the economics of the family by examining the effects of household income volatility on individual-level divorce propensities, the effects of inflation and other macroeconomic variables on national-level divorce rates, the role of the rising economic independence of women, increased access to oral contraception, and military conflict on aggregate divorce rates, and the impact of child-custody reform on marriage-specific investment in children.

Chapter 1 empirically examines the effects of household income volatility on individual-level divorce probabilities. In this essay, I use two measures of household income volatility, account for time-invariant, unobserved match quality, address the potential endogeneity issue associated household income and divorce rates, and examine the divorce behavior of lower- and higher-household income groups. The findings indicate that household income fluctuations have a substantial impact on individual-level

divorce propensities. Negative household income shocks raise the probability of divorce for both men and women, while the results are mixed for men and women when the household income shocks are positive. The results also differ between lower- and higher-household income individuals. Positive household income volatility decreases the probability of divorce for lower-household income individuals and raises the divorce risk for higher-household income individuals. Negative household income volatility raises the risk of divorce regardless of the level of household income.

Chapter 2 moves from analyzing individual-level divorce rates to examining macroeconomic factors that affect aggregate-level divorce rates in the United States. One of the advantages of my approach to this problem is the modeling of the divorce rate as an unobserved variable. This approach circumvents the problem of omitted variables and unobservables, which can bias estimates. Analyzing data from 1955 to 2004, the main finding is that inflation has a substantial, positive effect on the divorce rate. This study also encompasses covariates used in previous research, including the unemployment rate, female participation in higher education, and economic growth. The results for the unemployment differ across models, finding evidence of both positive and negative effects. The results for the relationship between economic growth and the divorce confirm previous research, indicating a positive relationship. An increase in female participation in higher education increases the divorce rate; however, the effect is small.

Chapter 3 (co-authored with Joachim Zietz) extends the work done in Chapter 2 by revisiting some past variables shown to affect national-level divorce rates and incorporating new covariates as predictors of the divorce rate. We employ a variety of

methods: single-equation models and systems estimators. While we find the systems estimator to be more appropriate, the results for the single-equation models are largely consistent with those from the multivariate model. We attempt to answer the question: what factors contributed to the sharp rise in the divorce rate in the 1960s and 1970s? Our econometric evidence supports the idea that increased access to oral contraception and the Vietnam War shifted the divorce rate to a new, higher level from the early-1960s to the mid- to late-1970s. We also find that the divorce rate and female participation in higher education are negatively related both in the short and long run. This result contests a large body of previous research.

Chapter 4 (co-authored with Alan Seals) examines whether child-custody reform in the early-1980s affected marriage-specific investment in children, measured as children's private school attendance. Child-custody reform alters the post-divorce allocation of children. As such, divorce-threat bargaining models predict that changes in policies that alter the post-divorce allocation of marital resources (including children) alter within-marriage distribution. Most research on custodial allocations focus on post-divorce investment by parents, with the investment behavior of noncustodial parents as the primary objective. By contrast, this chapter examines whether changes in the allocation of children (i.e. shared custody) affected the within-marriage investment behavior of spouses. We find that joint-custody reform negatively affects marriage-specific investment in children, with the effects becoming larger in states that have property-division laws that favor one spouse over the other.

CHAPTER 2

THE EFFECTS OF HOUSEHOLD INCOME VOLATILITY ON DIVORCE

2.1. INTRODUCTION

The theory developed by Becker et al. (1977) contends that “surprises,” whether positive or negative, should have a positive effect on the probability of divorce. In Becker et al.’s framework, household income volatility serves as a proxy for surprises or unexpected events. Their model predicts that household income volatility increases the risk of divorce because unexpected changes alter the couples’ expected returns from marriage. Negative shocks to household income could lower the returns from marriage below a particular threshold level, which may lead to divorce. Positive shocks could induce a self-reliance effect, which may also increase the risk of divorce. It could also be that positive or negative household income shocks change the value of the outside option, which is the divorced state.

A number of studies examine the effect of earnings shocks on consumption and other economic outcomes.¹ However, there have been few studies that examine the effects of earnings shocks on divorce. Previous attempts to measure the effects of earnings shocks on divorce have used actual minus predicted earnings (Becker et al. 1977), changes in

¹ For example, see Charles (1999), Cullen and Gruber (2000), Stephens (2001, 2004), and Blundell and Pistaferri (2003).

predicted earnings capacities (Weiss and Willis 1997), job displacement (Charles and Stephens 2004), and relative spousal income volatility (Hess 2004).

This paper provides an alternative proxy for earnings shocks by examining the effect of household income volatility on divorce. The measures of earnings shocks used in this paper differ from previous measures in two distinct ways: spousal incomes are jointly considered and positive and negative household income shocks are separately identified. Because the decision occurs over time, I use panel data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79). I construct two measures of household income volatility, one of which exploits potentially exogenous variation in the occupations of individuals. The first measure is the coefficient of variation over three-year periods. The second measure decomposes household income into permanent, transitory, and volatility components. The use of the two volatility measures relates to the differing assumptions governing the measures and the potential endogeneity problem associated with the first measure (i.e. coefficient of variation). The decomposition approach uses additional information in the first-stage regression that is shown to have no predictive power in the divorce equations, but has significant predictive power in the first-stage models.

The empirical models also include indicator variables that separately capture the effects of negative household income movements through time, since whether household income volatility stems from positive or negative household income shocks is otherwise not identified. For men, the effects of positive and negative household income volatility are statistically significant and positive; the effects are also consistent across the two

volatility measures. Statistical significance is not consistent across the volatility measures for women. The results for the coefficient of variation only indicate a statistically significant, positive effect with respect to positive household income volatility. The decomposition approach suggests the opposite; negative shocks to household income increase the probability of divorce. Positive shocks have no effect on the probability of divorce for women.

Supplementary models are also estimated for two different income groups: (i) lower-household income individuals and (ii) higher-household income individuals. The effects of household income volatility on divorce differ across the two income groups. For example, neither positive nor negative household income volatility changes the divorce risk for lower-household income men; however, increases in both raise the divorce risk for higher-household income men. Negative household income shocks increase the divorce risk and positive household income shocks decrease the divorce risk for women in the lower-income group. Both positive and negative household income shocks increase the divorce risk for women in the higher-income group.

2.2. BACKGROUND INFORMATION

2.2.1. The Role of Household Income Volatility in Marriage and Divorce Models

The theory of marriage developed by Becker (1973, 1974) contends that individuals sort into marriage based on economic and non-economic characteristics.² Becker's theory

² Becker (1973, 1974) also contends that couples marry to specialize in market and household work and to achieve higher levels of marriage-specific investment such as additions to human capital, property, and children.

contends that individuals marry others with like characteristics. For example, couples with similar education levels, intelligence, social background, race, and religion are more likely to marry and to be better matches once married. Marrying on the basis of like characteristics implies that the traits are complementary, also referred to as positive assortative mating. In contrast, Becker's theory suggests that negative assortative mating occurs with respect to earnings, which implies that spousal earnings are substitutes.³ If couples sort into marriages based on earnings, household income volatility should affect divorce behavior because the returns from marriage would change.

In their seminal article, Becker et al. (1977) posit that a decrease in the expected value of characteristics in which positive marital sorting occurs increases the risk of divorce. They also contend that unexpected events and differences in actual minus expected values of characteristics should also increase the risk of divorce.⁴ All of these factors, of which measures of household income volatility could proxy, affect divorce propensities by changing the returns from marriage.

The effects of household income volatility on divorce could also be tested in the context of a divorce-threat bargaining model.⁵ Divorce-threat bargaining models imply that the incomes received by husbands and wives shift bargaining power between spouses. For example, spouses who have higher incomes exert greater bargaining power,

³ Earnings may be substitutable because one spouse may specialize in market work while the other spouse may specialize in home production. Becker (1973, 1974) contends that specialization between spouses provides additional returns to marriage.

⁴ Their theory also predicts that increases in age at marriage, investments in marriage-specific capital, and increases in the length of marriage are expected to reduce divorce propensities. Additional educational attainment is expected to have an ambiguous effect on divorce.

⁵ See Lundberg and Pollak (1996) and Bergstrom (1996) for detailed discussion of models of household behavior including divorce-threat bargaining models.

as they possess more control over family resources. Individuals who divorce value the outside option, which is divorce, more than the option of remaining married.⁶ For example, large negative shocks to the husband's income may increase the value of the outside option for the wife; thus, divorce could occur. Alternatively, a large positive shock to the husband's income may either increase or decrease the value of the outside option for the wife. On the one hand, it could be that positive household income shocks to the husband's income stabilize marriages through increases in the returns from marriage. However, it could also be that a positive shock to the husband's income induces the wife to file for divorce because of the benefits associated with the divorce settlement. The directional effect of positive household income volatility on divorce may also depend on the underlying property-division laws in a particular state, as women typically receive greater benefits in community property states and men typically receive greater benefits in common-law states.

The model developed by Hess (2004) provides another theoretical channel to test the implications of household income volatility on divorce. As in Hess (2004), the decision to marry could provide couples with a way to hedge against income risk. Negative shocks to one spouse's income can be offset by the other spouse's income. Hence, marriage offers spouses a form of consumption insurance. If couples use marriage as a hedge against income risk, large reductions in household income could induce marital instability because of the ineffectiveness of the marriage hedge.

⁶ In a number of models, the threat point is interpreted as the utility associated with the divorced state; however, in others, the threat point is a noncooperative equilibrium within marriage. See Lundberg and Pollak (1993, 1996) and Bergstrom (1996) for a discussion of noncooperative marriage models.

2.2.2. Previous Empirical Findings

In contrast to Becker's predictions, there has been little empirical support for negative assortative mating based on earnings.⁷ However, Zhang and Liu (2003) find weak evidence of negative assortative mating with respect to wage rates. Smith (1979), Becker (1981), and Nakosteen et al. (2004) find evidence of positive assortative mating with respect to earnings and earnings residuals.⁸ Similarly, Chadwick and Solon (2002) find a substantial elasticity between daughters' (wives') earnings and the earnings of the family in which they were raised. The authors also find a similar elasticity between the daughter's husband's earnings and the earnings of the daughter's family, which provides more support for positive assortative mating with respect to earnings. The evidence supporting positive assortative mating based on earnings suggests that household income shocks should affect divorce propensities.

In Becker et al. (1977), unanticipated events or earnings shocks, measured as the actual earnings minus predicted earnings, tend to raise the probability of divorce for both men and women.⁹ Whether the measure of unexpected events is positive or negative has no bearing on the statistically significant, positive effect on divorce. Weiss and Willis (1997) use NLS data to examine the effects of unexpected changes in predicted earnings

⁷ Lam (1998) develops a theoretical model that discusses potential reasons for the lack of empirical support for negative assortative mating on wages. Lam also provides a brief survey of the other findings with respect to assortative mating.

⁸ Nakosteen et al.'s (2004) results are likely to be more accurate, since the authors are able to observe individuals both before and after marriage. Most data sets do not allow the pre-marital characteristics of individuals who eventually marry to be identified.

⁹ Becker et al. (1977) also find that difficulties in conceiving children also raise the probability of divorce. This is also used as a measure of unexpected events or surprises.

capacity on the likelihood of divorce.¹⁰ They find that an increase in predicted earnings capacity decreases the probability of divorce for men; however, the effect is positive for women.

Charles and Stephens (2004) use PSID data to examine the effect of negative earnings shocks, measured as job displacement, on divorce. They examine three types of job displacement: (i) layoffs, (ii) plant closings, and (iii) disability. They find no evidence that plant closings or disabilities translate into a greater risk of divorce. However, layoffs positively affect divorce propensities. Since plant closings, disabilities, and layoffs have similar long-run earnings effects, they conclude that a spouse's non-economic suitability may play a more significant role than pecuniary matters in divorce decisions.

The majority of these studies, with the exception of Hess (2004), examine the effect of own earnings measures on divorce propensities. Conversely, Hess incorporates the incomes of both spouses and examines their correlation, mean difference, and relative variances.¹¹ Using NLSY data, Hess finds that increases in relative spousal income volatility increases the probability of divorce.¹² Hess also accounts for the potential endogeneity bias associated with income in the divorce equation by using exogenous variation in the occupations of individuals. One possible limitation of Hess's analysis is that positive and negative income shocks are not identified. Hess examines relative

¹⁰ Weiss and Willis (1997) also add to the literature by incorporating match quality into the divorce equation. The authors suggest that match quality has permanent and transitory components. If constant mean and constant covariance assumptions hold, match quality can be accounted for by including fixed effects in the divorce equation.

¹¹ The results indicate that increases in the correlation of spousal incomes tend to raise the probability of divorce. Hess also finds that the mean difference of spousal incomes is not statistically significant from zero.

¹² Hess (2004) uses the variance of the breadwinner's income relative to the other spouse's income to construct the volatility measure.

income variances of spouses as a measure of income volatility, which makes identifying positive and negative household income shocks difficult. My approach differs from Hess's (2004) and other research in three distinct ways: (i) I jointly consider spousal incomes, (ii) I separately identify the effects of positive and negative household income shocks, and (iii) I estimate the effects of household income volatility on divorce for lower- and higher-household income individuals.

2.3. DATA

I use data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79) to examine the effects of household income volatility on divorce. The NLSY79 is a nationally representative panel data set, which is appropriate for analyzing dynamic processes such as divorce decisions. In 1979, the survey began interviewing 12,686 respondents between the ages of 14 and 22.¹³ The NLSY79 surveyed individuals annually until 1994 and then biennially thereafter. Each survey collects information on demographics along with individual labor-market and familial characteristics. The survey provides a way to analyze divorce decisions because the necessary information is available and is consistently provided in all survey years.

In each of the survey years, the NLSY79 collects information regarding respondents' marital status. Specifically, the survey identifies never married, married, separated, widowed, and divorced individuals in each year. The fact that the survey identifies the

¹³ The original sample contained 6,283 women and an oversample of blacks, Hispanics, low-income whites, and military personnel. In 1984 and 1990, the military and the low-income white oversamples were dropped, respectively.

marital status of each individual in all years allows me to construct the appropriate sample with which to examine divorce behavior in response to household income volatility over time. The key explanatory variables use a measure of family income, which contains all sources of family income and is provided in each of the survey years. To obtain a real measure of family income, I deflate family income by the implicit price deflator for Gross Domestic Product.¹⁴

The estimation procedure uses the entire sample period. Because the survey is biennial after 1994, examining divorce decisions could present problems. It is possible for an individual to divorce twice in a two-year period. However, most divorces take considerable time to become finalized, especially if property and children are involved. It may also take a considerable time to find a new spouse.¹⁵

To construct the appropriate sample, all individuals who marry during the survey are identified.¹⁶ I exclude anyone who is married at the beginning of the survey, as information on the individual and their spouse is not available for the years that they were married before the survey began. After identifying individuals who marry over the course of the survey, I construct a marriage duration variable, which is used to construct the appropriate sample. Individuals with a missing value for the duration of marriage exit the sample. Therefore, in the years following divorce, individuals will receive a missing

¹⁴ I discuss the ways in which the household income volatility measures are constructed in Section 4.

¹⁵ Limiting the sample to surveys conducted annually does not change the signs and statistical significance of the household income measures. The magnitudes of the effects do change slightly; in several cases, the effects become larger.

¹⁶ Some research uses measures of marital dissolution as a measure of marital instability (e.g., Becker et al. 1977; Weiss and Willis 1997), which implies that outcome variable is not only divorce but separation as well. Since individuals are legally married if they report being separated, I count separated individuals as being married.

value for the marriage duration variable and as a result will exit the sample unless they remarry in the following years.¹⁷ The divorce outcome variable, which is a zero/one indicator variable, is formed by using the marriage duration variable. For example, an individual who marries in 1981 and divorces in 1989 would receive a zero for each year of marriage and a one for the year the individual divorces. If the individual does not remarry in 1990 or the years thereafter, the individual will no longer be in the sample because they will have missing value for the divorce outcome and the marriage-duration variable.¹⁸

Next, I exclude all married individuals who have household income less than \$20,000 and greater than \$200,000. The household income restrictions are used because one of the household income volatility measures is sensitive to low levels of household income (see discussion of equation (1) in Section IV). It is also unlikely that negative shocks to household income for high-income households would have the same effect because the financial stress would not be as great. There are also few observations for individuals with of household income in excess of \$200,000.

After deletions are made, the sample contains only individuals who have married at some point over the sample period and who have household income fitting the previously-mentioned criteria. Since I use two different measures of household income volatility, the number of observations and the number of individuals observed in the empirical models differ. For the first measure of household income volatility, there are

¹⁷ Individuals who remarry re-enter the sample because their marriage duration variable no longer has a missing value.

¹⁸ Individuals who have been widowed or have never been married receive missing values for the divorce outcome; thus, they are not in the sample.

608 men examined with 3,001 person-year observations and there are 646 women examined with 3,169 person-year observations. The model for the second measure of household income volatility has 1,658 person-year observations for men and 1,637 person-year observations for women. The number of men and women analyzed are 441 and 448, respectively.

In supplemental analyses, I examine lower- and higher-income individuals separately. To conduct the supplemental analyses, I partition men and women into two household income groups: (i) the \$5,000 to \$40,000 range and (ii) the \$40,000 to \$200,000 range. These household income restrictions roughly divide the sample in half and provide ample observations to estimate the divorce equations for the two household income groups. With the exception of the household income restrictions, constructing the sample for the supplemental models follows the same rules as the full sample.

Using data from the NLSY79 offers a way to follow young individuals into their adulthood. At the end of the sample period, individuals would most likely have reached or would be approaching the peak of their earnings potential. Since individuals should experience household income shocks over the sample period, the NLSY79 is well suited to analyze changes in divorce behavior in response to household income fluctuations,

2.4. ECONOMETRIC METHODOLOGY

Measures of income uncertainty or volatility differ based on the assumptions governing

the individual's expectations of future income flows (Robst et al. 1999).¹⁹ The first measure of household income volatility, the coefficient of variation over three year periods (CV), measures dispersion in household income over time.²⁰ Formally, the coefficient of variation is

$$CV_{i,t} = \frac{\sigma_{i,t}^*}{\mu_{i,t}^*}. \quad (1)$$

The subscripts i and t index individuals and time, respectively. The term σ^* is the standard deviation of household income over three year periods and μ^* is the average of household income over three year periods. The divorce equation includes the CV measure along with the log of real household income (HI) to examine the relationship between household income volatility and divorce.

The specification of the divorce equation is

$$y_{i,t}^* = c_i + \beta_1 HI_{i,t} + \beta_2 CV_{i,t} + \beta_3 (D_{i,t} \times CV_{i,t}) + \beta_4 \mathbf{X}_{i,t} + \varepsilon_{i,t}. \quad (2)$$

The variable y^* is a binary variable taking on a value of one when the individual divorces and zero when married; c represents an individual-specific fixed effect; HI and CV are defined above; \mathbf{X} is a vector of control variables, which includes the individual's age,

¹⁹ The measures of volatility used here resemble the techniques used by Haurin (1991) and Robst et al. (1999).

²⁰ Using the CV as a measure of household income volatility assumes that the individual possesses little knowledge concerning future household income flows. As a result, this measure of household income volatility may be inadequate. Most individuals would expect some household income growth as they gain experience and more job skills. Problems also surface for the CV measure when the mean value of household income is close to zero. When this is the case, the CV measure is very sensitive to large changes in the standard deviation of household income. As I discussed in the previous section, I address this issue by excluding individuals with household income below \$20,000. However, I relax the household income restrictions in order to examine the divorce behavior of lower-income households.

educational attainment, the number of children, regional indicators, and time indicators; and ε is an error term. The term $(D \times CV)$ is an interaction term that captures decreases in μ^* over time. The variable D takes on a value of one when μ^* at period t is less than μ^* at period $t - 1$. Therefore, interacting D and CV allows for the effects of negative and positive household income shocks to be isolated. The β_i are parameters to be estimated. Attention focuses on the parameters β_2 and β_3 in equation (2), which measure the effect of positive and negative household income volatility, respectively.

There is a potential endogeneity problem with respect to household income and, as a result, the CV measure.²¹ The inclusion of c eliminates time-invariant traits that may induce bias because of correlation between unobservables and the household income variables or the variables in \mathbf{X} . The estimates are consistent if unobservables are time-invariant and there is no simultaneity bias. If unobservables are not time-invariant, then results may still be biased. Adding fixed effects to the model proxies for match quality, as in Weiss and Willis (1997).²²

Including c does nothing to correct for the potential simultaneity bias associated with household income and divorce. Johnson and Skinner (1986) find that women begin

²¹ Ressler and Waters (2000) implement a simultaneous equation model of divorce and female earnings. Their results imply that single equation divorce models will most likely overstate the relationship between female earnings and divorce, unless the identification strategy incorporates additional exogenous information in the model.

²² In some cases, one may prefer to control for a wide-range of covariates. Because of data limitations, some necessary control variables are not available. Other research has shown the importance of age at marriage, religious upbringing, cohabitation, and the presence of children from previous marriages in divorce decisions (e.g., Becker et al. 1977 and Weiss and Willis 1997). Unfortunately, some of this information is not available in the NLSY79. However, many of these variables are time-invariant. Therefore, the influence of these variables can be removed from the model by including fixed effects.

increasing their labor supply as the probability of divorce increases.²³ This would ultimately result in a simultaneous relationship between increases in the risk of divorce and increases in household income. As wives increase their labor supply in response to an increase in the risk of divorce, household income would also increase.

The assumptions governing equation (1) and the potential endogeneity problem associated with HI and CV in equation (2) are the reasons for including an alternative measure of household income volatility. The second measure assumes that individuals have knowledge about their future income streams, which are based on observable labor-market characteristics. Individuals know the characteristics of other individuals and the income they receive for their labor-market characteristics (Robst et al. 1999). As a result, this measure should be a more realistic measure of household income volatility.

The second measure uses a first-stage regression of HI on the variables in \mathbf{X} and other variables expected to predict household income.²⁴ Formally, I estimate

$$\ln HI_{i,t} = \gamma_0 + \gamma_1 \mathbf{O}_{i,t} + \gamma_2 \mathbf{L}_{i,t} + \gamma_3 \mathbf{C}_{i,t} + \gamma_4 \mathbf{X}_{i,t} + \xi_{i,t}. \quad (3)$$

HI is defined above; \mathbf{O} represents the occupation indicators; \mathbf{L} represents individual labor-market characteristics including job tenure, a squared term of educational attainment, and an indicator labor union status; \mathbf{C} represents county-level variables including the unemployment rate, the percent of the population that is (are) black, Hispanic, medical doctors, high-school educated, college educated, employed in the manufacturing sector,

²³ Sen (2002) finds that Johnson and Skinner's (1986) results still hold, but notes that the increase in the labor supply of women due to increases in the risk of divorce has diminished over time.

²⁴ Equation (3) takes a log-linear functional form because the specification yields a better model fit (Heckman and Polachek 1974). There is also no need to control for individual-unobserved heterogeneity in equation (3) because the specification of the divorce equation will eliminate all fixed effects; thus, nothing is lost by estimating the first-stage regression by ordinary least squares (OLS).

employed in the retail sector, and employed in the public sector; \mathbf{X} is defined above; and ξ is an error term. The γ_i are parameters to be estimated.

The occupation indicators in \mathbf{O} potentially provide a source of exogenous variation with which to identify the effects of household income and household income volatility on divorce. An individual's occupation should not be correlated with the divorce variable. However, household income and the occupations of individuals should be correlated. Hess (2004) also uses occupation indicators to identify the effect of relative spousal income volatility on the probability of divorce.

The second household income volatility measure uses the predicted values of HI and the predicted value of ξ from equation (3) to specify the different components of household income. The predicted value of ξ represents the uncertain portion of household income (HI). Interest does not focus on the parameter estimates of equation (3). There is a need, however, to control for as many factors as possible that are expected to influence household income. Omission of a key variable could lead to the uncertain portion (ξ) of household income not being attributable to uncertainty, rather an omitted variable.²⁵ After estimating equation (3), the mean of the residual series over three year periods (μ^ξ) and the standard deviation of the residual series over three year periods

²⁵ Although there are most likely omitted variables in equation (3), I have included as many factors as possible. However, the NLSY79 is limited in that it does not provide a great deal of information on spouse's educational attainment, job tenure, and experience, all of which should affect household income. However, it could be that the omitted factors are exogenous to the individual. If in fact the omitted factors are exogenous to the individual, the unobserved portion of household income should represent the uncertain portion of household income.

(σ^ξ) enter the divorce equation along with the permanent income component (μ^I).²⁶

The terms μ^ξ and σ^ξ represent the transitory and volatility components of household income, respectively. For the second specification, the divorce equation is

$$\begin{aligned}
 y_{i,t}^* &= c_i + \theta_1 \mu_{i,t}^I + \theta_2 \mu_{i,t}^\xi + \theta_3 \sigma_{i,t}^\xi \\
 &\quad + \theta_4 (D_{i,t} \times \mu_{i,t}^\xi) + \theta_5 (D_{i,t} \times \sigma_{i,t}^\xi) \\
 &\quad + \theta_6 \mathbf{X}_{i,t} + v_{i,t}.
 \end{aligned} \tag{4}$$

The terms y^* , c , μ^I , μ^ξ , σ^ξ , D , and \mathbf{X} are defined above. The variable v is an error term. The θ_i are parameters to be estimated. The estimation procedure focuses on the parameters θ_3 and θ_5 , which measure the effects of positive and negative household income volatility, respectively. Summary statistics depicting the difference in the household income volatility measure used in equation (2) and equation (4) between individuals who divorce and those who do not divorce are shown in TABLE 1. As TABLE 1 indicates, individuals who divorce, on average, have both higher levels of positive and negative household income volatility for both household income volatility measures.

Equations (2) and (4) are estimated by ordinary least squares (OLS).²⁷ Estimating a binary outcome by OLS does present a problem: the predicted outcome is not

²⁶ The permanent income component is the average predicted value of household income over three year periods.

²⁷ Logit and probit specifications do constrain predicted outcomes. However, when fixed effects enter the logit specification, individuals who have time-invariant outcomes are dropped. The result is a large reduction in the number of observations, which leads to insignificant results. The incidental parameters problem has the potential to surface when fixed effects enter the probit specification. The advantage of estimating the outcome by OLS with fixed effects is that time-invariant outcomes are not dropped from the model and no incidental problem exists.

constrained to be between zero and 100 percent.²⁸ It should be noted that if one examines the predicted divorce probabilities by evaluating the minimum and maximum summary statistics of the household income volatility measures, the predicted outcomes are *not* below zero percent and do *not* exceed 100 percent.

2.5. RESULTS

The estimates shown in TABLE 2 provides the key contribution of this paper, which show the results from equations (2) and (4) for men and women in the full sample. Recall that the occupation indicators included in equation (3) provide a potential source of exogenous variation with which to identify the effects of household income volatility on divorce. TABLE A1 shows the estimates for the occupation indicators in equation (3), which suggest that the occupation indicators have significant predictive power in the first-stage models.²⁹ TABLE A2 shows the estimates for the occupation indicators when included in equation (4). The results suggest that the occupation indicators are unrelated to the divorce decision. Since the occupation indicators are highly correlated with the household income measures and uncorrelated with divorce, they provide a source of exogenous variation with which to identify the effects of household income volatility on

²⁸ When fixed effects enter the logit specification, individuals who have time-invariant outcomes are dropped. The result is a large reduction in the number of observations, which leads to insignificant results. The incidental parameters problem has the potential to surface when fixed effects enter the probit specification. The advantage of estimating the outcome by OLS with fixed effects is that time-invariant outcomes are not dropped from the model and no incidental problem exists.

²⁹ Note that the number of observations used in the first-stage regression differ from the number of observations used to estimate the divorce equations. The numbers of observations differ because I estimate equation (3) for all household income groups. For the divorce equations, I partition individuals into different household income groups: (i) the \$20,000 to \$200,000 household income range, (ii) the \$5,000 to \$40,000 household income range, and (iii) the \$40,000 to \$200,000 household income range.

divorce. TABLE 3 shows the results for the supplementary models from equation (2). TABLE 4 presents the estimates for the supplementary models from equation (4).

The models with fixed effects are the preferred estimates because of the importance of many time-invariant factors that have been shown to impact divorce decisions in the literature (e.g., religious upbringing, previous marriages, the presence of children from previous marriages, etc.).³⁰ These factors, as well as other unobservables, could be correlated with the household income measures, which could bias estimates. I present the OLS estimates along with the fixed effects estimates in the TABLES because it is illuminating to observe the changes in the estimates when fixed effects enter the models. In many cases, the coefficients and the statistical significance of the estimates change dramatically once fixed effects enter the models, which may suggest that time-invariant unobservables that are removed by including fixed effects are likely to be correlated with the household income measures.

The results shown in TABLE 2 for equation (2) indicate that increases in the level of household income has a substantial stabilizing effect on marriages for both men and women, which is consistent with previous findings (e.g., Becker et al. 1977; Hoffman and Duncan 1995; Burgess et al. 2003). Both positive and negative household income volatility are statistically significant and positive for men. For women, positive household income volatility is statistically significant and positive. The effect of negative household income volatility for women is not statistically different from zero

³⁰ For example, see Weiss and Willis (1997) and Charles and Stephens (2004). They highlight the importance of accounting for match quality when examining divorce behavior in response to earnings shocks.

when fixed effects enter the model; however, the effect is statistically significant in the OLS specification.

The bottom of TABLE 2 shows the results from equation (4). The estimates from equation (4) suggest that increases in the permanent income component stabilize marriages for men; however, the effect is not present for women. There is a statistically significant, positive increase in the risk of divorce due to increases in positive and negative household income volatility for men. However, women only experience a statistically significant, positive increase in the divorce risk from increases in negative household income volatility; the effect of positive household income volatility is not statistically different from zero.

TABLE 3 presents the results from equation (2) for lower- and higher-income individuals, respectively. The stabilizing effect associated with increases in the level of household income is consistent with the findings in TABLE 2, regardless of whether the individual has a low or high level of household income. For the lower-income group, men face a decrease in the divorce risk in response to positive household income volatility and an increase in the divorce risk in response to negative household income volatility. Positive household income volatility has a stabilizing effect on marriages for lower-income women; there is no evidence that negative household income shocks affects their divorce propensity. The results for the higher-income group differ from the lower-income group. Increases in positive and negative household income volatility increase the divorce risk for men. For women, positive household income shocks do not affect the divorce risk; however, negative household income shocks increase the risk of

divorce.

TABLE 4 shows the results from equation (4) for lower- and higher-household income individuals. TABLE 4 also shows the joint-exclusion statistic for the occupation indicators when they enter the divorce equations (i.e. equation (4)) for men and women in the two household income groups. The exclusion statistics indicate the occupation indicators are unrelated to the divorce decision. Therefore, they also provide a source of exogenous information for the lower- and higher-household income individuals. The results suggest that the permanent household income component is not statistically different from zero for women in both household income groups. However, the permanent income component has a stabilizing effect on divorce for men in the lower-income group. There is no evidence that the permanent income component affects the divorce propensity for higher-income men. For lower-household income men, neither positive nor negative household income volatility changes the divorce propensity. For women, positive household income shocks reduce the divorce risk and negative household income shocks raise the divorce risk. The results for the higher-income group differ from the lower-income group, which was also the case for the estimates from equation (2) for the two income groups. Both positive and negative household income volatility raise the divorce risk for men and women.

Because of the potential simultaneity bias associated with household income and the divorce decision in equation (2), it is difficult to make any conclusions from the estimates. As can be seen by comparing the estimates, the estimates of equation (2) differ—sometimes dramatically—from the estimates generated by the two-stage

procedure (i.e. equations (3) and (4)). The use of the exogenous information in the first-stage model appears to be the point of departure with the estimates. Thus, the results from equation (4) are the preferred estimates for the full sample and the sample used for the supplementary models.

The main results for men (i.e. the estimates for equation (4) shown in TABLE 2) confirm Becker et al.'s (1977) theoretical predictions and their empirical findings. My findings for women are not completely consistent with their theory or empirical results. For women, I find that positive household income volatility has no effect on the divorce risk and negative household income volatility increases the divorce risk. The former is not consistent with Becker et al.'s (1977) theory and empirical findings, which suggests that unexpected events or earnings shocks raise the divorce risk regardless of whether they are positive or negative. A potential explanation is that positive household income volatility could stem from an increase in husband's earnings, which may imply that the value of the outside option for the wife is less when her husband's income increases. However, positive shocks could raise the value of the outside option for men. These competing effects may imply that the directional impact on the incentives of spouses to divorce—attributable to positive household income volatility—counter each other.

The differing results found for lower- and higher-household income individuals suggest that the two groups respond differently to fluctuations in household income. However, both groups seem to be affected similarly by negative household income volatility, which raises the risk of divorce for both groups regardless of gender. This is not the case for men in the lower-household income group who appear to be unaffected

by household income volatility. The positive impact of negative household income volatility on divorce may suggest that reductions in the returns from marriage precipitate a rise in divorce. Women in the lower-household income group face a reduction in the divorce risk because of positive household income volatility. The reduction in the divorce risk for women could be due to the additional returns associated with the positive household income shock. The results for the higher-income group confirm Becker et al.'s (1977) predictions and findings for both men and women.

Since it is not possible to determine which spouse filed for divorce, it is difficult to determine precisely how household income volatility affects the divorce decisions of spouses. That is, husbands and wives could be affected differently by shocks to household income. For example, it could be that men receive more outside marriage offers when they experience positive income shocks. However, if the positive shock is derived from the wife, then it could be that the positive income shock generates a self-reliance effect for women. Both the former and the latter may translate into a higher risk of divorce.

2.6. CONCLUSIONS

This paper estimates the effects of positive and negative household income volatility on divorce for men and women using two different measures constructed from the NLSY79. There are two major issues that must be addressed when examining the effects of earnings shocks on divorce behavior: *(i)* exogenizing measures of earnings or income in the divorce equations and *(ii)* controlling for the quality of marriage match. I address

these issues by using exogenous variation in the occupations of individuals, which is similar to the approach used by Hess (2004), and by including individual-specific fixed effects. The first-stage models show that the occupation indicators have significant predictive power and the second-stage models show that the occupation indicators are unrelated to the divorce decision. This provides a source of exogenous information with which to identify the effects of household income volatility on divorce.

The results largely confirm the majority of previous findings, which indicate the importance of earnings and earnings shocks in divorce decisions. Analyzing the full sample yields results that suggest that men face an increased risk of divorce from increases in household income volatility, regardless of whether the household income shocks are positive or negative. The preferred estimates indicate that women face an increased risk of divorce when there is an increase negative household income volatility. No effect is found with respect to positive household income volatility for women. The results for women differ from the theory and findings of Becker et al. (1977).

The results for the lower- and higher-household income individuals differ. There is no statistical evidence that men in the lower-household income group are affected by household income volatility. However, men in the higher-household income group experience a rise in the divorce risk in response to increases in both positive and negative household income volatility. Increases in positive household income volatility have a stabilizing effect on marriages for women in the lower-household income group; however, increases in negative household income volatility raise the divorce risk. Both

positive and negative household income volatility raises the divorce risk for women in the higher-household income group.

The results found in this paper suggest that household income measures have significant effects on divorce behavior. The main results are largely consistent with the findings by Becker et al. (1977) and Hess (2004). My findings are not consistent with the interpretation offered by Charles and Stephens (2004), who contend that nonpecuniary factors may better explain divorce behavior.

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TABLE 1
COMPARISON OF HOUSEHOLD INCOME
VOLATILITY MEASURES FROM EQUATIONS 2 AND 4

Variable	Not Divorced		Divorced	
	Mean	Std. Dev.	Mean	Std. Dev.
<i>Equation 2:</i>				
<i>Men:</i>				
Positive Household Income Volatility (CV)	0.2371	0.2637	0.2517	0.2369
Negative Household Income Volatility ($D * CV$)	0.0520	0.1314	0.1088	0.1921
<i>Women:</i>				
Positive Household Income Volatility (CV)	0.2441	0.2619	0.2929	0.2725
Negative Household Income Volatility ($D * CV$)	0.0544	0.1496	0.1482	0.2238
<i>Equation 4:</i>				
<i>Men:</i>				
Positive Volatility Component (σ^ξ)	0.2159	0.2140	0.3406	0.3691
Negative Volatility Component ($D * \sigma^\xi$)	0.0552	0.1335	0.1313	0.2919
<i>Women:</i>				
Positive Volatility Component (σ^ξ)	0.2536	0.2550	0.3237	0.2708
Negative Volatility Component ($D * \sigma^\xi$)	0.0685	0.1737	0.1646	0.2283

Notes: For the household income volatility measures used in equation 2, the numbers of observations for men are 3,001 and there are 3,169 for women. As for the measures of household income volatility used for equation 4, the numbers of observations are 1,658 and 1,637 for men and women, respectively. The statistics above are computed by restricting the sample to those who do not divorce and those who do divorce over the sample period. These statistics provide a source of comparison between the levels of household income volatility those who do not divorce and those who do divorce experience.

TABLE 2
RESULTS FOR THE FULL SAMPLE

Variable	Men		Women	
	OLS	OLS with Fixed effects	OLS	OLS with Fixed Effects
<i>Equation 2:</i>				
Log of Real Household Income (HI)	-0.0820*** (0.013)	-0.1076*** (0.019)	-0.0957*** (0.012)	-0.1692*** (0.017)
Positive Household Income Volatility (CV)	0.0283 (0.019)	0.0659*** (0.026)	0.0532*** (0.020)	0.0486** (0.023)
Negative Household Income Volatility (D * CV)	0.1402*** (0.045)	0.0948** (0.044)	0.1198*** (0.041)	0.0225 (0.036)
R-squared	0.2632	0.2404	0.2459	0.2224
Number of Observations	3,001	3,001	3,169	3,169
<i>Equation 4:</i>				
Permanent Component (μ^l)	-0.0958*** (0.029)	-0.1341** (0.055)	-0.0423 (0.030)	-0.0644 (0.055)
Positive Volatility Component (σ^ξ)	0.0918** (0.042)	0.1125*** (0.039)	0.0058 (0.034)	-0.0557 (0.037)
Negative Volatility Component (D * σ^ξ)	0.1520** (0.063)	0.2001*** (0.053)	0.1384** (0.057)	0.1374*** (0.047)
R-squared	0.2597	0.0924	0.2515	0.0810
Number of Observations	1,658	1,658	1,637	1,637

Notes: standard errors are in parentheses. * indicates statistical significance at the ten percent level, ** indicates statistical significance at the five percent level, and *** indicates statistical significance at the one percent level. All models are estimated using OLS. Each model contains demographic and regional covariates. The models also include time indicators.

TABLE 3
RESULTS FROM EQUATION 2 FOR LOWER- AND
HIGHER-HOUSEHOLD INCOME GROUPS

Variable	Men		Women	
	OLS	OLS with Fixed Effects	OLS	OLS with Fixed Effects
<i>\$5,000 ≤ Household Income < \$40,000</i>				
Log of Real Household Income (HI)	-0.0540*** (0.016)	-0.0638*** (0.025)	-0.0832*** (0.015)	-0.1076*** (0.021)
Positive Household Income Volatility (CV)	-0.0730*** (0.025)	-0.0818** (0.040)	-0.0368 (0.027)	-0.0921** (0.036)
Negative Household Income Volatility (D * CV)	0.1091*** (0.041)	0.1094** (0.053)	0.1005*** (0.039)	0.0511 (0.045)
R-squared	0.3004	0.2084	0.3234	0.1394
Number of Observations	1,980	1,980	2,511	2,511
<i>\$40,000 ≤ Household Income ≤ \$200,000</i>				
Log of Real Household Income (HI)	-0.0624*** (0.022)	-0.0937*** (0.030)	-0.0553*** (0.019)	-0.1520*** (0.025)
Positive Household Income Volatility (CV)	0.0630** (0.027)	0.0768** (0.032)	0.0577** (0.025)	0.0308 (0.027)
Negative Household Income Volatility (D * CV)	0.1321** (0.064)	0.1108* (0.058)	0.1249** (0.059)	0.0907** (0.044)
R-squared	0.2451	0.2430	0.2077	0.2428
Number of Observations	1,720	1,720	1,744	1,744

Notes: standard errors are in parentheses. * indicates statistical significance at the ten percent level, ** indicates statistical significance at the five percent level, and *** indicates statistical significance at the one percent level. All models are estimated using OLS. Each model contains demographic and regional covariates. The models also include time indicators.

TABLE 4
RESULTS FROM EQUATION 4 FOR LOWER- AND
HIGHER-HOUSEHOLD INCOME GROUPS

Variable	Men		Women	
	OLS	OLS with Fixed Effects	OLS	OLS with Fixed Effects
<i>\$5,000 ≤ Household Income < \$40,000</i>				
Permanent Component (μ^l)	-0.0717* (0.044)	-0.2249** (0.096)	0.0504 (0.048)	0.0310 (0.103)
Positive Volatility Component (σ^ξ)	0.0663 (0.051)	0.0490 (0.072)	0.0381 (0.053)	-0.1725** (0.080)
Negative Volatility Component ($D * \sigma^\xi$)	0.2034*** (0.078)	0.1391 (0.092)	0.0552 (0.064)	0.1749** (0.083)
R-squared	0.3098	0.0733	0.3365	0.1099
Number of Observations	952	952	970	970
Joint-Exclusion Test	1.14 [0.331]	1.40 [0.186]	0.55 [0.835]	0.31 [0.972]
<i>\$40,000 ≤ Household Income ≤ \$200,000</i>				
Permanent Component (μ^l)	-0.0391 (0.038)	-0.0542 (0.064)	-0.0078 (0.032)	0.0255 (0.062)
Positive Volatility Component (σ^ξ)	0.1184** (0.055)	0.0834* (0.042)	0.0630 (0.040)	0.0771* (0.044)
Negative Volatility Component ($D * \sigma^\xi$)	0.0777 (0.073)	0.1541** (0.063)	0.0990 (0.094)	0.1400** (0.067)
R-squared	0.2349	0.1166	0.2025	0.0613
Number of Observations	995	995	985	985
Joint-Exclusion Test	1.32 [0.230]	0.52 [0.844]	1.09 [0.366]	0.87 [0.551]

Notes: standard errors are in parentheses. * indicates statistical significance at the ten percent level, ** indicates statistical significance at the five percent level, and *** indicates statistical significance at the one percent level. All models are estimated using OLS. Each model contains demographic and regional covariates. The models also include time indicators. The statistics under the *Exclusion Statistic* heading are the *F*-statistics and the corresponding p-values are in brackets. The exclusion statistic tests the occupation indicators in equation (4) to determine if the occupation variables are jointly excludable in the divorce equation.

APPENDIX

TABLE A1
ESTIMATES FOR THE OCCUPATION INDICATORS
FROM THE FIRST-STAGE MODEL
(EQUATION 3)

Variable	Men		Women	
Professional	0.3718***	(0.074)	0.0793	(0.083)
Manufacturing	0.3281***	(0.074)	0.0980	(0.083)
Sales	0.3564***	(0.082)	0.0233	(0.087)
Clerical	0.2105***	(0.076)	0.0105	(0.079)
Craftsman	0.2029***	(0.071)	0.0449	(0.093)
Operations	0.1559**	(0.071)	-0.1196	(0.083)
Laborer	0.0613	(0.075)	-0.1662*	(0.100)
Service	0.0645	(0.074)	-0.1784**	(0.080)
Private	-0.2984	(0.270)	-0.1815**	(0.104)
Joint-Exclusion Test	12.40***		11.43***	
	[0.000]		[0.000]	
Number of Observations	3,810		4,072	

Notes: Equation (3) is estimated by OLS and also includes demographic characteristics, labor-market characteristics, county-level covariates, and time indicators as control variables. Standard errors are in parentheses and p-values are in brackets. * indicates statistical significance at the ten percent level, ** indicates statistical significance at the five percent level, and *** indicates statistical significance at the one percent level.

TABLE A2
ESTIMATES FOR OCCUPATION INDICATORS
WHEN ADDED TO EQUATION 4
(FULL SAMPLE)

Variable	Men		Women	
	OLS	OLS with Fixed Effects	OLS	OLS with Fixed Effects
Professional	0.0088 (0.097)	-0.0496 (0.103)	0.0139 (0.118)	0.1276 (0.100)
Manufacturing	0.0026 (0.097)	-0.0763 (0.103)	0.0409 (0.118)	0.1277 (0.101)
Sales	-0.0117 (0.098)	-0.1016 (0.108)	-0.0252 (0.119)	0.0476 (0.105)
Clerical	0.0069 (0.098)	-0.0731 (0.105)	-0.0014 (0.116)	0.1214 (0.098)
Craftsman	0.0279 (0.096)	-0.0113 (0.096)	0.0347 (0.129)	0.1312 (0.099)
Operations	0.0012 (0.096)	-0.0266 (0.098)	-0.0299 (0.117)	0.0156 (0.100)
Laborer	0.0215 (0.098)	-0.0583 (0.102)	-0.0483 (0.126)	0.0187 (0.113)
Service	0.0047 (0.098)	-0.0821 (0.106)	-0.0149 (0.116)	0.0826 (0.103)
Private	-0.0195 (0.102)	-0.2328 (0.293)	-0.0372 (0.118)	0.0762 (0.135)
Joint-Exclusion Test	0.53 [0.856]	0.92 [0.509]	1.39 [0.187]	1.45 [0.162]
Number of Observations	1,658	1,658	1,637	1,637

Notes: All models include demographic characteristics, labor-market characteristics, county-level characteristics, and time indicators as control variables. Standard errors are in parentheses and p-values are in brackets. Note that all of the occupation indicators are not statistically significant from zero in all divorce equations. Likewise, joint-exclusion tests indicate that the occupation indicators are excludable in the divorce equations.

CHAPTER 3

INFLATION AND OTHER AGGREGATE DETERMINANTS OF THE TREND IN U.S. DIVORCE RATES SINCE THE 1960S

3.1. INTRODUCTION

Throughout the 1960s and 1970s, divorce rates in the United States (U.S.) increased dramatically. After peaking in the late 1970s, the number of new divorces declined throughout the 1980s and continues to decline today (FIGURE 1). Both the rise and fall of divorce rates has been a topic of much debate (Michael 1978; Johnson and Skinner 1986; Ruggles 1997a; Ruggles 1997b; Oppenheimer 1997; Preston 1997; Goldstein 1999).¹ However, evidence on aggregate determinants of divorce is sparse.² In particular, there appears to be no study on the effects of inflation on the number of new divorces. Analysing this relationship is the primary contribution of this study. However, in analyzing the determinants of divorce rates for the U.S. over the period 1955 to 2004 this

¹ A few of these studies examine the increase and leveling of divorce rates, which refers to the stock of divorces, not the number of new divorces. This paper focuses on new divorces.

² South (1985) and Bremmer and Kesselring (1999, 2004) are exceptions.

study will also revisit several other determinants of the U.S. divorce rate that have been discussed, sometimes rather controversially, in the literature.

Previous studies on the determinants of the trend in U.S. divorce rates focus on the impact of changes in divorce laws, female labour-force participation, and economic growth. For example, Friedberg (1998) and Gruber (2004) attribute a substantial portion of the rise in divorce rates in the late sixties to the adoption of no-fault or unilateral divorce laws. However, Wolfers (2006) shows that the rise in divorce rates induced by divorce reform is small and temporary. Empirical research investigating the relationship between female labour-force participation and divorce rates has not been conclusive because the relationship is complicated by the potential simultaneity of the two variables (e.g., see Bremmer and Kesselring (1999, 2004) and Spitze and South (1985, 1986)). By contrast, a well-established positive relationship appears to exist between divorce rates and economic growth (e.g., see Ogburn and Thomas 1922; Goode 1971; Norton and Glick 1979) although South (1985) finds a negative relationship.

I contend that inflation accounts for a considerable portion of the sharp rise in divorce rates in the U.S. throughout the 1960s and 1970s. Inflation worsens the terms of trade for households through the reduction of household consumption and leisure. Therefore, the returns to marriage should decline in response to an increase in the inflation rate. I also expect the effects of inflation on divorce to be persistent. Price instability may interfere with married couples' long-term financial plans, which could lead to an increase in divorce rates.

The present study uses a structural time-series (unobserved component) model to circumvent potential identification issues associated with the trend in the divorce rate.³ Harvey (1989, 1997) and Koopman et al. (2000) advocate this method when there is a clear trend in the data series. The estimation approach moves omitted or unobserved variables out of the error term and into a stochastic trend component so that consistent estimates of included right-hand-side regressors can be obtained. Structural time-series models are also advantageous because they allow for structural change through time-varying trend components.

I estimate three different model specifications for the divorce rate: (i) a smooth-trend model that considers only inflation and unemployment, (ii) a stochastic-trend model that also considers only inflation and unemployment, and (iii) a stochastic-trend model that includes inflation, unemployment, the growth rate of U.S. Gross Domestic Product (GDP), and changes in women's educational attainment. I conclude that increases in the inflation rate contributed to the rise in divorce rates during the 1960s and 1970s and that the stabilization of inflation in the mid-1980s through the 1990s accounts for a portion of the decrease in divorce rates over the same period. The impact of inflation is positive, persistent, and statistically significant in all specifications. Unemployment's effect depends on the specification of the trend and the inclusion of additional covariates. Economic growth and the rise in the economic independence of women, as proxied by their educational attainment, also appear to raise the divorce rate. In contrast to South

³ I use the terms structural time-series and unobserved component models interchangeably throughout this paper. The structural time-series methodology has been used to analyze a variety of different economic relationships (e.g., Abeyasinghe 2000; Muscatelli and Tirelli 2001; Scuffham 2003; Hon and Yong 2004; Dimitropoulos et al. 2005; Mazzocchi et al. 2006; Adhikari et al. 2007)

(1985), no inverse relationship can be verified to exist between the divorce rate and economic expansions. Instead, I provide support for previous empirical findings that indicate a positive relationship between economic growth and divorce rates.

This paper proceeds as follows. Section II describes the institutional background and the channels through which the explanatory variables are expected to affect divorce rates. Section III describes the data and the econometric methodology. Section IV presents results. Section V concludes.

3.2. INSTITUTIONAL BACKGROUND

Changes in the macroeconomy and demographics should affect the returns from marriage by altering consumption, leisure, and household specialisation.⁴ The same dynamics may also affect fertility and marriage-specific investments, which the literature shows to have binding effects on marriages.⁵ Becker et al. (1977) contend that surprises or unexpected events raise the risk of divorce because such changes alter the returns from marriage. Previous studies use earnings shocks to estimate the impact of unexpected events on divorce (e.g., Becker et al. 1977; Weiss and Willis 1997; Charles and Stephens 2004; Hess 2004).⁶ Variability in the inflation rate from the 1960s to the mid-1980s offers an alternative proxy for unexpected events. Aggregate measures of job availability and

⁴ The returns associated with marriage are usually attributed to the couple's ability to specialize in market and household work. For example, increases in consumption, leisure, and the production of one's own children have been cited as determinants of marriage.

⁵ See Becker et al. (1977).

⁶ The results in the majority of these studies support the theory and findings of Becker et al. (1977). Charles and Stephens (2004) find that job displacement, measured as layoffs, increases the risk of divorce. However, they find that disability and plant closings have no effect on divorce. Their results cast doubt on pecuniary motives of divorce, since disability, plant closings, and layoffs have similar long-run consequences.

economic growth could be other proxies for unexpected events, as both have seen perceptible fluctuations over time.

The U.S. experienced significant macroeconomic and demographic change over the last 50 years. Inflation rose in the 1960s and remained relatively unstable and at high levels until the early- to mid-1980s, when it began to stabilize. Inflation erodes the purchasing power of money, which can place significant stress on marriages by reducing consumption of market- and home-based goods and of leisure. Periods of rising inflation can cause married couples to specialize in market and household work sub-optimally. Inflationary periods imply that the price of consumption increases. As a result, spouses may have to adjust their labour supply to achieve pre-inflation consumption and leisure levels. If market work increases for both spouses, the returns to marriage are reduced because less time will be allocated to leisure and household production. It is possible for increases in wages to offset rising prices; however, Christiano et al. (2005) show that prices tend to adjust more freely than wages to a positive money supply shock. The differing responses of wages and prices to increases in the money supply imply that inflation should worsen the gains from household specialisation. Inflation can also have a long-run impact on divorce. Because rising prices can cause greater uncertainty in the future returns to marriage, couples may be unable to invest in marriage-specific capital. Low levels of investment in marriage-specific capital lower the opportunity cost of divorce, which makes divorce more likely.⁷

⁷ Marriage-specific capital could be the production of children, investments in joint assets, and investing in additions to human capital for spouses.

The erratic behaviour of inflation from the 1960s to the mid-1980s (FIGURE 2) was roughly concomitant with fluctuations in unemployment. Unemployment began to behave erratically in the 1970s and continued through the early- to mid-1980s. Since the early- to mid-1980s, unemployment has remained relatively stable. The rise and fall of divorce appears to have been largely concurrent with the dynamics of inflation and unemployment (FIGURES 1 and 2).

Compared to inflation, the channels through which unemployment affects divorce are less clear. On the one hand, divorce may increase because higher unemployment reduces consumption of market- and home-based goods and of leisure. Consumption and leisure should decrease because layoffs occur and economic theory predicts that job seekers accept lower wages. On the other hand, it could be that the value of the outside option, which is divorce, is lower when unemployment is higher. If one spouse is considering divorce, high unemployment may stabilize marriages because of less job availability and lower wage offers. It could also be that unemployment insurance provides a means of consumption insurance, which may have binding effects on marriages.

The U.S. also experienced perceptible fluctuations in the growth rate of U.S. GDP over the same period as the rise in divorce. The upper portion of FIGURE 3 suggests that the growth rate of U.S. GDP experienced greater growth volatility from 1955 to 1980 compared with growth volatility since the 1980s, which is roughly concurrent with both the rise and fall of divorce.

South (1985) examines the role of expansions and recessions on divorce behaviour and finds that divorces increase in recessions and decrease during expansions. South uses

three proxies for economic growth: the unemployment rate, the change in Gross National Product (GNP), and the percentage change in GNP. South examines each of these variables independently from each other. Examining changes in GNP along with the unemployment rate should not introduce bias to the estimates. However, omitting one of these variables could bias estimates, as changes in GNP and the unemployment rate may be correlated.⁸ South's results may suggest that recessionary periods cause stress within marriages and expansionary periods create additional returns to marriage.

Economic growth could also have a positive effect on divorce rates. It could be that recessionary periods bind marriages because two incomes may be necessary to offset the adverse effects of the economic downturn. Expansionary periods may induce individuals to become more self-reliant. That is, economic expansions may allow individuals to earn more and to become more independent, which could increase divorce rates. In fact, most studies find a positive relationship between economic expansions and divorce rates (e.g., Ogburn and Thomas (1922), Goode (1971), and Norton and Glick (1979)).

A significant demographic transformation in the U.S. was the steady increase in women's educational attainment.⁹ Using women's educational attainment as a predictor of the trend in divorce rates, instead of their labour-force participation rate, provides another way to examine the effect of increases in the economic power of women on divorce behaviour. A number of studies analyze the effects of female labour-force

⁸ In fact, a simple OLS regression of the growth rate of gross domestic product on the unemployment rate, and vice versa, yields a negative, contemporaneous relationship between the two variables. As a result, South's estimates of the change (or percentage change) in GNP could be downwardly biased, which could be the reason for the negative effect found with respect to changes in GNP.

⁹ See the lower portion of FIGURE 3.

participation on divorce behaviour. However, estimating the effect of female labour-force participation on divorce is complicated by the potential simultaneity bias between the two variables. A comparison of the findings of Green and Quester (1982), Shapiro and Shaw (1983), Johnson and Skinner (1986), Bremmer and Kesselring (1999, 2004), and Lombardo (1999) with those of Spitze and South (1985, 1986) and Mincer (1985) suggest that the two variables may be simultaneously determined.¹⁰ The former group of studies concludes that divorce increases women's labour-force participation while the latter suggests the opposite. To circumvent identification issues associated with female labour-force participation, I use women's educational attainment as a proxy for the women's liberation movement that occurred in the 1960s and 1970s.¹¹ The rationale behind this choice is the fact that increases in the educational attainment of women create options for a single life that are independent of a current job.

Goldin and Katz (2000) contend that affordable contraceptives gave women greater control of fertility decisions and reduced the opportunity costs associated with investments in human capital. Increases in human capital improved the prospects of women for high-wage employment, which gave them greater bargaining power within

¹⁰ Bremmer and Kesselring (1999, 2004) are the only studies in this list to examine the aggregate relationship between divorce rates and female-labour force participation. Bremmer and Kesselring (1999) attempt to determine the causal direction of the two variables and find that the divorce rate 'Granger' causes female labour-force participation. Bremmer and Kesselring (2004) examine the long-run relationship between divorce rates and female labour-force participation. Their results suggest that rising divorce rates increases female labour-force participation and that rising female-labour force participation increases divorce rates.

¹¹ A related issue is the role of female earnings in divorce decisions. Ressler and Waters (2000) and Kesselring and Bremmer (2006) examine the relationship between female earnings and divorce. Ressler and Waters (2000) find evidence that the two variables are jointly determined. Kesselring and Bremmer (2006) find evidence confirming the results found by previous research; that is, the rising economic power of women increases the risk of divorce. After comparing these findings, it is difficult to determine the causal relationship between the two variables.).

households (Costa 2000). Achieving greater bargaining power and independence in the labour market could increase divorce rates because women could become more self-reliant.

3.3. DATA AND ECONOMETRIC STRATEGY

Data on the divorce rate come from the *Historical Statistics of the United States: Millennium Edition* and U.S. Statistical Abstracts and span the time period from 1955 to 2004. The measure for the divorce rate is the number of new divorces each year per 1,000 persons. TABLE 1 displays the variable definitions. Data on the inflation rate and the unemployment rate are taken from the Bureau of Labor Statistics (BLS). The measure of women's education attainment is derived from the higher education statistics of the U.S. Census Bureau by using the percentage of women enrolled in higher education relative to the total population enrolled. The measure of economic growth is the growth rate of U.S. Gross Domestic Product (GDP), which is calculated by the St. Louis Federal Reserve Board. TABLE 2 presents summary statistics and provides data sources for the variables considered. Note that the variable *weduc* is scaled to be made comparable to the other explanatory variables.

Tests for stationarity are shown in TABLE 3. They suggest that the variables *inflation*, *unemp*, and *growth* are stationary. However, the variable *weduc* is non-stationary and enters the model in first-differenced form. Since the divorce rate follows a trend according to TABLE 3 and FIGURE 1, it is necessary to include a trend in the empirical model to avoid spurious results (Harvey 1989, 1997). Harvey (1997) contends that

deterministic-trend models are, in many cases, too restrictive. The unobserved component modeling strategy does not rely on unit root tests to dictate the specification of the trend.¹² The initial specification of the trend includes stochastic level and slope components. The flexibility of the modeling strategy allows me to test the level and slope components to determine if another simpler specification of the trend is more appropriate.

The inclusion of a stochastic trend permits omitted factors to be moved out of the error term. Capturing theoretically relevant variables in a stochastic trend allows for the estimates to be unbiased assuming there is no simultaneity bias between the outcome variable and the right-hand-side variables. Unobserved component models also allow for structural change through time-varying level and slope components. Most other time-series models are sensitive to structural change and omitted variables (e.g., cointegration techniques and distributed-lag models).

The general form of the structural time series model is

$$y_t = \mu_t + \sum_i \sum_j \alpha_{ij} x_{i,t-j} + \varepsilon_t \quad \text{for } t=1,2,\dots,T. \quad (1)$$

The dependent variable is y_t ; μ_t is a time-varying intercept term; $x_{i,t-j}$ is the regressor variable i subject to time lag j ; α_{ij} represents the coefficient associated with the variable $x_{i,t-j}$; and ε_t is a zero mean constant variance disturbance term. The term μ_t enables the researcher to capture unobservables and omitted variables that influence the dependent

¹² Since unit root tests rely on autoregressive models, Harvey (1997) contends that such tests may exhibit poor statistical properties. In fact, Harvey and Jaeger (1993) show with simulations that unit root tests do not typically detect variables that are $I(2)$. Detecting a unit root process usually results in the researcher concluding that the series is $I(1)$.

variable, which may be correlated with the variables in $x_{i,t-j}$. The μ_t process takes the form:

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t \quad \eta \sim NID(0, \sigma_\eta^2) \quad (2)$$

$$\beta_t = \beta_{t-1} + \xi_t \quad \xi \sim NID(0, \sigma_\xi^2). \quad (3)$$

The term μ_t can be interpreted as the “level component” of a stochastic trend and β_t represents the drift parameter, which is the “slope” of the level component. The level component follows a random walk with drift and the slope component follows a random walk. The terms η_t and ξ_t are white noise disturbances. The white noise disturbances, η_t and ξ_t , are independent of each other and of ε_t . A Kalman filter recovers the state vectors μ_t and β_t .¹³ Equations (1) through (3) are in their most general form. The model can be tested down to contain a fixed level, a fixed slope, or other specifications including a fixed level and no slope, which is equivalent to ordinary least squares (OLS).¹⁴

3.4. RESULTS

I estimate three different models. Two of the models use only inflation and unemployment as explanatory variables. The third and final model considers inflation, unemployment, the growth rate of U.S. GDP, and the change in women’s educational attainment. There are three reasons for estimating three different model specifications: (i) to resolve the mixed results found for unemployment in the first two models, (ii) to

¹³ See Harvey (1989) for a detailed description of the Kalman filter and its application in structural time-series models. The statistical package used—Structural Time-Series Analyser, Modeller, and Predictor (STAMP)—offers a canned procedure for the Kalman Filter.

¹⁴ If the variance of the disturbance term η_t equals zero and the variance of the disturbance term ξ_t is nonzero, the model takes the smooth-trend specification, which is integrated of order two (Harvey 1997).

attempt to explain a greater portion of the trend in the data for the divorce rate, and *(iii)* to check the validity of the robust, positive, and persistent effect of inflation on the divorce rate.

3.4.1. Results from Models with only Inflation and Unemployment

This section presents two of the three unobserved component models, which use only inflation and unemployment as explanatory variables: *(i)* the smooth-trend model and *(ii)* the stochastic-trend model. The reason for the two trend specifications relates to different ways that I follow the general-to-specific methodology. The results suggest that the ways in which the methodology is carried out has a significant impact on the parameter estimates for unemployment, especially its long-run effect.

I begin with a stochastic level and slope specification with two lags of all variables including the dependent variable. The general specification applies to equations (1) through (3). The estimates from the general specification indicate that the variance of the disturbance term in equation (2) equals zero, which suggests that the trend should contain a fixed level; however, the slope remains stochastic. When the level is fixed and the slope is stochastic, the trend is smooth. This implies that—conditional on the included explanatory variables—the rate of new divorces is integrated of order two. I restrict the model to contain a smooth trend throughout successive parameter restrictions. After

restricting the level component to be fixed and the slope to be stochastic, I test the model down to a more parsimonious form.¹⁵

The second model reverts back to the general, stochastic specification each time a parameter restriction is made. I estimate the models with the stochastic specification to determine if restricting the model to contain a smooth trend throughout successive parameter restrictions is appropriate. After making a parameter restriction and reestimating the model with a stochastic level and slope, the estimated variances of the disturbance terms in equations (2) and (3) indicate that the stochastic-trend specification is appropriate. However, there is only one parameter restriction because all explanatory variables are at least marginally statistically significant different from zero after the first parameter restriction is made.

TABLE 4 shows the results from the smooth-trend model and TABLE 5 provides the results from the stochastic-trend model. For both models, I check for non-normality of residuals, higher-order autocorrelation and heteroskedasticity in the residuals, and the model's out-of-sample forecasting properties. I rely on the model's out-of-sample forecasting properties to validate any further parameter restrictions. The estimates for the smooth-trend and stochastic-trend models do not indicate any statistical adequacy problems, as evidenced by the battery of statistical adequacy tests shown at the bottom of TABLES 4 and 5 and the residual graphics provided in FIGURES 4 and 5.

¹⁵ I adopt the empirical methodology advocated by the London School of Economics (LSE). Each set of parameter restrictions are validated by checking the statistical properties of the model. The LSE approach assumes that all models are false. The goal of the LSE approach is to find an adequate model; one that captures the data generating process.

The remaining level and slope components from the two specifications are shown in FIGURES 6 and 7. The fact that neither the level nor slope components are flat but show distinctive patterns suggests that the included explanatory variables do not fully capture the data generating process. However, because unobservables or omitted variables can be isolated and that the estimates are not sensitive to structural change, the empirical approach allows for the effects of inflation and unemployment to be identified.

Consistent with my hypotheses, inflation is statistically significant, positive, and persistent in both specifications. As shown by the larger estimated coefficients at lagged values, the adverse impact of inflation seems to take more time to affect divorce rates. Regardless of the trend specification, unemployment has a contemporaneous, statistically significant, negative effect on divorce. The negative impact of unemployment is opposite to the findings of South (1985), who finds a positive effect. The smooth-trend specification does not indicate any persistent effects with respect to unemployment. However, when the model takes the stochastic trend specification (i.e. TABLE 5), unemployment's long-run effect is positive and substantial. The contemporaneous, negative effect found for unemployment may be due to the value of divorce being lower when unemployment is higher because obtaining a job would be more difficult and wage offers would be lower.

The long-run effects of inflation and unemployment on divorce are shown in TABLE 6. The long-run effects indicate that inflation has a considerable effect on divorce, regardless of the trend specification; however, the effects are larger in the stochastic-trend model. A doubling of the inflation rate from its mean value increases the number

of new divorces per 1,000 persons by 0.17 and 0.29 in the smooth- and stochastic-trend models, respectively. There are wide discrepancies with respect to the long-run effects of unemployment, as evidenced by the negative effect in the smooth-trend model and the positive effect in the stochastic-trend model. For the smooth-trend and stochastic-trend specifications, a doubling of the unemployment rate from its mean value decreases the number of new divorce per 1,000 persons by 0.37 and increases the number of new divorces per 1,000 persons by 0.25 in the long run, respectively.

The results for the stochastic-trend model seem more plausible. Persistent unemployment is likely to generate greater marital instability because jobs are scarce and wage offers become lower over time. Lower job availability and lower wage offers would reduce consumption of market- and home-based goods and of leisure both today and in the future. As a result, the long-run gains from household specialisation are reduced when there is persistent unemployment. A comparison of FIGURES 6 and 7 provides further support for the stochastic-trend model, which indicates that it accounts for a larger portion of the trend in the divorce rate compared with the smooth-trend model.

Although the results presented in this section do not indicate statistical problems, two issues remain unaddressed: (i) a large portion of the trend in divorce rates is not explained by the included explanatory variables and (ii) the results found with respect to unemployment are conflicting. I attempt to address these issues in the next section by including measures of economic growth and changes in women's educational attainment. Using a measure of economic growth provides a measure of the health of the economy,

which may help resolve the differing long-run effects associated with unemployment in the first two models. Changes in women's educational attainment offer a proxy for the women's liberation movement that occurred over the same period as the rise in divorce. The inclusion of these covariates should account for a larger portion of the trend in divorce rates and may aid in resolving discrepancies found with respect to unemployment. The final model, including additional covariates, also provides a way to check the robustness of inflation's persistent effect on divorce.

3.4.2. Results from Model with Additional Explanatory Variables

As in the first two specifications, I begin with a stochastic level and slope specification with two lags of the dependent variable and all explanatory variables except the change in women's educational attainment, which I only include one lag because it is differenced to be made stationary. Following the estimation of the general specification, I test the model down to a more parsimonious form. The estimated variances of the disturbance terms in equations (2) and (3) suggest that the trend should take the stochastic specification. The variances of the disturbance terms are also nonzero through successive parameter restrictions; thus, all of the models take the form of equations (1) through (3).

TABLE 7 shows the results from the final model, which considers all explanatory variables. In the final model, I also check the statistical adequacy of the model and follow the same methodological approach as outlined above. The statistical adequacy measures for the final model do not indicate any problems, as shown in TABLE 7 and FIGURE 8. The remaining trend components for the final model are given in FIGURE 9.

As was the case for the first two models, the included explanatory variables do not fully explain the trend in the divorce rate. However, adding other covariates to the basic specification does account for a larger portion of the trend in divorce rates. The long-run effects for the final model are provided in TABLE 8. The long-run effects from the other two models are also included in TABLE 8 in order to compare the long-run effects across different models. Note that the magnitude of inflation's long-run effect is similar in all models, especially the stochastic-trend models.

Consistent with my hypotheses, inflation remains statistically significant, positive, and persistent when additional regressors enter the model. In the long run, a doubling of the inflation rate increases the number of new divorces per 1,000 persons by 0.30. The change in women's educational attainment and economic growth are statistically significant, positive, and persistent. If the change in the ratio of women in higher education relative to the total population in higher education increases by ten percentage points, the magnitude of the long-run impact on the number of new divorces per 1,000 is an increase of 0.07. The results for economic growth are opposite to the findings of South (1985), who contends that the divorce rate rises in recessions and falls in expansions. However, my findings are consistent with the work of Ogburn and Thomas (1922), Goode (1971), and Norton and Glick (1979). Comparing the long-run effect of the growth rate of GDP at its mean value with a doubling of its mean value indicates a rise in the divorce rate of 0.22 divorces per 1,000 persons. The reversal of the sign associated with the coefficient for unemployment in the final model could be due to the inclusion of the growth rate of U.S. GDP, as the two variables measure similar aspects of

the macroeconomy. Unemployment has a statistically significant, positive effect, which is consistent with South's (1985) findings. A comparison of the mean value of the unemployment rate with a doubling of its means value suggests that the number of new divorce per 1,000 persons increases by 0.35.

The final model confirms the robustness of inflation's effect on divorce, explains a larger portion of the trend in the divorce rate, and aids in resolving conflicting estimates found for the effect of unemployment on divorce. The robust, positive effect of inflation on the divorce rate may be due to the additional strains placed on marriages through decreases in purchasing power, which may affect consumption, household specialisation, and investments in marriage-specific capital. The positive effect associated with unemployment is in line with Becker et al.'s (1977) theory, which suggests that increases in unemployment would reduce the returns to marriage by altering consumption, leisure, and household specialisation decisions; therefore, divorce should be more likely when there is higher unemployment.

The change in women's educational attainment also appears to explain a portion of the rise in divorce over the sample period. This suggests that the addition to human capital may have given women greater independence and bargaining power within households. Additions to the human capital enable women to compete effectively in the service-based economy because service-oriented work requires larger additions to human capital.

The persistent and positive effect on divorce rates found for economic growth suggests a channel through which increases in economic opportunities affect divorce.

Since economic growth implies greater job availability, higher wage offers, and higher returns on investment, divorcees have the potential to earn more and higher returns during expansionary periods. Thus, economic growth could induce a rise in divorce because the value of becoming divorced may be higher, as there is greater job availability and higher earnings potential.

3.5. CONCLUDING REMARKS

This paper adds to the empirical research on the determinants of divorce by examining the impact of inflation. I construct three unobserved component models for the divorce rate using annual data for the U.S. from 1955 to 2004. Two of the three specifications consider the effects of only inflation and unemployment on divorce. The other model includes inflation, unemployment, changes in women's educational attainment, and the growth rate of U.S. GDP as predictors of the divorce rate.

The empirical methodology circumvents potential identification problems because I model the trend in the divorce rate as an unobserved variable. The inclusion of an unobserved component allows for unobservables and omitted variables to be moved out of the error term into a stochastic trend component. This allows for the model's parameters to be estimated consistently. The empirical approach does not impose restrictive assumptions on the trend in the dependent variable but allows the data to generate the appropriate model specification.

The effects of inflation are statistically significant, positive, and persistent regardless of the trend specification and inclusion of additional explanatory variables. The long-run

effects of inflation are substantial. I also offer support for previous research that finds a positive relationship between economic growth and divorce rates. This result differs from earlier research by South (1985) who finds a negative relationship. Previous research on the link between female labour-force participation and divorce has suggested that the two variables are simultaneously determined. I use changes in women's educational attainment as a proxy for their rising economic power and show that it has a positive effect on divorce rates.

I conclude that inflation, economic growth, and changes in women's educational attainment account for a substantial portion of the trend in U.S. divorce rates. Because the unobserved trend components remain significant in all models, one has to conclude that the included explanatory variables do not fully explain the rise and fall in divorce rates. This suggests that some further research is warranted to explain the rise and fall of the divorce rate since the 1960s.

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TABLE 1
VARIABLE NAMES AND VARIABLE DEFINITIONS

Variable	Variable Definition
<i>divorce</i>	Number of new divorces per 1,000 persons
<i>inflation</i>	Log of the ratio of the Consumer Price Index (CPI) at period t relative to the CPI at period $t-1$
<i>unemp</i>	Percentage of the workforce that is unemployed but is actively pursuing employment
<i>growth</i>	Log of the ratio of U.S. Gross Domestic Product (GDP) at period t relative to U.S. GDP at $t-1$.
<i>weduc</i>	Percentage of women enrolled in higher education relative to the total population enrolled in higher education

Notes: All data relate to the United States and cover the period 1955 to 2004.

TABLE 2
SUMMARY STATISTICS AND VARIABLE SOURCES

Variable	Mean	Std. Deviation	Minimum	Maximum
<i>divorce</i>	4.0933	0.9708	2.20	5.30
<i>inflation</i>	4.2901	3.0475	0.67	13.26
<i>unemp</i>	5.9183	1.4415	3.49	9.71
<i>growth</i>	3.3700	2.1938	-1.90	7.20
<i>weduc</i>	4.9401	0.7238	3.54	5.89

Notes: All data relate to the United States. The data span the years 1955 to 2004 (obs. = 49). Data for the divorce rate come from the *Historical Statistics of the United States: Millennium Edition* and U.S. statistical abstracts. Data for inflation, unemployment, and the growth rate of GDP are accessed through www.economagic.com. Data for women's educational attainment come from the U.S. Census Bureau and are accessible at <http://www.census.gov/population/www/socdemo/school.html>. The variable *weduc* is scaled to be made comparable to the other explanatory variables.

TABLE 3
TESTS FOR STATIONARITY

Variable	KPSS Test	
	Trend {H0 = I(0)}	No-trend {H0 = I(0)}
<i>divorce</i>	0.7892**	0.2550*
<i>inflation</i>	0.2103	0.2099
<i>unemp</i>	0.1584	0.2060
<i>growth</i>	0.0487	0.1734
<i>weduc</i>	0.2218*	1.0295**

Notes: * indicates statistical significance at the five percent level and ** indicates statistical significance at the one percent level. Details of the KPSS test are outlined in Kwiatkowski et al. (1992). The KPSS uses stationarity as the null and tests against the alternative hypothesis of a unit root.

TABLE 4
EMPIRICAL RESULTS FOR THE DIVORCE RATE
(SMOOTH TREND)

Variable	Model 1		Model 2		Model 3		Model 4		Model 5	
	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)
μ	5.625	0.000	4.842	0.000	5.156	0.000	5.311	0.000	5.240	0.000
β_t (last year)	-0.264	0.015	-0.247	0.021	-0.218	0.032	-0.210	0.041	-0.166	0.090
$divorce_{t-1}$	-0.474	0.002	-0.398	0.002	-0.393	0.003	-0.409	0.002	-0.336	0.011
$divorce_{t-2}$	-0.135	0.336								
$inflation_t$	0.015	0.072	0.015	0.086	0.009	0.213				
$inflation_{t-1}$	0.027	0.004	0.028	0.003	0.020	0.004	0.020	0.005	0.015	0.018
$inflation_{t-2}$	0.036	0.000	0.036	0.000	0.033	0.000	0.031	0.000	0.024	0.000
$unemp_t$	-0.036	0.088	-0.043	0.032	-0.050	0.011	-0.060	0.001	-0.062	0.000
$unemp_{t-1}$	0.029	0.200	0.028	0.219						
$unemp_{t-2}$	0.042	0.041	0.044	0.031	0.032	0.085	0.030	0.112		
Statistical Adequacy Measures:										
R^2	0.9907		0.9904		0.9899		0.9895		0.9888	
AIC	4.3883		4.4065		4.4088		4.4115		4.3894	
SIC	3.9378		3.9969		4.0402		4.0838		4.1027	
Het. $F(13,13)$	0.9807		1.0934		1.0730		0.9846		1.0577	
Cusum (6)	-0.6447		-0.6253		-0.4738		-0.3973		-0.1971	
Cusum (10)	-0.4289		-0.3672		-0.2638		-0.1966		-0.1576	
<i>p-values:</i>										
Normality (2)	0.5852		0.2718		0.3421		0.1271		0.4295	
Box-Ljung (6)	0.3758		0.3914		0.5088		0.4028		0.7883	
Forecast (6)	0.9730		0.9701		0.9436		0.9453		0.9836	
Forecast (10)	0.9895		0.9875		0.9726		0.9879		0.9851	

Notes: There are 44 observations for each of the models. Columns (a) and (b) represent the coefficient estimates and the corresponding p-values, respectively. AIC represents the Akaike Information Criterion developed by Akaike (1974). SIC is the Schwarz Information Criterion. The SIC is sometimes referred to the Bayesian Information Criterion (BIC). Het. is an F -test for Heteroskedasticity. The critical value for the Heteroskedasticity test is 2.58. The Doornik and Hansen (1994) tests for normality; it has normality as the null hypothesis. The test Box-Ljung represents the Ljung and Box (1978) test for higher-order autocorrelation. The test Forecast (h) is a one-step-ahead χ^2 predictive test h observations into the future. Cusum (h) is a one-step-ahead predictive t -test h observations into the future for the residuals.

TABLE 5
EMPIRICAL RESULTS FOR THE DIVORCE RATE
(STOCHASTIC TREND)

Variable	Model 1		Model 2	
	(a)	(b)	(a)	(b)
μ	5.625	0.000		
μ_t (last year)			4.320	0.000
β_t (last year)	-0.264	0.015	-0.206	0.025
$divorce_{t-1}$	-0.474	0.002	-0.279	0.043
$divorce_{t-2}$	-0.135	0.336		
$inflation_t$	0.015	0.072	0.017	0.052
$inflation_{t-1}$	0.027	0.004	0.028	0.002
$inflation_{t-2}$	0.036	0.000	0.033	0.000
$unemp_t$	-0.036	0.088	-0.038	0.061
$unemp_{t-1}$	0.029	0.200	0.037	0.099
$unemp_{t-2}$	0.042	0.041	0.042	0.042
Statistical Adequacy Measures:				
R^2	0.9907		0.9905	
AIC	4.3883		4.3719	
SIC	3.9378		3.9214	
Het. $F(13,13)$	0.9807		1.0865	
Cusum $t(6)$	-0.6447		-0.8073	
Cusum $t(10)$	-0.4289		-0.5336	
<i>p-values:</i>				
Normality (2)	0.5852		0.3095	
Box-Ljung (6)	0.3758		0.2868	
Forecast (6)	0.9730		0.9504	
Forecast (10)	0.9895		0.9794	

Notes: There are 44 observations for each of the models. Columns (a) and (b) represent the coefficient estimates and the corresponding p-values, respectively. AIC represents the Akaike Information Criterion developed by Akaike (1974). SIC is the Schwarz Information Criterion. The SIC is sometimes referred to the Bayesian Information Criterion (BIC). Het. is an F -test for Heteroskedasticity. The critical value for the Heteroskedasticity test is 2.58. The Doornik and Hansen (1994) tests for normality; it has normality as the null hypothesis. The test Box-Ljung represents the Ljung and Box (1978) test for higher-order autocorrelation. The test Forecast (h) is a one-step-ahead χ^2 predictive test h observations into the future. Cusum (h) is a one-step-ahead predictive t -test h observations into the future for the residuals.

TABLE 6
LONG-RUN EFFECTS FOR
VARIOUS TREND SPECIFICATIONS

Variable	Smooth Trend	Stochastic Trend
<i>inflation</i>	0.039	0.068
<i>unemp</i>	-0.062	0.041

Notes: Long-run multipliers are calculated by dropping the time subscripts in the final models and solving for the dependent variable. Note that the long-run multiplier for *unemp* in the smooth-trend specification equals the impact multiplier.

TABLE 7
EMPIRICAL RESULTS FOR THE DIVORCE RATE
(WITH ADDITIONAL REGRESSORS)

Variable	Model 1		Model 2		Model 3		Model 4	
	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)
μ_t (last year)	2.406	0.005	2.622	0.002	3.016	0.000	3.026	0.000
β_t (last year)	-0.101	0.142	-0.113	0.088	-0.129	0.049	-0.129	0.051
<i>divorce</i> _{<i>t-1</i>}	0.097	0.568	0.051	0.743				
<i>divorce</i> _{<i>t-2</i>}	0.073	0.656	0.053	0.722				
<i>inflation</i> _{<i>t</i>}	0.023	0.026	0.023	0.016	0.023	0.013	0.023	0.011
<i>inflation</i> _{<i>t-1</i>}	0.020	0.049	0.021	0.019	0.021	0.012	0.021	0.010
<i>inflation</i> _{<i>t-2</i>}	0.021	0.051	0.022	0.017	0.023	0.008	0.023	0.007
<i>unemp</i> _{<i>t</i>}	0.059	0.183	0.059	0.073	0.060	0.059	0.058	0.064
<i>unemp</i> _{<i>t-1</i>}	-0.007	0.875						
<i>unemp</i> _{<i>t-2</i>}	0.002	0.961						
<i>growth</i> _{<i>t</i>}	0.036	0.027	0.035	0.002	0.035	0.002	0.035	0.001
<i>growth</i> _{<i>t-1</i>}	0.021	0.275	0.023	0.031	0.023	0.019	0.023	0.019
<i>growth</i> _{<i>t-2</i>}	0.007	0.396	0.009	0.163	0.009	0.120	0.010	0.094
Δ <i>weduc</i> _{<i>t</i>}	0.026	0.541	0.022	0.564	0.018	0.612		
Δ <i>weduc</i> _{<i>t-1</i>}	0.095	0.029	0.090	0.020	0.086	0.017	0.074	0.007
Statistical Adequacy Measures:								
R ²	0.9928		0.9928		0.9928		0.9928	
AIC	4.2110		4.3028		4.3940		4.4313	
SIC	5.5352		3.7065		3.8772		3.9543	
Het. <i>F</i> (14,14)	0.5163		0.5234		0.5332		0.5499	
Cusum (6)	-0.6634		-0.5919		-0.6076		-0.5957	
Cusum (10)	-0.8333		-0.4831		-0.5369		-0.5347	
<i>p-values:</i>								
Normality (2)	0.9040		0.8953		0.9761		0.9962	
Box-Ljung (6)	0.8240		0.9012		0.9008		0.8230	
Forecast (6)	0.9098		0.9119		0.8964		0.8788	
Forecast (10)	0.9924		0.9928		0.9882		0.9829	

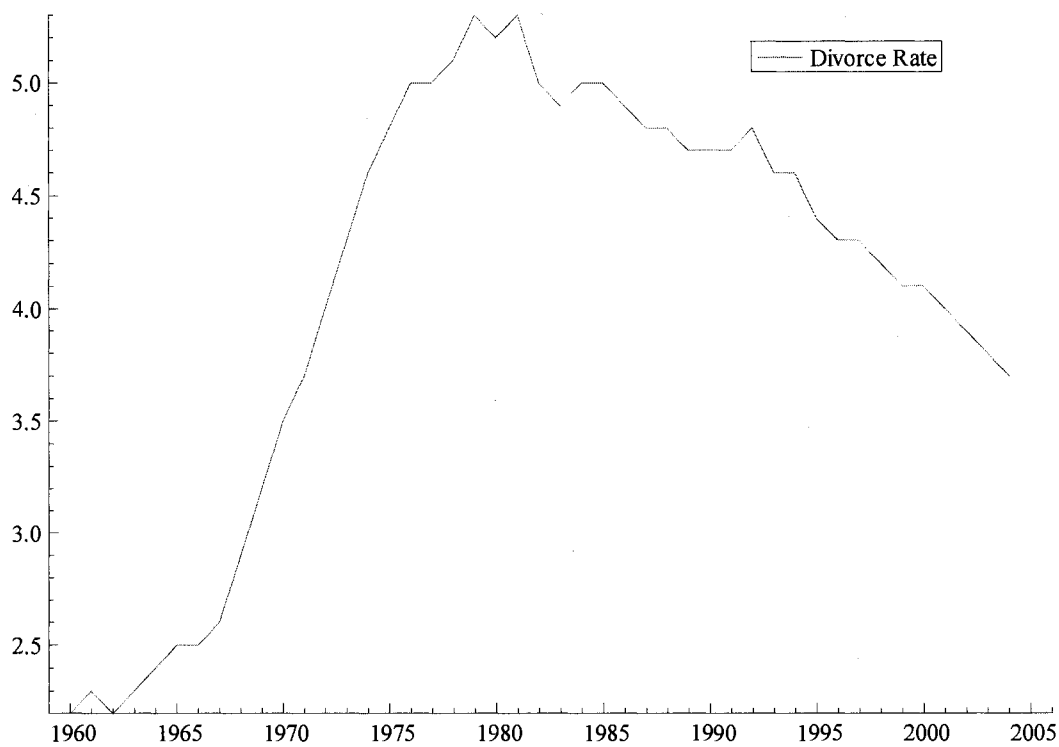
Notes: There are 44 observations for each model. Columns (a) and (b) represent the coefficient estimates and the corresponding p-values, respectively. AIC represents the Akaike Information Criterion. SIC is the Schwarz Information Criterion. Het. is a test for Heteroskedasticity (critical value: 2.48). The Doornik and Hansen (1994) test checks for non-normality. The test Box-Ljung is Ljung and Box's (1978) test for higher-order autocorrelation. The test Forecast (*h*) is a one-step-ahead χ^2 predictive test *h* observations into the future. Cusum (*h*) is a one-step-ahead predictive *t*-test *h* observations into the future for the residuals.

TABLE 8
LONG-RUN EFFECTS OF THE EXPLANATORY VARIABLES
ON THE DIVORCE RATE

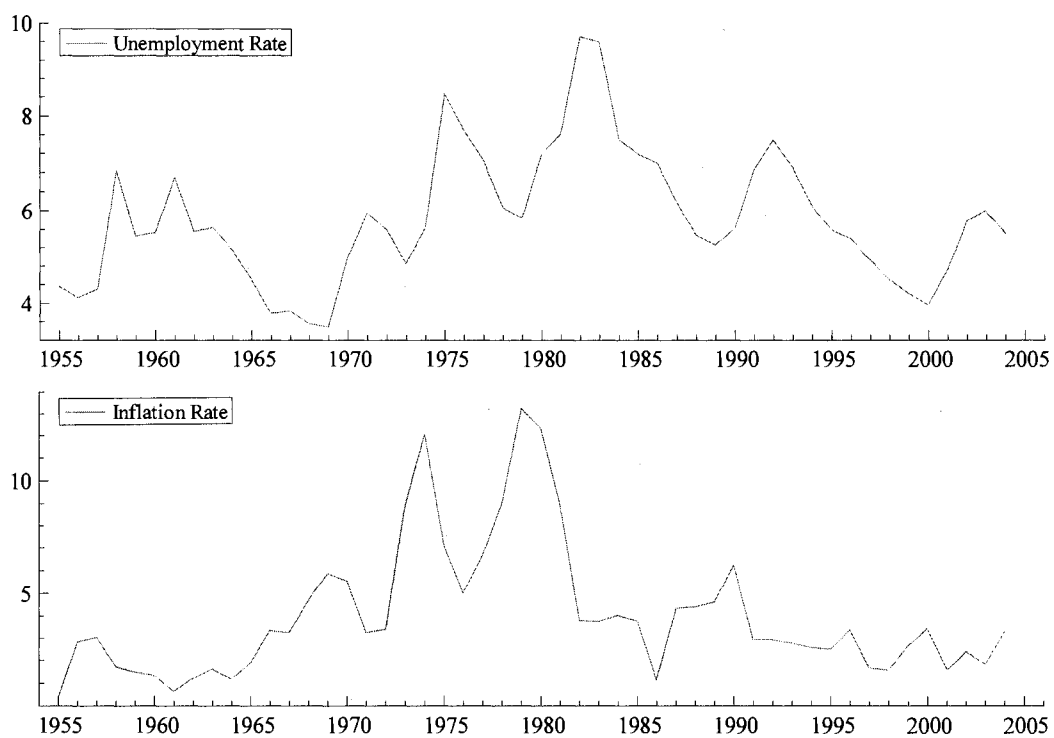
Variable	Smooth Trend	Stochastic Trend	Final Model
<i>inflation</i>	0.039	0.068	0.069
<i>unemp</i>	-0.062	0.041	0.058
<i>growth</i>			0.068
Δ <i>weduc</i>			0.074

Notes: The long-run effects under the heading Smooth Trend are from the Model 5 in TABLE 3. The long-run effects under the heading Stochastic Trend are from the Model 2 in TABLE 4. The long-run effects under the heading Final Model are from the Model 4 in TABLE 5. Long-run multipliers are calculated by dropping the time subscripts in the final models and solving for the dependent variable. Note that the long-run multipliers for *unemp* under the headings Smooth Trend and Final Model equal the impact multipliers.

FIGURE 1: THE RATE OF NEW DIVORCES OVER TIME

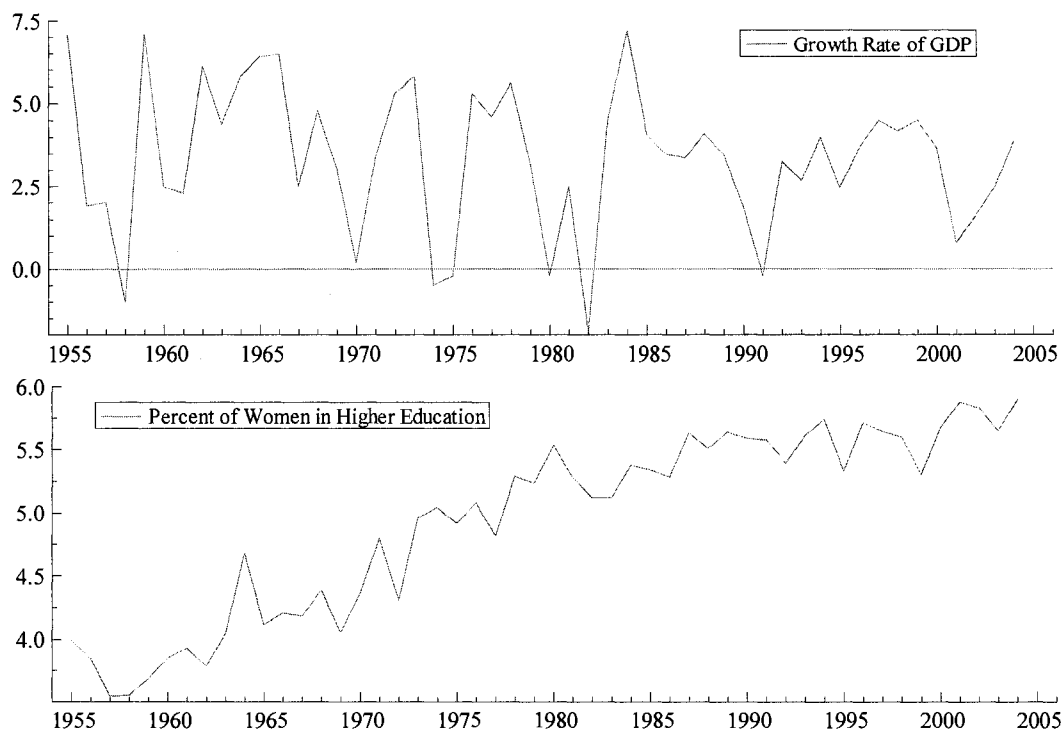


Note: The y-axis measures the number of new divorces per 1,000 people.

FIGURE 2: INFLATION AND UNEMPLOYMENT OVER TIME

Note: The y-axis measures the rate of the explanatory variable.

FIGURE 3: ECONOMIC GROWTH AND WOMEN'S EDUCATIONAL ATTAINMENT OVER TIME



Notes: The y-axis in the upper graph represents the rate of the explanatory variable. In the lower graph, the y-axis represents the percentage of the explanatory variable. However, in the lower graph, the percentage is scaled to be made comparable to the other explanatory variables.

FIGURE 4: RESIDUAL GRAPHICS FOR THE SMOOTH-TREND MODEL

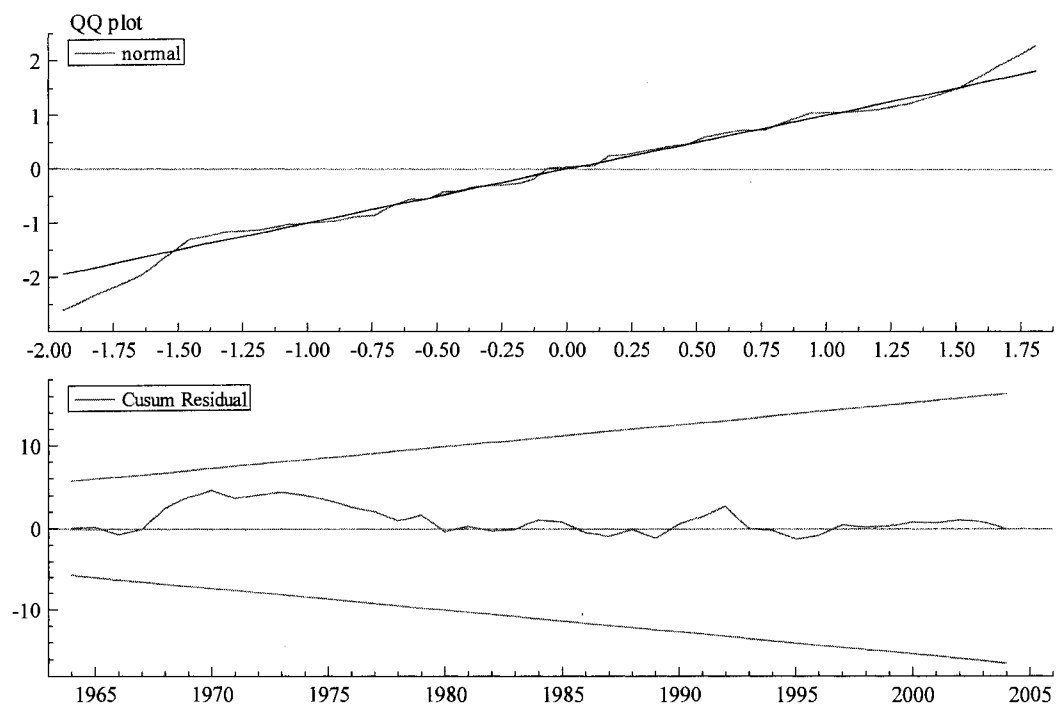
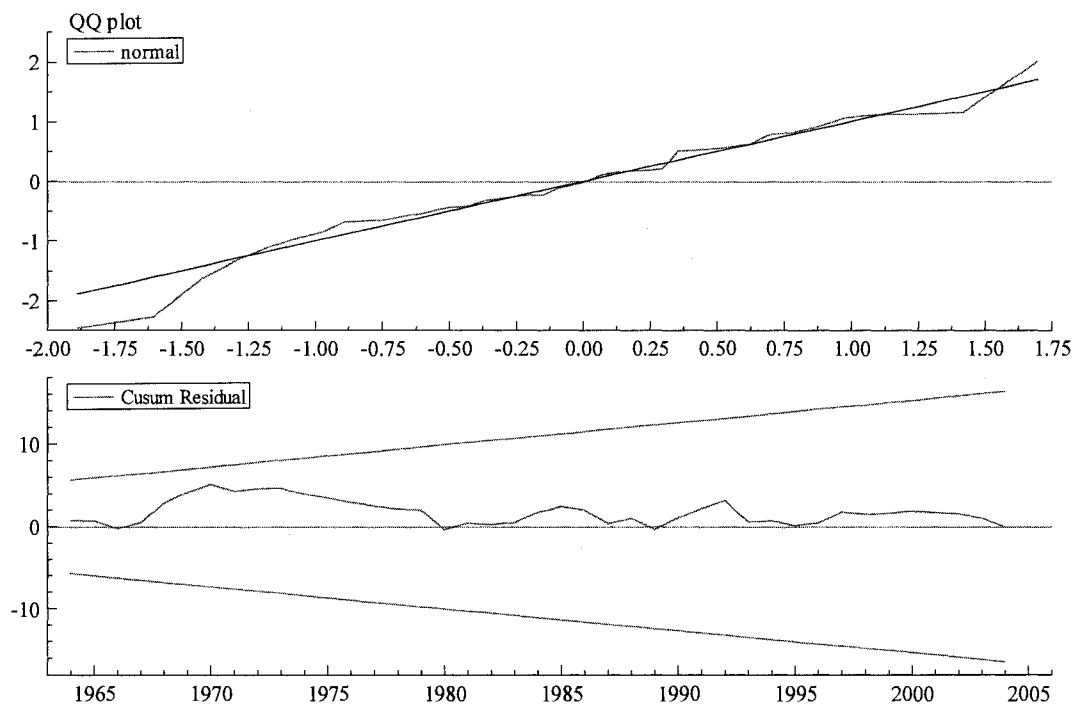


FIGURE 5: RESIDUAL GRAPHICS FOR THE STOCHASTIC-TREND MODEL



**FIGURE 6: REMAINING COMPONENTS FROM THE
SMOOTH-TREND MODEL**

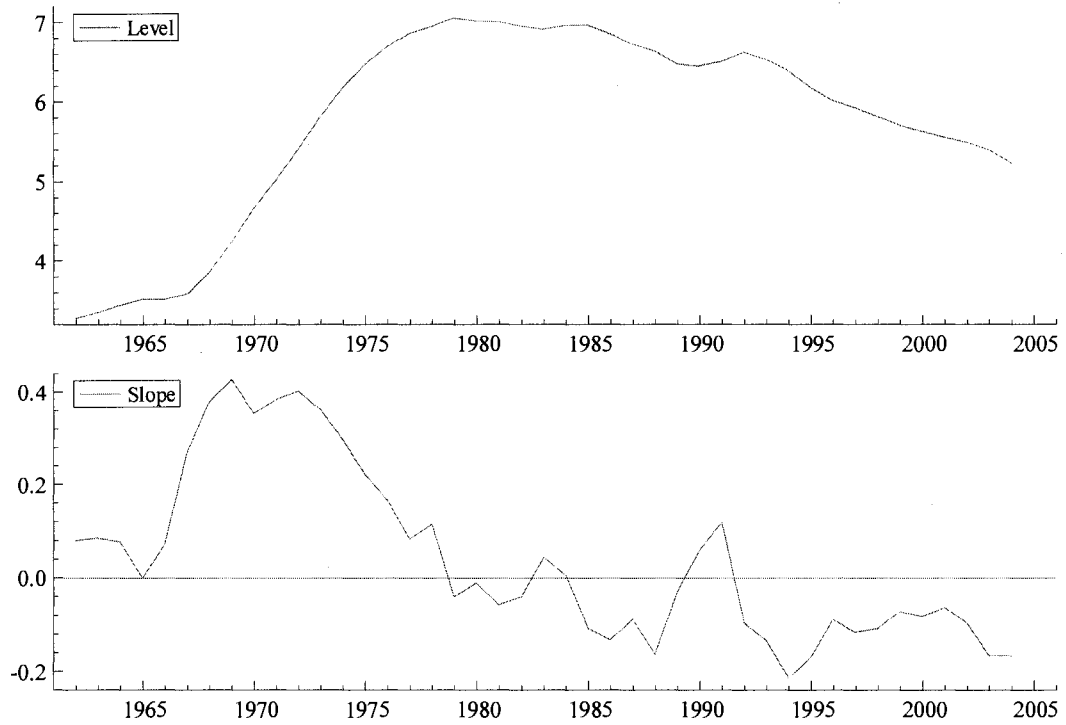
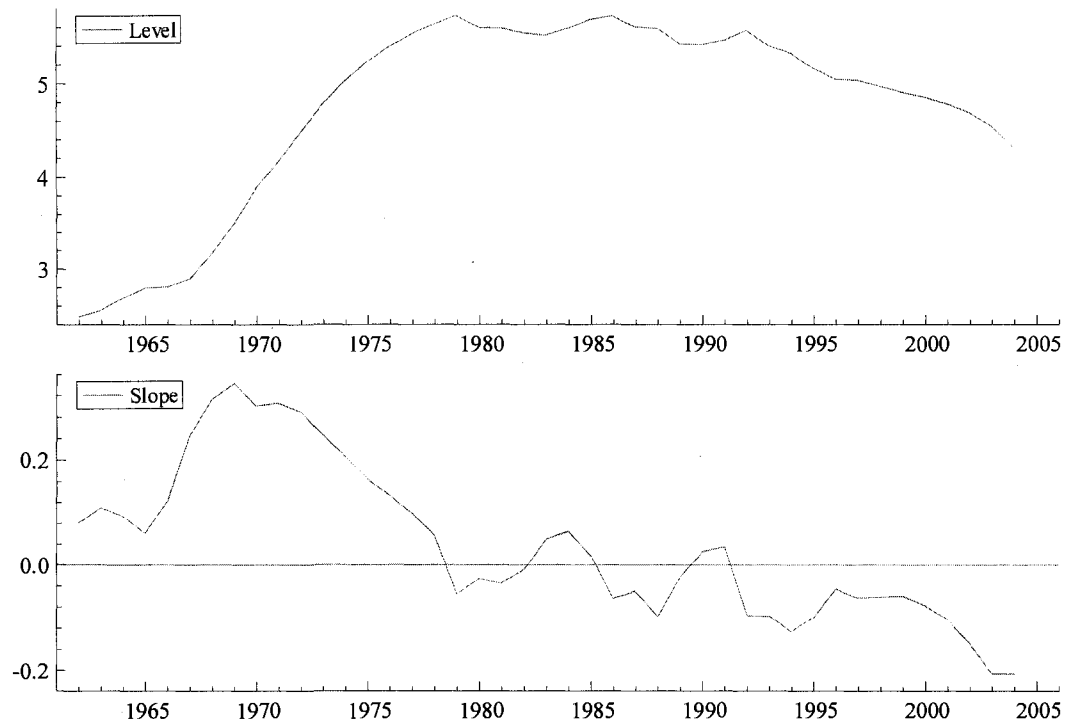
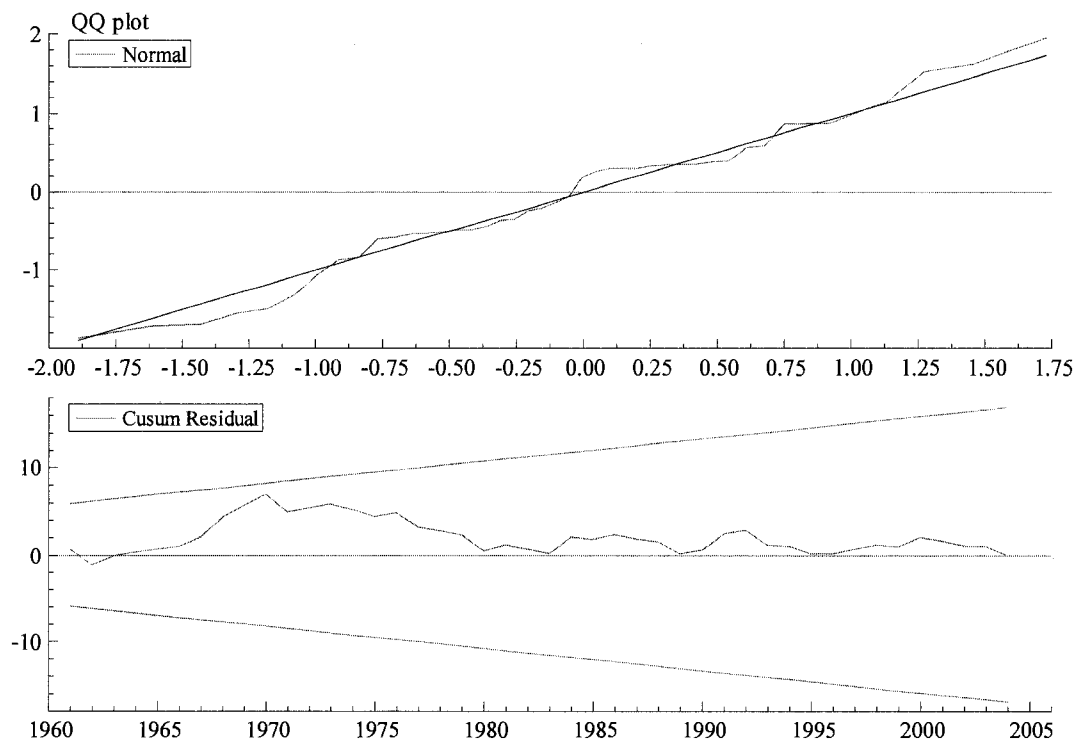


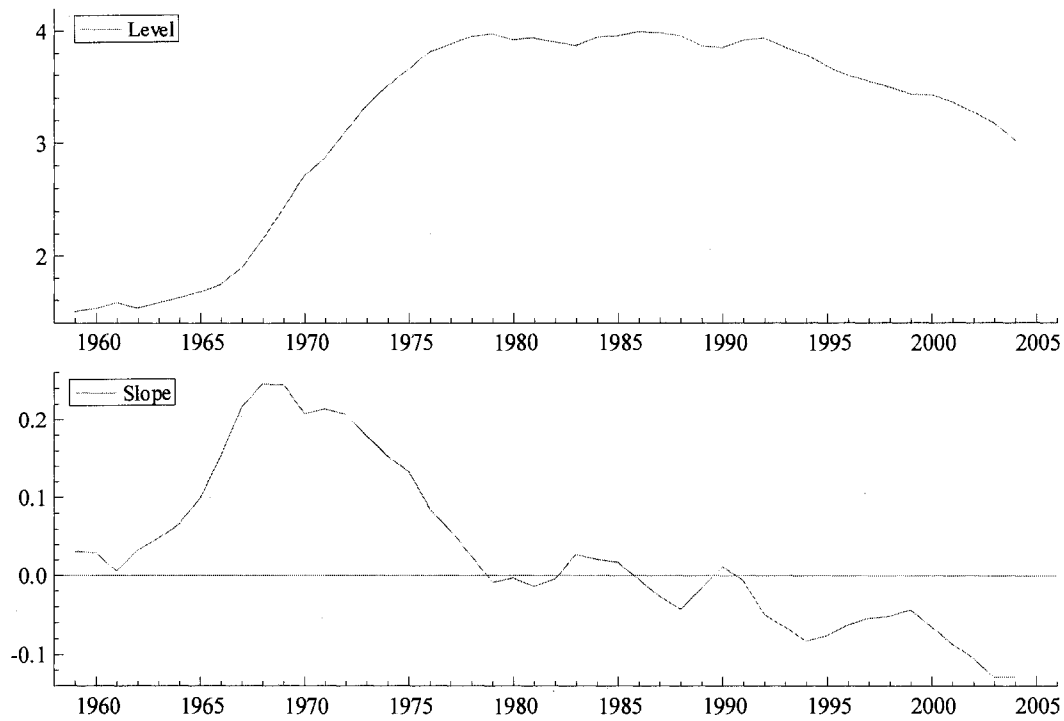
FIGURE 7: REMAINING COMPONENTS FROM THE STOCHASTIC-TREND MODEL



**FIGURE 8: RESIDUAL GRAPHICS FOR THE
FINAL MODEL**



**FIGURE 9: REMAINING TREND COMPONENTS
FROM THE FINAL MODEL**



CHAPTER 4

EXPLAINING THE EVOLUTION OF THE U.S. DIVORCE RATE

(with Joachim Zietz)

4.1. INTRODUCTION

The steady rise in the United States (U.S.) divorce rate throughout the 1960s and 1970s has been a topic of much debate among demographers and economists (Michael 1978; Weitzman 1985; Ruggles 1997a; Ruggles 1997b; Oppenheimer 1997; Preston 1997; Friedberg 1998; Goldstein 1999; Gruber 2004; Wolfers 2006). However, there has been little empirical research that has successfully explained this strong, upward trend (Figure 1). Researchers have focused primarily on the effects on divorce rates of changes in the female labor-force participation rate (FLFPR) and changes in divorce laws. Most research on the relationship between divorce and the FLFPR indicate that causality runs from divorce to a rise in the FLFPR as opposed to the other way around (Johnson and Skinner 1986; Sen 2002). Recent research on the impact of the adoption of unilateral

divorce laws indicates a small, transitory rise in divorce rates, with the effects dissipating within a decade (Wolfers 2006).

The empirical evidence suggests that neither the rise in the FLFPR nor the unilateral divorce law reform fully capture the trend in the aggregate divorce rate (Figure 1), especially that observed for the 1960s and 1970s. The purpose of this study is to capture the trend of the 1960s and 1970s by extending the analysis of the aggregate U.S. divorce rate in a number of ways. First, following Smith (1997), we try to identify to what extent the legal availability of oral contraceptives and divorce law changes have had a measurable impact on the divorce rate. Second, we explicitly consider the impact of the Vietnam War. Third, we extend the analysis back to 1929 to allow for more variation in sample observations. This extension necessitates the construction of meaningful variables for the impact on divorces of World War II (WWII) and the Korean War. In addition, it necessitates substituting a variable for the FLFPR, which does not reach back that far. Female participation in higher education is chosen for this purpose.

In our analysis, we also include covariates used in earlier research. In particular, we include the growth rate of Gross National Product (GNP), the inflation rate, and changes in the unemployment rate. We extend the previous research on the impact of economic growth on the divorce rate by allowing for asymmetric effects inside and outside of recessions.

We find that the econometric evidence provided by the U.S. time series data is not strong enough to identify separately the impact of divorce-law reform and the impact of increased access to the “pill” on the U.S. divorce rate. This is not surprising as both

changes were implemented in many states at about the same time toward the end of the 1960s and early-1970s. Also, the separate estimates for these two variables are so similar that it is impossible to choose among these two variables on the basis of statistical fit. However, by relying on the evidence presented in previous research (Smith 1997; Wolfers 2006), we conclude that the availability of oral contraception is the more likely causal factor for changes in the U.S. divorce rate than divorce-law changes. The Vietnam War is shown to have had a very significant impact on the U.S. divorce rate.

Our key result pertains to the relationship between the FLFPR, or more specifically our proxy for it, and the divorce rate. Significant controversy has been present in the literature on the relationship between the FLFPR and divorce, with some researchers finding that increases in the women's participation in the labor market increased divorce rates (Spitze and South 1985, 1986; Mincer 1985) and others finding that rising divorce rates led to an increase in female labor-force participation (Green and Quester 1982; Shapiro and Shaw 1983; Johnson and Skinner 1986; Lombardo 1999; Sen 2002; Bremmer and Kesselring 2004). We show that the uncertainty about the direction of causality is likely a result of discounting the possibility that the FLFPR and divorce rates are jointly determined or endogenous. More importantly, we demonstrate that the commonly accepted idea that divorce rates and the FLFPR are positively related stems from the fact that previous studies have been too narrowly focused on the years immediately surrounding the introduction of oral contraception, divorce-law changes, and the Vietnam War without explicitly taking into account the impact of these variables on divorce rates. Accounting for these variables and extending the sample back to 1929, we

are able to identify a strong, negative relationship between the divorce rate and the FLFPR or its proxy, female participation in higher education. The years from the mid-1960s to the mid-1970s, which are marked by the diffusion of oral-contraceptives, divorce-law changes, and the Vietnam War, simply shift the negative relationship between divorce rate and the FLFPR, which exists before the mid-1960s and again after the mid-1970s, toward a higher level of both the divorce rate and the FLFPR.

The remainder of this study is organized as follows. Section 2 provides background information on the key variables used in our analysis. In the same section, we also discuss previous work on each of these variables and the theoretical impact that we expect these covariates to have on the divorce rate. Sections 3 and 4 describe the data and econometric methodology, respectively. Section 5 presents our findings. Section 6 provides a brief summary and some concluding remarks.

4.2. INSTITUTIONAL FRAMEWORK

4.2.1. Female Labor-Force Participation and Participation in Higher Education

A number of researchers have analyzed the impact of the rising economic independence of women on the rise in U.S. divorce rates (e.g., Bremmer and Kesselring 2004; Nunley 2008).¹ Economic independence is typically associated with increases in the FLFPR or in the participation of females in higher education. However, using the FLFPR as a proxy for women's rising economic independence may not be ideal because until the late-1960s

¹ There have been numerous micro-level studies examining the impact of educational attainment and female labor-force participation on the probability of divorce (e.g., see Becker et al. 1977; Van Der Klaauw 1996; Weiss and Willis 1997; South 2001; Jalavaara 2003).

and 1970s many women remained secondary earners within households, continued to take their husband's labor-market choices as given, and worked part time with little opportunity for on-the-job advancement (Goldin 2006). Similar to the FLFPR, female participation in higher education has grown steadily since the late-1940s.² Over this period, Goldin et al. (2006) document how the rate of females taking math and science courses in high school converged to that of men. This better prepared them for college and supplied the necessary skills to sort into professionalized fields of study, such as medical, law, business, and dental schools.³ As women increased their economic independence through participation in professional jobs, household labor-market decisions became interdependent, perhaps indicating a shift in bargaining power toward women within households (Costa 2000). The gain in bargaining power from increased participation in professionalized fields suggests that female participation in higher education proxies well for the rising economic independence of women.

Previous empirical work centers on explaining the divorce rate in the 1960s and 1970s (South 1985; Bremmer and Kesselring 2004; Nunley 2008). However, comparison of Figures 1 and 2 suggests that both the FLFPR and female participation in higher education began to rise well before the dramatic rise in the divorce rate in the early-1960s

² Female participation in higher education rose from the late 1930s to the early WWII years, but fell substantially following the war's end, when many war veterans began attending college (Figure 3). Male participation in higher education increased dramatically following WWII because of incentives created by the GI bill, which provided college funding for WWII veterans (Goldin et al. 2006).

³ Since the late-1960s, female participation in the medical, law, dental, and business fields increased substantially (Goldin and Katz 2000; Goldin 2004, 2006; Goldin et al. 2006).

and 1970s.⁴ Identifying the effect of the FLFPR on the divorce rate has proven to be difficult, as there is evidence suggesting that the two variables may be simultaneously determined (Johnson and Skinner 1986; Sen 2002; Bremmer and Kesselring 2004). Using a vector-error-correction mechanism, Bremmer and Kesselring (2004) find a positive, long-run relationship between the divorce rate and the FLFPR for the period from 1960 to 2001. Figure 3, which shows a scatterplot of the divorce rate and the FLFPR over this period, provides insight as to why they find a positive, long-run relationship. The figure reveals that a positive relationship between the two variables existed throughout much of the sample period. Similarly, Nunley (2008) finds a small, persistent, positive effect of an increase in the change in female participation in higher education on the divorce rate. One potential problem with Nunley's (2008) study is the sample period examined, which is 1955 to 2004. Visual inspection of the scatterplot shown in Figure 4 reveals a strong, positive relationship between the divorce rate and female enrollment in higher education during the 1960s and 1970s. However, Figure 4 also indicates a strong, negative relationship between the divorce rate and female participation in higher education prior to 1965 (excluding the WWII years) and from the mid-1970s onward. The negative relationships for these time spans are shown separately in Figures 5 and 6.

Figure 1 shows a transitory rise in the divorce rate during WWII, which returns to pre-war levels following the war's end. By contrast, in the mid-1960s and 1970s, the

⁴ See Goldin et al. (2004) and Goldin (2004, 2006) for a discussion of trends in the FLFPR and female educational attainment.

divorce rate permanently shifts to a higher level. There are several factors that altered family life over this period: the diffusion of oral contraceptives, the Vietnam War, and changes in divorce laws, with all of these having potentially a positive impact on divorce rates. As such, studies too narrowly focused on the rise in the divorce rate in the 1960s and 1970s will necessarily identify a positive relationship between female participation in higher education and divorce.

The theoretical relationship between female participation in higher education and the divorce rate is not clear.⁵ Increases in female participation in higher education improve women's employment opportunities by allowing selection into professionalized fields. Improvements in employment options through women's rising educational attainment may lead to a rise in divorce rates because there may be a reduction in the returns from marriage. Likewise, women may no longer be reliant on the incomes of husbands. The returns from marriage fall in Becker's (1973, 1974, 1991) traditional family model because the household production function is no longer maximized, as increases in wives' participation in the labor market, with no change in husbands' labor-market behavior, imply less of households' time devoted to leisure and home production. However, the advent of labor-saving technologies and the ability to purchase household services in the market could mean that the traditional model of household behavior no

⁵ Most studies use microdata to investigate the relationship between education and divorce (South and Spitze 1986; Weiss and Willis 1997; South 2001; and Jalovaara 2003). South and Spitze (1986) find that the effect of educational attainment on divorce depends crucially on the duration of the marriage, with a decrease in the probability of divorce for newlyweds and an increase in the probability of divorce for marriages of lengthy durations. Researchers have also found that the divorce risk is lower when both spouses have similar education levels, but the divorce risk increases when spouses have heterogeneous education levels (Weiss and Willis 1997; Jalovaara 2003). By contrast, no relationship between women's rising educational attainment and the divorce rate is found by South (2001).

longer applies (Stevenson and Wolfers 2007). As such, both spouses participating in the labor market, especially if both spouses work in professional fields, may allow greater returns from marriage through increased efficiency from household technologies, purchases of household services, and increases in consumption and leisure. In this sense, there could be returns from marriage when both spouses work, implying a reduction in divorce rates.

4.2.2. Increased Access to Oral Contraceptives

In 1957, the Food and Drug Administration (FDA) approved the use of Envoid—the first oral contraceptive—for medical use. Three years later, the drug was approved for the purpose of oral contraception (Goldin 2001). Although most states only allowed married women access to “the pill,” states passed laws providing unmarried women access to the pill in the 1960s and 1970s.⁶ Complete diffusion of the pill to all women was brought about by the passage of Twenty-Sixth Amendment to the U.S. Constitution in 1971, which provided young individuals additional rights (Goldin and Katz 2000, 2002; Bailey 2006). By 1976, all states had adopted some form of legislation permitting all women access to oral contraception (Bailey 2006).⁷

Goldin and Katz (2000, 2002) contend that access to the “pill” gave women more control over fertility decisions, which meant a decline in the opportunity costs of human-capital investment and an increase in age at marriage. Additions to human capital

⁶ 41 percent of married women under the age of 30 were using oral contraceptives by 1965 according to Goldin and Katz (2002).

⁷ See Goldin (2001), Goldin and Katz (2000, 2002), and Bailey (2006) for a detailed discussion of state-level and national reforms allowing access to the pill.

increase the opportunity for higher earnings and favorable career options. This provides additional bargaining power to women within households (Costa 2000).⁸ The increase in bargaining power of women stems from a higher value of options outside of marriage, which includes the time after a divorce, as women become economically independent. The incidence of divorces may increase as a consequence. By contrast, higher education may also lower the attractiveness of divorce because two-income couples, who are able to delay fertility, may be able to invest more in other forms of marriage-specific capital. Two-earner couples, especially when both spouses work in professional fields, may have the means for significantly greater consumption and leisure, which lowers the attractiveness of divorce.

Goldin and Katz (2000, 2002) find that access to the pill increased women's age at marriage. Because increases in age at marriage can improve marital sorting through a reduction in the opportunity costs of postponing marriage, oral-contraceptive use created a "thicker" marriage market (Goldin and Katz 2000, 2002). Improvements in the marriage market imply an increase in marriage-match quality because the cost of marital search is lower.⁹ Increases in marriage-match quality and age at marriage may have the potential to reduce divorce rates.

Although Goldin and Katz (2002) find a negative effect of access to oral contraceptives on divorce among college-educated women, there is also credible evidence of a positive effect. Using time-series data from England and Wales, Smith

⁸ Goldin and Katz (2002) also find that the pill had a positive effect on women entering into professional schools.

⁹ See Weiss and Willis (1997) and Charles and Stephens (2004) for comprehensive analyses and discussion on the importance of marriage-match quality.

(1997) shows that access to oral contraceptives increased the divorce rate in both regions. The evidence is somewhat tenuous, however, as Smith's (1997) "pill effect" is based on a smooth, diffusion function in which only the starting point and the ending point of the diffusion process are known.

Access to the pill could have a positive effect on the divorce rate for several reasons. First, the traditional model of marriage developed by Becker (1973, 1974, 1991) implies a reduction in the returns from marriage, as the pill led women to participate more in market work (Bailey 2006) and allowed college educated women to become doctors, lawyers, and enter into other professions (Goldin and Katz 2000, 2002). The traditional family model posits separate spheres for husbands and wives, suggesting that spouses should specialize in the sphere in which they have a comparative advantage. If spouses choose to specialize in the same sphere (i.e. market work), the traditional family model predicts a decline in the returns from marriage. Second, the pill has been shown to reduce fertility (Bailey 2006). It is well established that increases in marriage-specific capital, such as children, decrease the risk of divorce (Becker et al. 1977). Third, individuals may sort into "bad" marriages, as they can delay fertility. The ability to delay child births may allow spouses to try out potential mates before having children. Put differently, spouses may sort into riskier marriages because the costs of divorce are lower when fertility can be controlled. Fourth, the pill may make extramarital affairs more likely by reducing the perceived costs, as the likelihood of an unwanted pregnancy is lower.

Figure 8 plots the percentage of the U.S. population affected by early access to the pill over time. The figure reveals a sharp increase in the diffusion of the pill in the mid-1960s, which coincides with the start of the sharp increase in the divorce rate. The full distribution of the pill to the population is achieved by 1976. Figure 9 provides a scatterplot of the percentage of the population affected by early access to oral contraceptives and the divorce rate. It indicates that the two variables are positively related over the time period from the mid-1960s to the mid-1970s. However, the points for the years 1968-1971 are above the fitted trend line, an indication that another force was at work during this time. As discussed in the next section, these years coincide with the heightened intensity of the Vietnam War, which may have shifted the divorce rate to an elevated level above and beyond what can be attributed to the use of oral contraceptives.

4.2.3. World War II, the Korean War, and the Vietnam War

From the Civil War to the present, war has had significant effects on family outcomes (Povalko and Elder 1990). During wartime, some marriage decisions come about more quickly, while others are delayed. Of those marriages that form during war time, many end in divorce. Most researchers use time-specific, indicator variables to capture the effects of war on divorce (South 1985; Anderson and Little 1999).¹⁰ South (1985) finds that the divorce rate rose during the Vietnam War period, but no statistical evidence

¹⁰ Typically, major wars and their impact are not captured in numerical form. Most of the time, the war years are left out of empirical work completely or they are absorbed, but not explained, with a set of indicator variables (e.g., see South 1985; Anderson and Little 1999).

linking the Korean War to the divorce rate. By contrast, no relationship is identified between the divorce rate and the Vietnam War years by Anderson and Little (1999), but they substantiate South's (1985) results for the Korean War. Using cross-section data, Pavalko and Elder (1990) find that WWII veterans were more likely to divorce than nonveterans. Confirming Pavalko and Elder's (1990) estimates with time-series data, Anderson and Little (1999) also establish a statistically significant, positive effect of the WWII years on the divorce rate.

Figure 1 shows sharp increases in the divorce rate during and following WWII and the Vietnam War.¹¹ However, the sharp increase in the divorce rate during and shortly following WWII appears to be temporary, with the divorce rate returning to pre-WWII levels following the war's end. For the time period of the Vietnam War, the increase in the divorce rate appears to be prolonged, with the divorce rate increasing steadily until the mid-1970s. The divorce rate continued to increase in the late-1970s, but at a much slower rate than in the mid-1960s to mid-1970s.

During the years of the Korean War (1950-1953), there appears to be little variability in the divorce rate although the Korean War had a similar number of casualties and intensity as the Vietnam War. A potential explanation is that the Korean War closely followed WWII, which may suggest that it affected essentially the generation that had grown up during WWII. The generation affected by both WWII and the Korean War experienced either the wars itself or the stress of the war as it was brought home by

¹¹ Large increases in divorce rates can be seen at the end of WWII both for the U.S. (Figure 1) and Great Britain (Figure 11). Despite a strong similarity in the change in divorce rates there is a notable difference in the level of divorce rates for the two countries.

millions of returning soldiers. In other words, war was not a new experience for this generation, as it had been actively or passively involved in war. This is different for the Baby Boom generation born after WWII. The Vietnam War was its first exposure to the reality of war. To that generation it was a shock, which was brought about by the generation of the parents.

What added significantly to the shock were a number of things. First, the Vietnam War was the first war with large scale TV coverage (Farenick 1993). It was present in the daily news reports not only of the U.S. but of any western nation. Second, the support of a corrupt regime in Vietnam was seen as unjust, especially by a Baby Boom generation that had started rebelling against what was perceived as something rather similar: the straightjacket of 1950s style societal rules and conventions, which included the 1950s shift toward family values with increasing marriage and fertility rates (Cherlin 1981, Ch.2). The rising intensity of the Vietnam War fueled the latent demand for radical changes that was present all over the western world.¹² The Vietnam War had in some sense become a catalyst for the desire to change society (Buzzanco 1999, pp. 147). Third, the war exposed millions of young men, a large cross section of the younger population of the U.S. who are prime candidates for marriage and divorce, to the horrors and stress of war. Fourth, as an outgrowth of the war there was for the first time a significant influx of narcotics into the U.S. and the western world at large. All four changes that are connected with the Vietnam War had a significant impact on the Baby

¹² The anti-war movement during the Vietnam era is well documented (e.g., see DeBenedetti and Chatfield 1990; Olson 1993; Wells 1994; Garfinkle 1995; Buzzanco 1999; Tischler 2002).

Boom generation. As the 1950s lifestyle was abandoned, divorce turned into an acceptable option.

4.2.4. Divorce-Law Reform

Over the period of rising divorce rates, many states adopted unilateral divorce laws, thereby allowing divorce on demand by either spouse. Enactment of unilateral divorce transfers the right to exit the marriage to the spouse who prefers divorce over remaining married.¹³ The impact of divorce-law changes on divorce rates has been studied extensively by researchers in sociology (Nakonezny et al. 1995; Glenn 1999; Rogers et al. 1999), law (Brinig and Buckley 1998; Ellman and Lohr 1998; Ellman 2000), and economics (Peters 1986, 1992; Friedberg 1998; Gruber 2004; Méchoulan 2006; Rasul 2006; Wolfers 2006). The findings of these studies are somewhat mixed; however, a majority of studies identify a small, positive effect of unilateral divorce laws on divorce rates.¹⁴

Figure 11 shows the diffusion process through which unilateral divorce spread throughout the U.S. A comparison of Figures 1 and 11 shows that the divorce rate was trending upward prior to the implementation of unilateral divorce reform, regardless of the divorce-law coding used. Divorce rates began trending upward around 1965, while the majority of states that adopted unilateral divorce did so during the late-1960s and 1970s, with most adopting unilateral divorce in the early-1970s. Because Friedberg's

¹³ See Weitzman (1985) and Jacob (1988) for a more detailed discussion of divorce laws.

¹⁴ Fella et al. (2004) posit a theoretical model providing support for the view that changes in social norms instead of divorce-law reform led to the permanent rise in divorce rates.

(1998) empirical specification begins in 1968, it is likely that the estimated effects presented in her study reflect preexisting trends rather than the impact of the legal change. Friedberg's model may also be misleading, as the empirical specification only captures a one-time (permanent) change in divorce rates from the adoption of unilateral divorce laws. This assumption is restrictive because the short-run effect could be very different from the long-term response.¹⁵ Wolfers (2006) provides support for this idea by extending Friedberg's model by increasing the sample length—beginning 1956 and ending in 1998—which allows for the short- and long-run effects of unilateral divorce reform to be identified.¹⁶ Another advantage of Wolfers' specification is it allows for preexisting trends to be separated from the adoption of unilateral divorce laws.

Until the work of Wolfers (2006), Friedberg's (1998) results were considered the most accurate. According to her estimates, one-sixth of the trend in U.S. divorce could be explained by the adoption of unilateral divorce. Wolfers (2006) finds a small, temporary rise in divorce rates following the reform, with the effects dissipating within a decade. In fact, somewhat puzzlingly, Wolfers concludes that unilateral divorce laws reduce divorce rates in the long run. The small transitory effects found by Wolfers (2006) suggests an alternative cause for the rise in divorce rates in the 1960s and 1970s.

¹⁵ Similar to Wolfers (2006), Smith (2002) suggests a need to estimate the full adjustment path of divorce rates to divorce-law reform, as there could be "pent-up" demand for marriages with low match quality.

¹⁶ Friedberg's (1998) sample ends in 1988.

4.2.5. Macroeconomic Conditions

There seems to be some consensus on the relationship between economic growth and divorce rates. Divorce rates tend to rise during economic expansions and decline during economic contractions (e.g., see Goode 1971; Nunley 2008).¹⁷ However, South (1985) finds the opposite.¹⁸ The traditional marriage model, as posited by Becker (1973, 1974), predicts a decline in the returns from marriage if both spouses participate at a higher rate in the labor market. This is because the gains from specialization decline when both spouses work in the market, which means that less of the household's time is allocated to leisure and home production. However, this implies that the substitution effect dominates the income effect generated by economic growth. It is possible for the income effect to dominate. Because economic growth generally implies increases in earnings and higher returns on investment, both spouses could achieve a higher degree of specialization, which translates into an increase in the returns from marriage. However, as suggested by Stevenson and Wolfers (2007), household specialization may have changed, or at least may have begun to take on a new meaning. As such, the returns to marriage could increase because higher labor-force participation for spouses implies increases in earnings, which could generate greater household consumption, leisure, and gains from purchasing household services in the market.

¹⁷ See Figure 1A in the Appendix for a time-series plot of the growth rate of GNP.

¹⁸ South (1985) considers the percentage change in GNP and the unemployment rate as proxies for the overall health of the economy. In South's model specification, he includes the percentage change in GNP but not the unemployment rate, and vice versa. It is possible that the growth rate of GNP and the unemployment rate are correlated. In fact, a reduced form ordinary least squares (OLS) regression of the growth rate of GDP on the unemployment rate suggests that the two variables are negatively correlated, which may explain the negative effect found with respect to the percent change growth rate of GNP.

Divorce-threat (also referred to as exit-threat) bargaining models, as posited by Manser and Brown (1980) and McElroy and Horney (1981), predict a rise in the value of options outside of marriage (i.e. divorce). Economic growth improves spouses' outside options through greater job availability, higher incomes, and higher return on investment. Economic growth, by improving outside options, could generate a self-reliance effect, which could raise divorce rates.

Another macroeconomic indicator that has been shown to have sizable, persistent, and positive effects on the divorce rate is the inflation rate (Nunley 2008).¹⁹ An increase in the inflation rate worsens the terms of trade between spouses, which reduces the returns from marriage, as spouses who specialize in home production may be forced to enter the labor market in order to achieve pre-inflation, consumption levels. If spouses are forced to work more in the market and continue to work in the home, time allocated to leisure declines. A decrease in leisure also reduces the returns from marriage. However, it is possible for wage increases to offset rising prices. In fact, it has been shown that prices respond more quickly than wages to positive money-supply shocks (Christiano et al. 2005), which suggests a decrease in the returns from marriage when the inflation rate rises. Regardless of whether one uses the family model that implies gains from specializing in separate spheres or one in which consumption complementarities define the returns from marriage, inflation reduces these returns because it acts as a tax on the household.

¹⁹ See Figure 2A in the Appendix for a time-series plot of the inflation rate.

4.3. DATA

This study uses time-series data from 1929 to 2006, while the majority of studies use cross-section or panel data to investigate factors that affect divorce rates (e.g., Becker et al. 1977; Weiss and Willis 1997; Charles and Stephens 2004; Hess 2004). One advantage of using time-series data is that we are able to examine a much longer time horizon, which allows us to analyze both the short- and long-run dynamics of the divorce rate. However, using this long sample also requires some solutions to apparent data problems, such as the unavailability of some variables all the way back to 1929, which includes the FLFPR. We use female participation in higher education as a substitute for the FLFPR, because annual data on the FLFPR are only available from 1948-2006. There are missing years of data even for female participation in higher education: only odd years are reported for the years from 1929 to 1945. We replace the missing years of data with the average of the odd years. For example, female participation in higher education in 1930 is taken to equal the average of the 1929 and 1931 values. Figure 2 plots the FLFPR and female participation in higher education over time. The figure shows that the two variables display very similar trends from 1948 to 2006. The similarities of the two data series are revealed more explicitly in Figure 7, which displays a scatterplot of the FLFPR and female participation in higher education with a least-squares fit line.

Our primary outcome variable is divorces per 1,000 persons. We use this variable instead of divorces per 1,000 married couples because the Center for Disease Control (CDC) stopped collecting data on divorces per 1,000 married couples in 1997. This

measure can be approximated by manipulating the divorces per 1,000 person's variable.²⁰ Despite our ability to estimate reasonable estimates for divorces per 1,000 married couples, we continue to use divorces per 1,000 persons because the two variables display similar behavior over time (Figure 12). In fact, a scatterplot of divorce per 1,000 persons and divorces per 1,000 married couples reveals a clean, linear relationship between the two variables (Figure 13). This suggests that the estimated effects of our explanatory variables would be similar regardless of which divorce measure used.²¹

Figures 8 and 11 depict the behavior of two variables of some importance for explaining changes in the divorce rate: the percentage of the population affected by legal changes allowing increased access to oral contraceptives and unilateral divorce. These variables are constructed by dividing aggregated state populations that adopt the law in a year by the total U.S. population in that year, which effectively form diffusion functions for each variable. The complete diffusion of "the pill" occurred in 1976, 16 years after the first legal change providing access to young, unmarried women. The diffusion process for unilateral divorce depends on the law coding used. We primarily use Friedberg's (1998) and Gruber's (2004) codings but check the robustness of our results to

²⁰ Multiplying the number of divorces per 1,000 persons by the U.S. population per 1,000 persons gives the total number of divorces. Dividing this number by the stock of married couples creates the variable of interest: divorces per 1,000 married couples. This measure is created using various U.S. Statistical Abstracts. Data on divorces per 1,000 married couples are available until 1995. Therefore, to check our estimates for the years 1996-2006, we use the same calculation method described above for the available years and find that any difference in the estimates is in the decimal places.

²¹ In fact, our findings are robust to the divorces per 1,000 married couples' measure. The effects are somewhat larger. This is not surprising, as the divorce rate per 1,000 married couples is a much larger number than the divorce rate per 1,000 persons. The results are presented in the Appendix.

Méchoulan's (2006) coding. The diffusion process for each of the law codings for unilateral divorce is shown in Figure 11.²²

To incorporate the effects of major wars on divorce rates, we construct variables that proxy for the intensity or "stress" of the war. We use casualties relative to deployments for this purpose, which allows us to create an objective measure of stress and/or the intensity of the war (Figure 14). More specifically, the variable that proxies for the "stress" of the Vietnam War is defined as U.S. military personnel deaths due to the Vietnam War as a fraction of U.S. military deployments in East Asia. Deaths include not only combat casualties but also deaths due to accidents and death from wounds suffered in Vietnam but occurring elsewhere.²³ The overall death figures are highly correlated with the combat death figures and the large number of wounded: all three are driven by combat intensity. The death counts in the numerator of the Vietnam War variable relate to deaths associated with the Vietnam War. The denominator of the variable measures

²² One potential problem with examining laws allowing unilateral divorce and increased access to oral contraceptives and divorce rates is whether the reforms were determined independent of divorce rates. Historical accounts provided by Jacob (1988) suggest that unilateral divorce reform is credibly exogenous. One of the main reasons that states adopted unilateral divorce was to simplify divorce proceedings. Before unilateral divorce reform, proof of marital wrongdoing was required for a divorce to be granted. These restrictions led spouses who preferred divorce over marriage to admit fault even when no such fault was committed. Deception was also a concern among lawmakers, because one spouse could extort more of the marital surplus from the party wishing to exit the marriage. Legal reforms that allow young, unmarried women increased access to oral contraceptives are also plausibly exogenous. Historical accounts indicate that access to contraceptives was brought about by issues pertaining legal rights, further supporting the exogeneity of the reform. Goldin and Katz (2002) and Baily (2006) argue that the Vietnam War helped fuel the demand for extending the rights of younger individuals.

²³ There is no official account of the deaths occurring outside of Vietnam due to wounds suffered in Vietnam. These deaths are not officially counted as Vietnam casualties and have not been made public. However, their number is significant and is estimated to exceed the number of combat deaths starting in 1971. Some estimates of the combined number of casualties are available for the Army, but not for all service branches. The Army figures are used in the calculations of deaths. They come from the following source: <http://www.thetruthseeker.co.uk/article.asp?ID=2703>.

U.S. troop deployment in all of East Asia, not just Vietnam.²⁴ This is to account for the fact that many soldiers killed in Vietnam operated from bases in East Asia outside of Vietnam. For all practical purposes, any deployment for a soldier to East Asia during the Vietnam War could mean to get into harm's way in Vietnam.

The data series for the Korean War and WWII are derived in analogy to those of the Vietnam War. Casualties by year are divided by troop deployments.²⁵ Troop deployments in East Asia are used for the Korean War.²⁶ Troop deployments for WWII are culled from Matloff (1990).²⁷

We also include two standard macroeconomic variables: the inflation rate and economic growth. We also include squared terms of certain variables, as well as an interaction term between economic growth and an indicator variable for whether the economy is in a(n) recessionary or expansionary period. The interaction variable captures asymmetric effects of economic growth in recessionary and expansionary

²⁴ Annual figures on U.S. troop deployments by region and country from 1950 onwards are taken from the March 1, 2006, "U.S. Troop Deployment Dataset" as compiled by Tim Kane of the Heritage Foundation, Center for Data Analysis, Washington, D.C. (<http://www.heritage.org/Research/NationalSecurity/cda06-02.cfm>).

²⁵ For the Korean War, casualties by year are derived from "State-level Lists of Casualties from the Korean War (1951-1957)," The U.S. National Archives and Records Administration, College Park, MD, (<http://www.archives.gov/research/korean-war/casualty-lists/>). For WWII, Navy casualties by year are calculated from data provided in "U.S. Navy Personnel in World War II: Service and Casualty Statistics", as taken from "Annual Report, Navy and Marine Corps Military Personnel Statistics, 30 June 1964," Bureau of Naval Personnel and U.S. Marine Corps Headquarters, Naval Historical Center, Department of the Navy, Washington, D.C. (http://www.history.navy.mil/library/online/ww2_statistics.htm); Army casualties by year are taken from page 99 of "Army battle casualties and nonbattle deaths in World War II: Final Report." Office of the Comptroller of the Army, Program Review & Analysis Division, Office of the Adjutant General, Washington, DC, 20310, 1946 (<http://stinet.dtic.mil/oai/oai?verb=getRecord&metadataPrefix=html&identifier=ADA438106>).

²⁶ The source of the deployment figures is the same as that used for the Vietnam War.

²⁷ Available at <http://www.history.army.mil/books/wwii/sp1943-44/index.htm> (chapter 17, Tables 4 and 5 and Appendix E). The deployment figures for 1946 that are related to WWII deployments are assumed to be one sixth of all troops stationed overseas during that year.

periods on the divorce rate. Including squared terms of some explanatory variables allows for nonlinear responses. These variables are required in some models because Ramsey's (1969) RESET test reveals significant functional form misspecification.

Table 1 provides variable names, definitions, and sources, while Table 2 presents basic statistics for 1929-2006, 1929-1948, and 1949-2006. Because of the apparent structural break for some of the models around 1948/49, we estimate models using different sample periods: 1949-2006 and 1929-2006. Examining the summary statistics for periods 1929-1948 and 1949-2006, there is far less volatility in the variables from 1949-2006. Figures 1A and 2A further support this, as more pronounced volatility appears before WWII of the inflation rate (excluding the 1970s) and economic growth. When we examine the full sample period, we use the least-absolute-deviations estimator to allow for outliers in the data.

We test each of the variables used in our analysis for the presence of a unit root and stationarity, indicating that the variables *divorce*, *fem_ratio*, *fem_ratio*² are I(1). The evidence is somewhat ambiguous for the *Vietnam* variable. The variables *WWII*, *Korea*, and *Vietnam* are in one sense dummy variables because they have nonzero values only for a limited number of observations. However, in another sense, the variables are different from a dummy variable because their nonzero values are not equal to unity but are based on observed figures of casualties and the degree of military involvement by the U.S. The variables *pill*, *gruber*, *friedberg*, and *mechoulan* are also nominally I(1). However, they are not following random walks but are the result of a stable diffusion process. The variables *inflation*, *ygrowth*, *ygrowth*², *das*ygrowth*, and $\Delta unemp$ are I(0).

4.4. ECONOMETRIC METHODOLOGY

To encompass earlier results as much as possible, the study uses two basic types of approaches, single-equation estimators and system estimators. Since the divorce rate and some other variables are non-stationary, some care is needed to avoid spurious results. We employ a large number of statistical specification tests to rule out spurious results for single-equation estimates. In addition, we check whether the single-equation models contain any stochastic or non-stochastic trend components that are not captured by included right-hand-side variables. For that purpose, the study uses the unobserved component methodology (UCM) of Harvey (1989) as further elaborated by Durbin and Koopman (2001).

Accounting for unobserved stochastic trends is important for single-equation estimators because the model contains a nonstationary variable as the dependent variable (i.e. the divorce rate) and one or more nonstationary variables on the right-hand side. We check for the presence of unobserved-trend components by expanding the OLS regression into an UCM, in which the unobserved components are modeled as flexible stochastic-trend components. More formally, the unobserved-component model takes the form

$$y_t = \mu_t + \sum_i \sum_j \alpha_{ij} x_{i,t-j} + \varepsilon_t \quad \text{for } t=1,2,..,T \quad (1)$$

where y_t is the dependent variable and μ_t the time-varying unobserved component; $x_{i,t-j}$ represents explanatory variable i subject to time lag j ; $\alpha_{i,j}$ denotes the coefficient

associated with the variable; and ε_t is a zero mean, constant variance, irregular component. The term μ_t represents the unobserved stochastic trend component that differentiates equation (1) from an OLS regression equation. It captures the impact of unobservables and omitted variables that influence the dependent variable. By removing their influence from the error term, the irregular component is uncorrelated with the variables in $x_{i,t,j}$. This makes for unbiased coefficient estimates. The stochastic trend, μ_t , takes the form:

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t \quad \eta_t \sim NID(0, \sigma_\eta^2) \quad (2)$$

$$\beta_t = \beta_{t-1} + \xi_t \quad \xi_t \sim NID(0, \sigma_\xi^2) \quad (3)$$

The term μ_t is the “level component” of the unobserved stochastic trend and β_t its “slope”. Equation (2) is modeled as a random walk with drift and equation (3) as a pure random walk. The terms η_t and ξ_t are white noise disturbances that are assumed independent of each other and of ε_t . The terms σ_η^2 and σ_ξ^2 are the hyper-parameters that define the stochastic trend μ_t , which need to be estimated. Once they are known, the state vectors μ_t and β_t can be retrieved from the model.²⁸ To the extent that one or both of the hyper-parameters are zero, the unobserved stochastic trend simplifies. In the limiting case, in which both hyper-parameters are zero, the stochastic trend model collapses to OLS, either with or without a deterministic trend, depending on whether the

²⁸ See Harvey (1989) for a detailed description of structural time-series models. The statistical package used—Structural Time-Series Analyser, Modeller, and Predictor (STAMP)—offers a convenient estimation procedure (see Koopman et al. 2000).

drift term β_t is different from zero. If σ_η^2 equals zero and σ_ξ^2 is nonzero, the model takes the smooth-trend specification, which is integrated of order two (Harvey 1997).

If no stochastic or deterministic trend can be verified, a model can be trusted not to be subject to any stochastic trend.²⁹ The absence of a trend component implies that the right-hand-side variables capture the data generating process (DGP) without the help of an unobserved component, which captures all other influences either correlated or uncorrelated with the included explanatory variables. As we shall see, single-equation models with “traditional” right-hand-side variables cannot fully explain the U.S. divorce rate in the late 1960s and early 1970s. This tends to give rise to a strong, stochastic-trend component for this time period, which is an indication of omitted variables. It appears that the observed variables that we include to explain the divorce rate eliminate any sign of or need for a stochastic trend.

Single-equation methods, whether in the form of OLS or UCM, assume that the right-hand-side variables are at least weakly exogenous. That may be a strong assumption in the present context with divorce and female participation in higher education. Therefore, we investigate the need for a system estimator, one that allows for both variables to be endogenous and can also account for the nonstationarity of both variables. The empirical evidence strongly indicates the need for systems approach to cointegration as suggested by Johansen and Juselius (1990). As this methodology is well

²⁹ It should be noted that spurious stochastic trends will in most cases also be detected by simple specification tests of the type provided in conjunction with all least-squares models of Tables 4 and 7 (e.g., see Zietz 2000).

established, we refrain from a formal discussion but refer to standard, up-to-date sources such as Juselius (2006) and Lütkepohl (2007).

4.5. RESULTS

4.5.1. Primary Specifications

Table 4 provides least-squares estimation results. The table is organized around four models. Model 1 uses a minimal number of regressors. It indicates that the divorce rate is subject to a moderate degree of persistence. The lagged dependent variable is less than 0.9, which deflects potential problems associated with unit-root processes. However, Ramsey's RESET test indicates that the functional form of the model is improperly specified. Structural stability, as tested by Quandt's likelihood-ratio test, is also rejected.³⁰ Hence, Model 1 cannot be accepted as a representation of the data generating process (DGP). Adding non-linear terms to Model 1, as in Model 2, removes these specification problems.³¹ Additional terms, as added to Models 3 and 4, do not significantly improve the model fit, with the Schwarz Bayesian Criterion (SBC) and the Hannan-Quinn Criterion (HQC) statistics becoming worse. Only the Akaike Information Criterion (AIC) improves. Tests for correct functional form, homoskedasticity, normality, as well as the absence of ARCH effects and of structural change do not indicate any problems with Models 2, 3, and 4. While there is no problem with

³⁰ This test statistic is not shown in Table 4. The *F-value* is 5.437, which exceeds the critical value of 4.53. There is no statistical evidence indicating a structural break in Models 2, 3, and 4.

³¹ All models are tested also for ARCH effects. This is somewhat uncommon for non-financial data. However, as recently suggested by Hamilton (2007), there is strong evidence that ignoring ARCH effects can induce spurious results in typical macroeconomic models.

autocorrelation in Models 2 and 3, there appears to be higher-order autocorrelation present in Model 4.

Table 5 provides additional evidence that Models 2, 3, and 4 capture the DGP: no deterministic or stochastic trend remains, as the estimated variances of the level and slope components are zero. By contrast, the OLS specification of Model 1 does not fully capture the DGP. This model contains an underlying stochastic or deterministic trend when estimated as an UCM (Figure 15). Adding a stochastic trend to Model 1 effectively removes the statistical problems associated with the equivalent OLS model. Table 5 reveals that the hyper-parameters of Model 1 (i.e. the variances of the stochastic level and slope components) are relatively large. The hyper-parameter for the slope is different from zero. The size of the variance of the level is, by contrast, approximately zero. However, there exists a sizable stochastic trend (Figure 15). It identifies a strong increase in the divorce rate from the mid-1960s to the mid-1970s. It is precisely this increase in the divorce rate that has not been explained so far in the literature. The unobserved component version of Model 1 does not explain it either. It can capture it, but only as an unobserved and, hence, unexplained component. By contrast, Model 2 does not give rise to such an unobserved component. It captures the DGP without the need for an underlying stochastic trend to capture movements in the divorce rate that are not explained by the model's right-hand-side variables. If the UCM model excludes the variables *Vietnam* and *pill*, the resulting stochastic trend is substantial (Figure 16). As such, these covariates appear to be strong predictors of the divorce rate.

In Model 2 of Table 4, the variables *Vietnam* and *pill* are both highly statistically significant and positive. The coefficients of *fem_ratio*, *fem_ratio*², *ygrowth*, and *ygrowth*² are also statistically significant. They indicate that the divorce rate rises at a decreasing rate with an increase in *fem_ratio* or *ygrowth*. Inflation is also statistically significant and positive, as found in recent work by Nunley (2008).

The impact of *fem_ratio* and *ygrowth* is rather different when it is evaluated in terms of marginal effects at the mean values of the variables. As shown in Table 6, the short-run implied marginal effect of *fem_ratio* is negative and statistically significant, while the short-run effect of *ygrowth* remains positive but is only marginally statistically significant.³² Table 6 also provides the long-run implied marginal effects and elasticities of the models presented in Table 4. Again, using Model 2 from Table 6 as the point of reference, the long-run marginal effects for *fem_ratio*, *pill*, *ygrowth*, and *Vietnam* are substantially larger than the short-run effects. However, the long-run effect of *ygrowth* is only marginally statistically significant, while there appears to be consistent short-run effects. The long-run effects of *fem_ratio*, *pill*, and *Vietnam* are each statistically significant at the one-percent level. The former has a negative effect, while the latter two variables have positive effects on the divorce rate.

In Model 4 of Table 4, we allow for asymmetric effects of economic growth inside and outside of recessionary periods on the divorce rate. We also include an explicit measure for the potential effects of the Korean War. While this model is not chosen as

³² We only present the implied short-run marginal effects for the variables *fem_ratio* and *ygrowth* because only these variables contain more than one term in three of the four single-equation models (Table 4). However, we report the short-run implied elasticities for all variables considered in Table 6.

the preferred specification, it does fit the data reasonably well. Once we allow economic growth to have a different impact on the divorce rate in recessionary periods than in times of positive growth, the effects of $ygrowth$ and $ygrowth^2$ are no longer statistically significant. Economic growth in recessionary periods ($das*ygrowth$) has a statistically significant, positive effect on the divorce rate. Overall, however, the short-run and long-run marginal effects and elasticities for Model 4 in Table 6 are largely consistent with those from Model 2.

Next, we extend the sample from 1949 back to 1929. Table 7 shows least-squares estimates for the full sample period, 1929-2006. Extending the sample in this way necessitates the inclusion of variables for WWII. In each of the models, there are some statistical specification problems. For this reason, we report p-values that are based on a heteroskedasticity and autocorrelation robust covariance matrix (HAC). The Quandt likelihood-ratio test indicates a structural break around 1949, which confirms the selection of 1949 as the starting point for the calculations shown in Tables 4, 5, and 6.³³ Structural change appears associated in particular with the coefficients of the macroeconomic variables. For example, if one compares Tables 4 and 7, *inflation* is positive, but no longer statistically significant once the sample is extended back to 1929. Also, *ygrowth* has the opposite directional effect in the longer sample.

We also estimate UCM models for the time period 1929 to 2006. Table 8 shows that no unobserved stochastic trend is present, regardless of the explanatory variables

³³ Again, the Quandt likelihood-ratio test is not shown in Table 7. In each model, the Quandt likelihood-ratio test, which has no structural break as the null, is rejected at the one-percent level, with each model indicating a break in 1949. For our reference model (i.e. Model 7), the *F-statistic* is 3.969 and the one-percent critical value is 3.57.

included. They each appear to capture the DGP without the need for a stochastic or deterministic trend component. Because the AIC, SBC, and HQC improve as additional covariates enter the model, we select Model 7 as the preferred model specification.

The estimates of Model 7 are largely consistent with the results of Table 4 for the variables *fem_ratio*, *pill*, and *Vietnam*. However, economic growth in a recessionary period leads to a decrease in the divorce rate according to Model 7, while the opposite response is recorded for Model 4 of Tables 4 and 6. The statistically significant, positive and negative effects of *WWII* and *WWII*² indicate that the divorce rate responds nonlinearly to an increase in the variable *WWII*. The variable *Korea* has no statistically significant effect in Model 7, which is consistent with previous work (South 1985; Anderson and Little 1999).

Because there are multiple terms for *fem_ratio*, *WWII*, and *ygrowth*, we present the short-run implied marginal effects of these variables in Table 9. The variable *fem_ratio* has a negative impact, and is statistically significant at the one-percent level in each specification. By contrast, the short-run implied marginal effect of *WWII* is positive and statistically significant in each of the models. Economic growth during expansionary periods is not statistically different from zero. The long-run marginal effects for *fem_ratio*, *pill*, and *Vietnam* shown in Table 9 indicate the same directional effects and similarities in terms of size and statistical significance as the long-run marginal effects presented in Table 6.

Because there may be outliers present resulting from the depression of the 1930s, the years of WWII, and its immediate aftermath, we re-estimate the models of Table 7 for the

full sample by the method of least absolute deviations (LAD) (Table 10). This approach is less sensitive to outliers in the data and, therefore, provides more robust parameter estimates than OLS. Models 8 and 9 of Table 10 provide the best fits to the data, as there are no ARCH effects present. However, we select Model 9 as the reference model because it is more parsimonious. The results shown in Table 10 are largely consistent with those in Table 7. However, *inflation* and *ygrowth* are statistically significant and positive when the model is estimated using the LAD estimator. These variables have the same sign but are statistically insignificant in the preferred specification from Table 7.

Table 11 reports the short- and long-run implied marginal effects that correspond to the estimates of Table 10. As is apparent, there is little difference between the estimates presented in implied marginal effects and elasticities shown in Tables 9 (OLS) and 11 (LAD). The effects of *fem_ratio*, *pill*, and *Vietnam* are consistent in both specifications. However, the statistical significance of *inflation* and *ygrowth* are different. The LAD estimates indicate the short-run implied marginal effect of *ygrowth* in expansionary periods is positive and statistically significant, while the OLS estimate is not statistically different from zero. The long-run effects of *inflation* and *ygrowth* have the same signs, but are only statistically significant when estimated by LAD.

The overall conclusion from the single-equation models is that increased access to the pill has a consistent, positive short- and long-run impact on the divorce rate across all models, while female participation in higher education has negative short- and long-run effects on the divorce rate. The finding of a negative relationship between *divorce* and *fem_ratio* is opposite to that of previous research, which typically identifies a positive

relationship between divorce and the FLFPR, or its proxy, female participation in higher education. For the short sample period extending from 1949 to 2006, we estimate that a 0.1 increase in *fem_ratio* decreases the divorce rate by 2.12 in the long run. For the longer sample from 1929 to 2006, the equivalent decrease is estimated to be at 1.73 divorces. The persistent and positive effect found for increased access to the pill differs from the findings of Goldin and Katz (2002). However, our results are consistent with those of Smith (1997). According to our estimates, a 0.1 increase in the percentage of the population affected by increased access to the pill leads to an increase in divorces per capita by about 0.4 in the long run, regardless of which sample period is used.

The econometric evidence strongly supports the idea that the Vietnam War served as a catalyst of major social changes, which also involved divorce rates. The long-run marginal effect of the *Vietnam* variable on the divorce rate is significant at conventional levels of statistical significance across all models. Inflation is only statistically significant in the models using data from 1949 to 2006, while we find different statistical evidence when the models are estimated by OLS and LAD. The former indicates no effect, while the latter indicates both short- and long-run positive effects on divorces of increases in the inflation rate. The effects of economic growth on divorce are also different in the two sample periods, with positive short- and long-run effects for the shorter sample. For the longer sample, economic growth during recessions is statistically significant using both OLS and LAD estimators. The effect of economic growth during expansionary periods is positive, but not statistically significant when estimated by OLS.

However, when the model is estimated by LAD, economic growth remains positive but is statistically significant at conventional levels.

In the remaining analysis, we consider the possibility that divorce and female participation in higher education are jointly determined or endogenous. Our analysis thus far has used variations of the autoregressive-distributed lag (ADL) model, which requires that all right-hand-side variables be at least weakly exogenous. Although there is little research indicating endogeneity of female participation in higher education, such literature exists for the FLFPR. In fact, Bremmer and Kesselring (2004) treat the divorce rate and the FLFPR as jointly endogenous variables in their empirical work and identify a positive long-run relationship between the divorce rate and the FLFPR.³⁴

Table 12 presents the cointegrating vectors and the underlying tests for cointegration based on the Johansen/Juselius vector autoregression (VAR) approach. We use the AIC, SBC, and HQC to select the lag length of the underlying VAR. The cointegration rank tests suggest exactly one cointegrating vector regardless of variations in the specification. Model 4 is chosen as the preferred model because it has fewer statistical problems than the other models. The Ljung-Box Q-statistic rejects the null of no higher-order autocorrelation in Models 2, 3, 5, and 6, but is not statistically significant in Model 4. In addition to the two endogenous variables, *divorce* and *fem_ratio*, we also include an exogenous variable, *pill*, in cointegration space.³⁵ We also include several exogenous variables in the vector error correction model (VECM), in particular the war variables

³⁴ They also treat fertility rates and female earnings as endogenous variables, thereby modeling each simultaneously allowing each variable to affect the others.

³⁵ We also included other variables, such as *Vietnam* and *WWII*, and found no statistically significant relationship between *divorce* and the two variables.

WWII, *Korea*, *Vietnam*, and two observation-specific dummy variables, *d41* and *d47*. They do not have a very significant impact on the resulting cointegrating vector. More importantly, their presence or absence does not affect that exactly one cointegrating vector results for the two endogenous variables *divorce*, *fem_ratio*, and the exogenous variable, *pill*.³⁶

The results from the cointegrating vector (CIV) largely match the long-run effects estimated for the single-equation models. In the preferred specification from Tables 6 and 9, the long-run effects of *fem_ratio* are -21.20 and -16.01, respectively. The long-run estimate for *fem_ratio* in the CIV is -22.42, which is somewhat larger but overall is consistent with the single-equation, long-run effects. Therefore, a 0.1 increase in *fem_ratio* translates into 2.24 fewer divorces per capita in the long run. The long-run effects for *pill* in Table 6 and 9 are 4.43 and 3.80, respectively, while the long-run effect is 6.59 in the CIV from Table 12.

Although the results for the CIV are somewhat larger than those from the single-equation models, the long-run effects for increased access to the pill in each model are positive and statistically significant at conventional levels. However, without including the nonlinear term for *fem_ratio* in the single-equation models, there is no guarantee that the short- and long-run effects would be negative. Therefore, finding negative effects for *fem_ratio* in the single-equation models appears to hinge on the inclusion of a nonlinear term.

³⁶Adding any of these terms changes the statistical distribution of the cointegration tests. Bootstrap simulations are conducted to approximate the correct statistical significance level for the cointegration test in each case.

Models 5 and 6 consider the long-run relationship between the divorce rate and the diffusion of unilateral divorce to the population. Although we cannot separately identify the effects of pill access and the adoption of unilateral divorce, the findings of Smith (1997) and Wolfers (2006) lend support for the pill's positive effect on the divorce rate.

Although not reported in Table 12, the VECM equations have adjustment parameters in front of the error correction terms that are different from zero at any reasonable level of statistical significance. This confirms that neither the variable *fem_ratio* nor the variable *divorce* are weakly exogenous. As a consequence, we expect the OLS regressions to underestimate the impact of the variable *fem_ratio* on the divorce rate. This is borne out by the estimates: the cointegrating vectors of Table 12 all imply a larger long-run value for the variables *fem_ratio* and *pill* when the cointegrating vector is normalized on the divorce rate.

Figure 4 highlights the problems of focusing on a time horizon that is too narrow and of relying on single-equation linear regressions. The linear regression line of Figure 4 hides two distinct negative relationships between the divorce rate and the variable *fem_ratio*, one at the lower left end of the figure and the other at the upper right end of the figure. The two negative relationships are shown separately in Figures 5 and 6. The regression coefficients on *fem_ratio* in Figures 5 and 6 (-19.1 and -11.2, respectively) are close to the estimates from the single-equation models and the cointegration relationships of Table 12. The regression coefficient for variable *pill* in Figure 9 is also close to the estimated long-run effects of *pill* from Tables 6, 9, 11, and 12. It appears that the impact

of the Vietnam War pushes the divorce rate for some years above the regression line in Figure 9, which is also the case in Figure 4.

There appears to be a distinct difference between WWII and the Vietnam War. WWII led to a temporary shift of the negative relationship between the divorce rate and *fem_ratio*. The Vietnam War together with access to the pill led to a permanent, rightward shift in the curve linking *divorce* and *fem_ratio*. We find that diffusion of the pill and unilateral divorce have quantitatively similar effects on the divorce rate. Based on statistical fit, we are unable to disentangle which of these factors is driving the rise in divorce during the 1960s and 1970s from the time-series data for the U.S. But the findings of other studies shed some light on the causal relationships. For example, Wolfers (2006), using state-level panel data from the U.S., concludes that the adoption of unilateral divorce led to a small, transitory rise in the U.S. divorce rates.³⁷ Comparing the U.S. divorce rate with those in European countries, along with the careful econometric evidence collected by Smith (1997) for the pill and Wolfers (2006) for divorce-law changes, it seems reasonable to conclude that the pill is more likely the key factor for the rise in divorce rates in the 1960s and early 1970s than changes in divorce laws. In fact, visual inspection of time plots of the divorce rates of most European countries indicate a rise in divorce rates prior to the implementation of either no-fault or unilateral divorce laws (e.g., Gonzalez and Viitanen 2006, Figure 2).³⁸ Smith (1997) finds evidence that

³⁷ Others find similar results (e.g., Friedberg 1998), but do not estimate the full adjustment process as in Wolfers (2006).

³⁸ Go to <http://repec.iza.org/RePEc/Discussionpaper/dp2023.pdf> to access the manuscript. Although Gonzalez and Viitanen (2006) find a persistent, positive effect of liberalization of divorce laws on divorce rates in European countries, Smith (2002) suggests that the strictness of divorce laws and divorce rates may

access to the pill increased the divorce rate in England and Wales but concludes that reforms allowing easier access to divorce led to a temporary rise in divorce rates, with no evidence of a long-run relationship.

4.5.2. Sensitivity Analysis

In this section, we check our results to a number of alternative specifications. First, we examine whether the results are materially affected by having employed the variable female participation in higher education in lieu of the variable female participation in the labor force (FLFPR). The single-equation estimates analogous to those of Table 4 but with FLFPR substituting for female participation in higher education are presented in Table A1. The coefficient estimates for the female labor-force participation rate (*flfpr*) are very similar to those of the variable female participation in higher education as reported in Table 4. The coefficients of the other covariates in Table A1 are also quantitatively similar to those of Table 4. Hence, no bias is apparently introduced by relying on female participation in higher education as a proxy for female labor force participation.

Next, we check whether the chosen definition of the dependent variable, divorce per 1,000 persons, leads to materially different results than the alternative divorce variable “divorces per 1,000 married couples.” For that purpose we re-estimate the models presented in Tables 4, 7, and 10 with divorces per 1,000 married couples as the dependent

be simultaneously determined. Therefore, one should perhaps interpret their findings cautiously. Passage of unilateral divorce in the U.S., on the other hand, has been argued as plausibly exogenous (See Jacob 1988).

variable. These estimates are shown in Tables A2, A3, and A4. While the estimated effects are much larger than those of Tables 4, 7, and 10, the directional effects and statistical significance are very similar to those that result for the dependent variable “divorces per 1,000 persons”. The substantially larger estimates simply result from the fact that divorces per 1,000 married couples exceed divorces per 1,000 persons by a substantial margin. Table A5 presents the resulting cointegrating equations when divorces per 1,000 married couples is used as the dependent variable. Similar to the results of Table 12, we find only one cointegrating equation, with the same directional relationship and statistical significance.

Lastly, we revisit Bremmer and Kesselring’s (2004) study by considering the potential cointegrating relationship between the divorce rate and the FLFPR for the years 1960-2001. We find the presence of two cointegrating vectors: one indicating a positive relationship and the other a negative one between the divorce rate and the FLFPR. The resulting cointegrating equations both normalized on the divorce rate per 1,000 married couples are

$$\text{divorce} = -23.460 + 0.097 * \text{flfpr}$$

$$\text{divorce} = 115.776 - 2.037 * \text{flfpr} .$$

For the first equation, the estimated effect of the FLFPR on the divorce rate is positive but small. This is the relationship identified by Bremmer and Kesselring (2004). Their estimate for the FLFPR, normalized on the divorce rate, is slightly smaller than the estimate shown in the first cointegrating equation above. With the second, the estimate remains small, but is larger than the effect found for the first cointegrating equation.

Our sensitivity analyses support the results of Section 5.1. The estimates appear to be unaffected by the measure of the divorce rate. Likewise, when using the FLFPR instead of female participation in higher education, there appears to be little difference in the estimated effects. Lastly, we can encompass the result of a positive long-run relationship between the divorce rate and the FLFPR found by Bremmer and Kesselring (2004) but we show that there are in fact two cointegrating equations for the time period 1960-2001: one indicating a positive relationship and the other a negative relationship.

4. 6. Conclusions

This study examines the evolution of the U.S. divorce rate from 1929 to 2006, with the primary aim of explaining the rise in the divorce rate during the mid-1960s to the mid-1970s. We extend previous research in a number of ways. First, following Smith (1997), we consider whether increased access to oral contraception contributed to the rise in the divorce rate. Second, we construct objective measures for the effects of WWII, the Korean War, and the Vietnam War on the divorce rate, while previous research has used time-specific, indicator variables to capture the effects of wars on divorce rates. Third and perhaps most importantly, we extend the analysis back to 1929 to allow for more variation in sample observations.

We show that previous work that has been too narrowly focused on the 1960s and 1970s necessarily identifies a positive relationship between the female labor-force participation rate or its close proxy, female participation in higher education, and the divorce rate. Examining a sample that is overweighted by the observations of the 1960s

and 1970s hides two distinct negative relationships between the divorce rate and female participation in the labor force and higher education: one before and after WWII and another from the late-1970s onwards. The sharp rise in the divorce rate from the mid-1960s to the mid-1970s is marked by increased access to oral contraception, divorce-law reform, and the Vietnam War. After incorporating the impact of these variables and extending the sample back to 1929, we identify a strong, negative relationship between female participation in higher education and the divorce rate. In particular, we find that female participation in higher education and the divorce rate are endogenous variables linked by a negative, long-run relationship. However, we find similar, albeit slightly smaller, effects when we estimate single-equation models, which treat female participation in higher education as a weakly exogenous variable.

Our econometric evidence supports the idea that access to oral contraception, divorce-law changes, and the Vietnam War shifted the divorce rate to a permanently higher level from the mid 1960s to the mid 1970s. Unfortunately, the econometric evidence from time-series data is not strong enough to identify separately the effects of access to oral contraception and divorce-law reform, as both came about almost concurrently. However, recent research by Wolfers (2006) and Smith (1997) suggest that increased access to the oral contraception in the U.S. is the more likely factor that led to the sharp increase in divorce rates in the 1960s and 1970s rather than divorce-law reforms. WWII and the Vietnam War also significantly increased the divorce rate, while no such effect can be found for the Korean War.

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TABLE 1: VARIABLE NAMES, DEFINITIONS, AND SOURCES

Name	Definition	Source
<i>divorce</i>	Number of divorces per 1,000 persons	<i>Historical Statistics of the United States</i> and U.S. Statistical Abstracts.
<i>WWII</i>	U.S. military personnel deaths due to WWII as a fraction of U.S. military deployments for WWII.	http://www.history.navy.mil/library/online/ww2_statistics.htm ; http://stinet.dtic.mil/oai/oai?verb=getRecord&metadataPrefix=html&identifier=ADA438106 ; http://www.history.army.mil/books/wwii/sp1943-44/index.htm
<i>Korea</i>	U.S. military personnel deaths as a fraction of U.S. military deployments in East Asia.	http://www.archives.gov/research/korean-war/casualty-lists
<i>Vietnam</i>	U.S. military personnel deaths due to the Vietnam War as a fraction of U.S. military deployments in East Asia.	http://www.thetruthseeker.co.uk/article.asp?ID=2703 ; http://www.heritage.org/Research/NationalSecurity/cda06-02.cfm
<i>pill</i>	Diffusion function, measured as the percentage of the U.S. population affected by increased access to oral contraceptives.	Law coding is from Bailey (2006); population data are from <i>Historical Statistics of the United States</i> and U.S. Statistical Abstracts.
<i>friedberg/gruber</i>	Diffusion function, measured as the percentage of the U.S. population affected by unilateral divorce law reform.	Law codings are from Friedberg (1998) and Gruber (2004); population data are from <i>Historical Statistics of the United States</i> and U.S. Statistical Abstracts.
<i>inflation</i>	Log of the ratio of the Consumer Price Index at period t relative to the CPI at period $t-1$	http://www.inflationdata.com/Inflation/Inflation_Rate/HistoricalInflation.aspx
<i>ygrowth</i>	Log of the ratio of U.S. Gross National Product (GNP) at period t relative to U.S. GNP at $t-1$.	http://research.stlouisfed.org/fred2/series/GNPCA?cid=106
<i>das*ygrowth</i>	Log of the ratio of U.S. Gross National Product (GNP) at period t relative to U.S. GNP at $t-1$ multiplied by a dummy variable that equals one when the economy is in a recessionary period and zero otherwise.	http://research.stlouisfed.org/fred2/series/GNPCA?cid=106
<i>fem_ratio</i>	Percentage of women enrolled in higher education relative to the total population enrolled in higher education	<i>Historical Statistics of the United States</i> and U.S. Statistical Abstracts.
$\Delta unemp$	Change in the unemployment rate, measured as the percentage of the population who is unemployed but actively pursuing employment.	http://www.bls.gov/cps/cpsaat1.pdf ; http://research.stlouisfed.org/fred2/series/UNRATE/downloaddata?cid=12

TABLE 2: BASIC SUMMARY STATISTICS

Variable	Mean	Med.	Min.	Max.	Std. Dev.	C.V.	Skew.	Kurt.
1929-2006								
<i>divorce</i>	3.753	4.000	2.100	5.300	1.088	0.290	-0.222	-1.510
<i>fem_ratio</i>	0.466	0.493	0.296	0.577	0.087	0.186	-0.264	-1.462
<i>fem_ratio</i> ²	0.224	0.243	0.087	0.333	0.079	0.353	-0.136	-1.568
<i>pill</i>	0.620	1.000	0.000	1.000	0.465	0.751	-0.462	-1.730
<i>inflation</i>	3.778	3.020	-0.950	13.580	2.938	0.778	1.357	1.774
<i>ygrowth</i>	3.406	3.507	-1.869	8.738	2.367	0.695	-0.151	-0.287
<i>ygrowth</i> ²	17.110	12.303	0.027	76.357	16.908	0.988	1.440	1.813
<i>das*ygrowth</i>	-0.094	0.000	-1.869	0.000	0.305	3.247	-4.251	19.731
Δ <i>unemp</i>	0.015	-0.246	-2.092	2.833	1.056	71.354	0.916	0.735
<i>Vietnam</i>	0.125	0.000	0.000	2.157	0.398	3.183	3.592	12.462
<i>Korea</i>	0.228	0.000	0.000	9.478	1.276	5.587	6.811	46.476
<i>WWII</i>	0.138	0.000	0.000	2.796	0.501	3.624	3.657	12.807
1929-1948								
<i>divorce</i>	2.215	1.900	1.300	4.300	0.807	0.365	1.090	0.404
<i>fem_ratio</i>	0.408	0.416	0.289	0.499	0.053	0.130	-1.055	0.839
<i>fem_ratio</i> ²	0.169	0.173	0.083	0.249	0.041	0.240	-0.684	0.617
<i>inflation</i>	2.004	2.270	-10.300	14.650	6.319	3.153	-0.099	-0.210
<i>ygrowth</i>	3.849	5.240	-13.048	18.431	9.584	2.490	-0.192	-1.048
<i>ygrowth</i> ²	101.83	73.48	0.623	339.69	102.57	1.007	1.094	0.099
<i>das*ygrowth</i>	-2.410	0.000	-13.048	0.000	4.154	1.724	-1.563	0.969
Δ <i>unemp</i>	0.032	-0.700	-4.960	7.710	3.841	119.65	0.726	-0.572
<i>WWII</i>	0.539	0.000	0.000	2.796	0.889	1.648	1.231	0.112
1949-2006								
<i>divorce</i>	3.737	4.000	2.100	5.300	1.086	0.291	-0.186	-1.526
<i>fem_ratio</i>	0.463	0.487	0.289	0.577	0.089	0.193	-0.274	-1.410
<i>fem_ratio</i> ²	0.222	0.237	0.083	0.333	0.080	0.363	-0.130	-1.547
<i>pill</i>	0.609	1.000	0.000	1.000	0.468	0.768	-0.418	-1.771
<i>inflation</i>	3.845	3.030	-0.950	13.580	2.958	0.769	1.280	1.501
<i>ygrowth</i>	3.424	3.54528	-1.869	8.738	2.351	0.687	-0.173	-0.253
<i>ygrowth</i> ²	17.156	12.5690	0.027	76.357	16.765	0.977	1.443	1.875
<i>das*ygrowth</i>	-0.092	0.000	-1.869	0.000	0.302	3.277	-4.293	20.138
Δ <i>unemp</i>	0.018	-0.242	-2.092	2.833	1.047	57.919	0.914	0.783
<i>Vietnam</i>	0.165	0.000	0.000	2.157	0.451	2.728	3.017	8.349
<i>Korea</i>	0.224	0.000	0.000	9.478	1.265	5.636	6.873	47.358

**TABLE 3: UNIT ROOT AND STATIONARITY TESTS
(1929-2007 OR MAXIMUM AVAILABLE)**

Variable	ADF – H0: I(1)			KPSS – H0: I(0)	
	Constant	Constant with Trend	Lag Order	Trend	No Trend
Continuous Variables:					
<i>divorce</i>	0.445	0.852	1	1.441	2.797
<i>fem_ratio</i>	0.943	0.576	4	0.283	1.245
<i>fem_ratio</i> ²	0.941	0.542	3	1.586	1.586
<i>pill</i>	0.790	0.563	1	1.785	2.500
<i>friedberg</i>	0.860	0.573	4	0.204	1.427
<i>gruber</i>	0.871	0.592	4	0.261	1.455
<i>mechoulan</i>	0.869	0.622	5	0.193	1.226
<i>inflation</i>	0.001	0.018	2	0.369	0.353
<i>ygrowth</i>	0.000	0.000	3	0.063	0.063
<i>ygrowth</i> ²	0.088	0.000	3/2	0.843	0.843
<i>das*ygrowth</i>	0.000	0.000	0	0.624	1.141
Δ <i>unemp</i>	0.000	0.000	3	0.064	0.064
Quasi Dummy Variables:					
<i>WWII</i>	0.005	0.010	2	0.329	0.401
<i>WWII</i> ²	0.006	0.015	2	0.301	0.358
<i>Korea</i>	0.000	0.000	0	0.167	0.240
<i>Vietnam</i>	0.001	0.007	1	0.281	0.281

Notes: ADF stands for the Augmented Dickey-Fuller test; the statistics for the ADF test are p-values. The ADF tests whether the variables follows a unit-root process, with unit root as the null hypothesis. The null hypothesis for the KPSS test is stationarity. The critical values for the KPSS test are 0.347 (10%), 0.463 (5%), 0.574 (2.5%), and 0.739 (1%). The column denoting lag order represents the number of lags used for both the ADF and KPSS tests.

TABLE 4: OLS ESTIMATES, 1949-2006

	Model 1		Model 2		Model 3		Model 4	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
constant	0.8753	0.000	-0.7614	0.236	-0.7529	0.237	-0.8318	0.234
<i>divorce</i> (-1)	0.8761	0.000	0.8705	0.000	0.8732	0.000	0.8881	0.000
<i>fem_ratio</i>	-1.9070	0.000	5.8624	0.045	5.3303	0.066	5.8633	0.072
<i>fem_ratio</i> ²			-8.6083	0.006	-7.8313	0.011	-8.3934	0.016
<i>pill</i>	0.5912	0.000	0.5738	0.000	0.5453	0.000	0.5238	0.000
<i>inflation</i>	0.0143	0.019	0.0101	0.073	0.0125	0.032	0.0091	0.122
<i>ygrowth</i>	0.0112	0.042	0.0343	0.011	0.0562	0.005	0.0145	0.566
<i>ygrowth</i> ²			-0.0040	0.035	-0.0049	0.014	0.0002	0.938
<i>das*ygrowth</i>							0.1624	0.018
Δ <i>unemp</i>					0.0393	0.130	0.0443	0.086
<i>Vietnam</i> (-1)	0.1850	0.000	0.1319	0.000	0.1284	0.000	0.1247	0.000
<i>Korea</i>							-0.0120	0.334
Unadjusted R ²	0.9935		0.9952		0.9955		0.9960	
Adjusted R ²	0.9927		0.9947		0.9946		0.9950	
Log-likelihood	59.408		68.305		69.704		73.351	
AIC	-104.815		-118.610		-119.408		-122.703	
SBC	-90.392		-100.066		-98.804		-97.977	
HQC	-99.173		-111.386		-111.382		-113.072	
RESET		0.001		0.179		0.149		0.236
Homoskedasticity		0.098		0.530		0.303		0.240
Normality		0.951		0.551		0.293		0.363
Autocorrelation LM(1)		0.876		0.280		0.275		0.040
Autocorrelation LM(2)		0.853		0.281		0.228		0.033
Autocorrelation LM(3)		0.652		0.323		0.408		0.121
ARCH (1)		0.312		0.161		0.260		0.484
ARCH (2)		0.462		0.290		0.525		0.655
ARCH (3)		0.756		0.456		0.710		0.774
Harvey-Collier (cusum)		0.019		0.706		0.492		0.623

Notes: (-1) denotes a lag order of one. AIC stands for the Akaike Information Criterion, SBC for the Schwarz-Bayesian Criterion, HQC for the Hannan-Quinn Criterion; RESET is Ramsey's test for correct functional form; Homoskedasticity is White's test; Normality is a test for normality of the residuals; the null of no autocorrelation at various lag lengths is tested with the Breusch-Godfrey test; ARCH tests the null of no relationship between the current error variance and its past values; Quandt LR tests for the lack of structural breaks (Stock and Watson 2007); and Harvey-Collier tests parameter stability using cumulated recursive residuals.

TABLE 5: MODELS FROM TABLE 4 ESTIMATED AS UNOBSERVED COMPONENT MODELS

	Model 1		Model 2		Model 3		Model 4	
	Estimated Variance	Q Ratio	Estimated Variance	Q Ratio	Estimated Variance	Q Ratio	Estimated Variance	Q Ratio
σ_{ε}^2	0.0065	1.0000	0.0065	1.0000	0.0062	1.0000	0.0056	1.0000
σ_{η}^2	0.0000	0.0000	0.0000	0.0000	0.0001	0.0129	0.0000	0.0000
σ_{ξ}^2	0.0000	0.0009	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
μ	-1.006	0.265	-2.135	0.057	-2.177	0.050	-3.401	0.018
β	-0.019	0.030	-0.006	0.506	-0.007	0.397	-0.012	0.140

Notes: Only parameters that relate to stochastic or deterministic trends are provided. This table shows the estimated variances for the level and slope components, and shows the coefficient estimates for the level and slope components once they are restricted to be fixed. The Q Ratio provides the ratio of the estimated variances relative to the irregular component. To test this for the presence of a deterministic trend, we fix the level and slope components and then test for whether they are different from zero. The p-values for the level and slope coefficients in Models 2, 3, and 4 that the models collapse to OLS, which implies that these models fully captures the trend in the divorce rate.

TABLE 6: IMPLIED MARGINAL EFFECTS AND ELASTICITIES FOR TABLE 4

	Model 1		Model 2		Model 3		Model 4	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
Short-Run Implied Marginal Effects:								
<i>fem_ratio</i>	-1.9070	0.000	-2.7460	0.000	-2.5010	0.000	-2.5301	0.000
<i>ygrowth</i> (if positive)	0.0112	0.042	0.0302	0.054	0.0513	0.047	0.0148	0.602
<i>ygrowth</i> (if negative)	0.0112	0.042	0.0302	0.054	0.0513	0.047	0.1772	0.000
Short-Run Implied Elasticities:								
<i>fem_ratio</i>	-0.2363		-0.3402		-0.3099		-0.3135	
<i>pill</i>	0.0963		0.0935		0.0889		0.0854	
<i>inflation</i>	0.0147		0.0104		0.0129		0.0094	
<i>ygrowth</i> (if positive)	0.0103		0.0277		0.0470		0.0136	
<i>ygrowth</i> (if negative)	0.0103		0.0277		0.0470		0.1624	
<i>Vietnam</i>	0.0082		0.0058		0.0057		0.0055	
$\Delta unemp$					0.0002		0.0002	
<i>Korea</i>							-0.0007	
Long-Run Implied Marginal Effects:								
<i>fem_ratio</i>	-15.390	0.002	-21.201	0.000	-19.725	0.000	-22.600	0.000
<i>pill</i>	4.7714	0.000	4.4299	0.000	4.3010	0.000	4.6794	0.000
<i>inflation</i>	0.1151	0.084	0.0782	0.111	0.0987	0.066	0.0809	0.251
<i>ygrowth</i> (if positive)	0.0905	0.098	0.2335	0.053	0.4044	0.079	0.1319	0.607
<i>ygrowth</i> (if negative)	0.0905	0.098	0.2335	0.053	0.4044	0.079	1.5825	0.028
<i>Vietnam</i>	1.4933	0.005	1.0185	0.005	1.0124	0.008	1.1140	0.007
$\Delta unemp$					0.3097	0.232	0.3953	0.203
<i>Korea</i>							-0.1069	0.316
Long-Run Implied Elasticities:								
<i>fem_ratio</i>	-1.9068		-2.6267		-2.4439		-2.8001	
<i>pill</i>	0.7776		0.7219		0.7009		0.7626	
<i>inflation</i>	0.1184		0.0805		0.1016		0.0832	
<i>ygrowth</i> (if positive)	0.0829		0.2139		0.3705		0.1209	
<i>ygrowth</i> (if negative)	0.0829		0.2139		0.3705		1.4500	
<i>Vietnam</i>	0.0659		0.0450		0.0447		0.0492	
$\Delta unemp$					0.0015		0.0019	
<i>Korea</i>							-0.0064	

Notes: We only report the short-term implied marginal effects for the variables that have multiple terms. Short-run implied elasticities, long-run implied marginal effects, and long-run implied elasticities are reported for all variables.

TABLE 7: LEAST SQUARES ESTIMATES, 1929-2006, HAC COVARIANCE MATRIX

Variable	Model 5		Model 6		Model 7	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
constant	-2.1561	0.000	-2.0441	0.000	-1.7393	0.002
<i>divorce</i> (-1)	0.7931	0.000	0.7760	0.000	0.8275	0.000
<i>fem_ratio</i>	13.9000	0.000	12.7335	0.000	10.7761	0.000
<i>fem_ratio</i> ²	-17.8613	0.000	-16.3184	0.000	-13.7624	0.000
<i>pill</i>	0.8529	0.002	0.8516	0.001	0.6356	0.000
<i>inflation</i>	-0.0037	0.634	0.0043	0.625	0.0091	0.128
<i>ygrowth</i>	-0.0153	0.153	0.0296	0.066	0.0163	0.293
<i>ygrowth</i> ²	-0.0013	0.002	-0.0038	0.000	-0.0030	0.001
<i>das*ygrowth</i>			-0.0790	0.020	-0.0660	0.007
Δ <i>unemp</i>	-0.0479	0.016	-0.0269	0.046	-0.0389	0.033
<i>WWII</i>	0.9219	0.000	1.0776	0.000	0.8789	0.000
<i>WWII</i> ²	-0.2581	0.000	-0.3194	0.000	-0.2534	0.000
<i>Vietnam</i> (-1)	0.1370	0.000	0.1395	0.000	0.1292	0.000
<i>Korea</i>	0.0227	0.002	0.0167	0.014	0.0111	0.381
<i>d47</i>					-0.5746	0.001
Unadjusted R ²	0.9910		0.9921		0.9938	
Adjusted R ²	0.9893		0.9905		0.9924	
Log-likelihood	57.497		62.721		71.876	
AIC	-88.995		-97.442		-113.753	
SBC	-58.525		-64.629		-78.596	
HQC	-76.807		-84.317		-99.690	
RESET		0.513		0.413		0.169
Homoskedasticity		0.000		0.000		0.006
Normality		0.003		0.130		0.048
Autocorrelation LM(1)		0.045		0.098		0.018
Autocorrelation LM(2)		0.075		0.109		0.071
Autocorrelation LM(3)		0.029		0.018		0.018
ARCH (1)		0.001		0.002		0.236
ARCH (2)		0.003		0.005		0.404
ARCH (3)		0.009		0.010		0.626
Harvey-Collier (CUSUM)		0.686		0.309		0.530

Notes: (-1) denotes a lag order of one. AIC stands for the Akaike Information Criterion, SBC for the Schwarz-Bayesian Criterion, HQC for the Hannan-Quinn Criterion; RESET is Ramsey's test for correct functional form; Homoskedasticity is White's test; Normality is a test for normality of the residuals; the null of no autocorrelation at various lag lengths is tested with the Breusch-Godfrey test; ARCH tests the null of no relationship between the current error variance and its past values; Quandt LR tests for the lack of structural breaks (Stock and Watson 2007); and Harvey-Collier tests parameter stability using cumulated recursive residuals.

TABLE 8: MODELS FROM TABLE 7 ESTIMATED AS UNOBSERVED COMPONENT MODELS

	Model 5		Model 6		Model 7	
	Estimated Variance	Q Ratio	Estimated Variance	Q Ratio	Estimated Variance	Q Ratio
σ_ε^2	0.0154	1.0000	0.0123	1.0000	0.0114	1.0000
σ_η^2	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
σ_ξ^2	0.0000	0.0003	0.0000	0.0017	0.0000	0.0000
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
μ	-2.425	0.001	-2.379	0.001	-1.873	0.003
β	0.002	0.409	0.002	0.271	0.001	0.649

Notes: Only parameters that relate to stochastic or deterministic trends are provided. This table shows the estimated variances for the level and slope components, and shows the coefficient estimates for the level and slope components once they are restricted to be fixed. The Q Ratio provides the ratio of the estimated variances relative to the irregular component. To test this for the presence of a deterministic trend, we fix the level and slope components and then test for whether they are different from zero. The p-values for the level and slope coefficients in Models 5, 6, and 7 that the models collapse to OLS, which implies that these models fully captures the trend in the divorce rate.

TABLE 9: IMPLIED MARGINAL EFFECTS AND ELASTICITIES FOR TABLE 7

	Model 5		Model 6		Model 7	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
Short-Run Implied Marginal Effects:						
<i>fem_ratio</i>	-3.9613	0.000	-3.5849	0.000	-2.9863	0.000
<i>ygrowth</i> (if negative)	-0.0166	0.264	-0.0532	0.032	-0.0527	0.007
<i>ygrowth</i> (if positive)	-0.0166	0.264	0.0258	0.178	0.0133	0.431
<i>WWII</i>	0.6637	0.000	0.7582	0.000	0.6255	0.000
Short-Run Implied Elasticities:						
<i>fem_ratio</i>	-0.4919		-0.4451		-0.3708	
<i>pill</i>	0.1409		0.1407		0.1050	
<i>inflation</i>	-0.0037		0.0043		0.0092	
<i>ygrowth</i> (if negative)	-0.0151		0.0234		0.0121	
<i>ygrowth</i> (if positive)	-0.0151		-0.0483		-0.0478	
<i>Vietnam</i>	0.0046		0.0046		0.0043	
<i>Korea</i>	0.0014		0.0010		0.0007	
<i>WWII</i>	0.0244		0.0279		0.0230	
Δ <i>unemp</i>	-0.0002		-0.0001		-0.0002	
Long-Run Implied Marginal Effects:						
<i>fem_ratio</i>	-19.147	0.000	-16.006	0.000	-17.313	0.000
<i>pill</i>	4.1224	0.000	3.8021	0.000	3.6850	0.000
<i>inflation</i>	-0.0178	0.582	0.0192	0.590	0.0525	0.212
<i>ygrowth</i> (if negative)	-0.0803	0.275	-0.2374	0.026	-0.3055	0.008
<i>ygrowth</i> (if positive)	-0.0803	0.275	0.1154	0.141	0.0769	0.434
<i>Vietnam</i>	0.6623	0.002	0.6229	0.003	0.7490	0.005
<i>Korea</i>	0.0198	0.013	0.0744	0.117	0.0641	0.382
<i>WWII</i>	3.2083	0.000	3.3854	0.000	3.6264	0.000
Δ <i>unemp</i>	-0.2316	0.100	-0.1199	0.246	-0.2255	0.053
Long-Run Implied Elasticities:						
<i>fem_ratio</i>	-2.3774		-1.9874		-2.1497	
<i>pill</i>	0.6810		0.6281		0.6088	
<i>inflation</i>	-0.0179		0.0193		0.0528	
<i>ygrowth</i> (if negative)	-0.0729		-0.2155		-0.2773	
<i>ygrowth</i> (if positive)	-0.0729		0.1047		0.0698	
<i>Vietnam</i>	0.0221		0.0207		0.0249	
<i>Korea</i>	0.0012		0.0045		0.0039	
<i>WWII</i>	0.1180		0.1245		0.1333	
Δ <i>unemp</i>	-0.0009		-0.0005		-0.0009	

Notes: We only report the short-term implied marginal effects for the variables that have multiple terms. Short-run implied elasticities, long-run implied marginal effects, and long-run implied elasticities are reported for all variables considered.

TABLE 10: LEAST ABSOLUTE DEVIATIONS ESTIMATES, 1929-2006

Variable	Model 8		Model 9		Model 10	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
constant	-0.9690	0.038	-0.8806	0.040	-1.3401	0.001
<i>divorce</i> (-1)	0.8485	0.000	0.8404	0.000	0.8499	0.000
<i>fem_ratio</i>	6.8833	0.001	6.3547	0.001	8.6233	0.000
<i>fem_ratio</i> ²	-9.5605	0.000	-8.8390	0.000	-11.7249	0.000
<i>pill</i>	0.5579	0.000	0.5609	0.000	0.6069	0.000
<i>inflation</i>	0.0176	0.000	0.0175	0.000	0.0106	0.021
<i>ygrowth</i>	0.0324	0.008	0.0432	0.000	0.0431	0.000
<i>ygrowth</i> ²	-0.0035	0.000	-0.0041	0.000	-0.0039	0.000
<i>das*ygrowth</i>	-0.0842	0.000	-0.0954	0.000	-0.0879	0.000
Δ <i>unemp</i>	-0.0136	0.373				
<i>WWII</i>	0.8805	0.000	0.9229	0.000	0.9486	0.000
<i>WWII</i> ²	-0.2554	0.000	-0.2731	0.000	-0.2816	0.000
<i>Vietnam</i> (-1)	0.1111	0.001	0.1130	0.000	0.1202	0.001
<i>Korea</i>	0.0013	0.910				
<i>d47</i>	-0.7747	0.000	-0.7278	0.000		
Sum of Absolute Residuals	5.1071		5.1719		5.8638	
ARCH (1)		0.828		0.787		0.000
ARCH (2)		0.957		0.966		0.000
ARCH (3)		0.988		0.992		0.000

Notes: (-1) denotes a lag order of one. Δ is the first-difference operator. *d47* is an indicator variables for year 1947. We begin with the final model from Table 7, and test the model down to its most parsimonious form. We only report tests for ARCH effects, as the LAD estimator allows for outliers in the data. Model 9 provides the best fit to the data, as it is more parsimonious than Model 8 and has no ARCH effects present.

TABLE 11: IMPLIED MARGINAL EFFECTS AND ELASTICITIES FOR TABLE 10

	Model 8		Model 9		Model 10	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
Short-Run Implied Marginal Effects:						
<i>fem_ratio</i>	-2.6772	0.001	-2.4843	0.000	-3.1016	0.000
<i>ygrowth</i> (if negative)	-0.0553	0.011	-0.0563	0.000	-0.0486	0.002
<i>ygrowth</i> (if positive)	0.0289	0.175	0.0391	0.000	0.3925	0.001
<i>WWII</i>	0.6231	0.000	0.6498	0.000	0.6670	0.000
Short-Run Implied Elasticities:						
<i>fem_ratio</i>	-0.3324		-0.3085		-0.3851	
<i>pill</i>	0.0922		0.0927		0.1003	
<i>inflation</i>	0.0177		0.0176		0.0107	
<i>ygrowth</i> (if negative)	-0.0502		-0.0511		-0.0441	
<i>ygrowth</i> (if positive)	0.0262		0.0355		0.3562	
<i>Vietnam</i>	0.0037		0.0038		0.0040	
<i>Korea</i>	0.0001		0.0000		0.0000	
<i>WWII</i>	0.0229		0.0239		0.0245	
Δ <i>unemp</i>	-0.0001		0.0000		0.0000	
Long-Run Implied Marginal Effects:						
<i>fem_ratio</i>	-17.666	0.036	-15.567	0.001	-20.660	0.003
<i>pill</i>	3.6811	0.000	3.5148	0.000	4.0426	0.000
<i>inflation</i>	0.1162	0.074	0.1094	0.006	0.0707	0.071
<i>ygrowth</i> (if negative)	-0.3651	0.048	-0.3528	0.002	-0.3240	0.015
<i>ygrowth</i> (if positive)	0.1904	0.167	0.2451	0.004	0.2615	0.016
<i>Vietnam</i>	0.7334	0.067	0.7079	0.006	0.8004	0.016
<i>Korea</i>	0.0088	0.893				
<i>WWII</i>	4.1248	0.005	4.0718	0.000	4.4431	0.001
Δ <i>unemp</i>	-0.0895	0.607				
Long-Run Implied Elasticities:						
<i>fem_ratio</i>	-2.1935		-1.9329		-2.5653	
<i>pill</i>	0.6081		0.5806		0.6678	
<i>inflation</i>	0.1170		0.1101		0.0712	
<i>ygrowth</i> (if negative)	-0.3313		-0.3202		-0.2940	
<i>ygrowth</i> (if positive)	0.1728		0.2224		0.2373	
<i>Vietnam</i>	0.0244		0.0236		0.0266	
<i>Korea</i>	0.0005					
<i>WWII</i>	0.1517		0.1497		0.1634	
Δ <i>unemp</i>	-0.0004					

Notes: We only report the short-term implied marginal effects for the variables that have multiple terms. Short-run implied elasticities, long-run implied marginal effects, and long-run implied elasticities are reported for all variables considered.

TABLE 12: COINTEGRATION AND VECM RESULTS, 1929 TO 2006

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
<i>Implicit CIV Normalized on Divorce</i>						
constant	-11.450 (-10.01)	-11.333 (-9.97)	-11.263 (-9.49)	-10.685 (-13.22)	-14.627 (-9.96)	-14.172 (-11.03)
fem_ratio	24.839 (8.232)	24.908 (8.02)	24.0230 (7.64)	22.710 (10.68)	33.865 (8.68)	33.167 (9.49)
pill	-7.156 (-13.21)	-6.673 (-12.50)	-6.7230 (-12.14)	-6.589 (-17.58)		
gruber					-15.847 (-12.93)	
friedberg						-15.857 (-14.11)
<i>Unrestricted Exogenous Variables</i>		<i>WWII</i> <i>Vietnam</i> <i>Korea</i>	<i>WWII</i> <i>WWII</i> ² <i>Vietnam</i> <i>Korea</i>	<i>WWII</i> <i>WWII</i> ² <i>Vietnam</i> <i>Korea</i>	<i>WWII</i> <i>WWII</i> ² <i>Vietnam</i> <i>Korea</i>	<i>WWII</i> <i>WWII</i> ² <i>Vietnam</i> <i>Korea</i>
				<i>d41</i> <i>d46</i>	<i>d41</i> <i>d46</i>	<i>d41</i> <i>d46</i>
<i>Cointegration Trace Test</i>						
p-values:						
null of rank = 0	0.000	0.006	0.016	0.000	0.000	0.000
null of rank = 1	0.206	0.347	0.387	0.249	0.318	0.345
<i>Statistics and Specification Tests of VECM</i>						
Log-likelihood	477.164	489.628	504.619	577.257	576.638	576.908
SBC	-11.190	-11.240	-11.533	-13.288	-13.271	-13.278
HQC	-11.832	-11.938	-12.269	-14.100	-14.083	-14.090
R ² divorce equation	0.690	0.685	0.723	0.821	0.858	0.854
R ² second equation	0.621	0.676	0.706	0.956	0.943	0.945
P-values of system tests						
ARCH (1)	0.000	0.000	0.013	0.084	0.047	0.062
ARCH (2)	0.000	0.000	0.209	0.241	0.381	0.410
Ljung Box Q (18)	0.133	0.030	0.011	0.071	0.000	0.005
Autocorrelation (1)	0.134	0.015	0.358	0.001	0.026	0.015
Autocorrelation (2)	0.693	0.857	0.003	0.836	0.285	0.397
Normality	0.000	0.000	0.000	0.166	0.065	0.225

Note: T-values are in parenthesis. 5 lags are used. The p-values of the cointegration rank test are based on simulated values with 2,500 replications and random walks of length 400. The effective sample is 1934 to 2006. Calculations are done in CATS in RATS, version 2 (Dennis et al. 2005).

FIGURE 1—DIVORCE RATE PER CAPITA IN THE U.S.

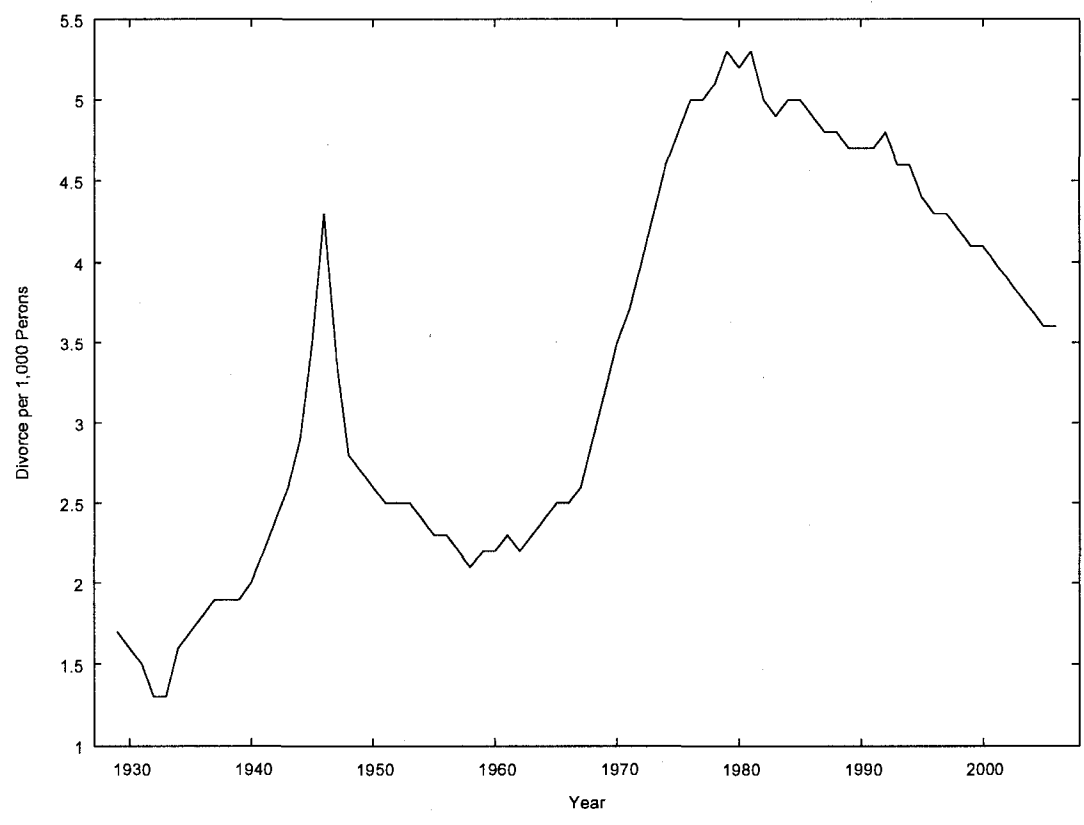
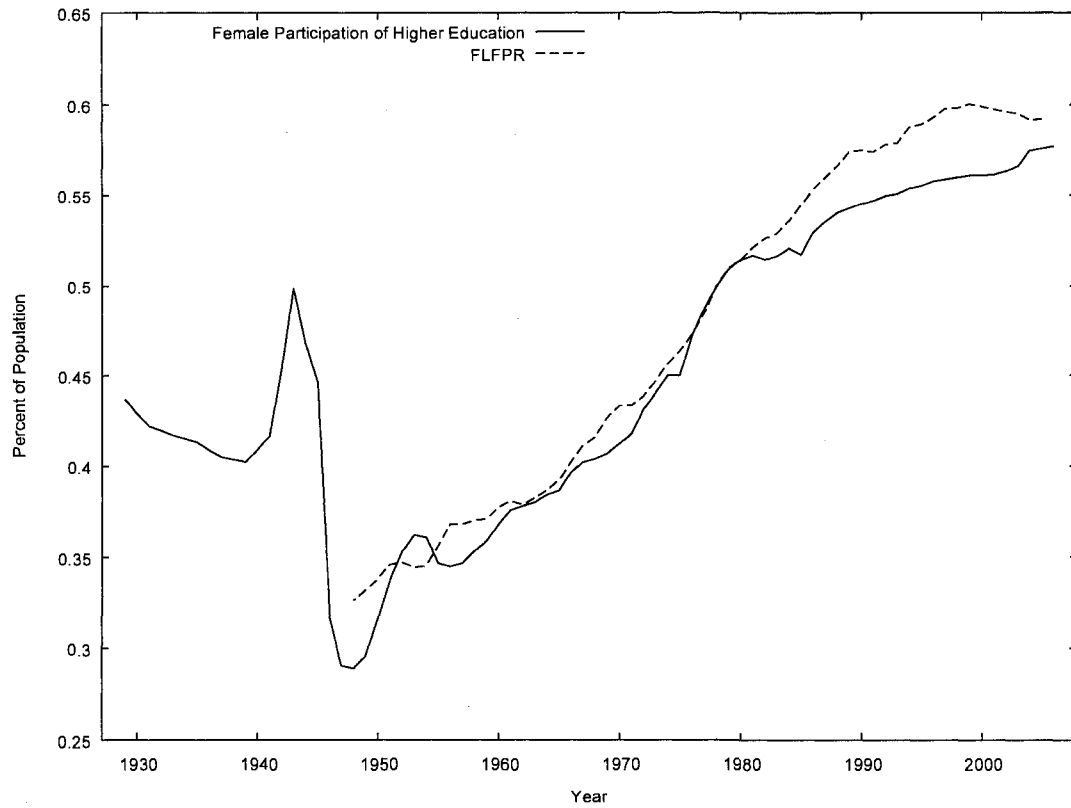


FIGURE 2—TIME SERIES PLOT OF FEMALE PARTICIPATION IN HIGHER EDUCATION AND THE FEMALE LABOR-FORCE PARTICIPATION RATE (FLFPR)



**FIGURE 3—SCATTERPLOT OF THE DIVORCE RATE AND
THE FEMALE LABOR-FORCE PARTICIPATION RATE
(1960-2001, LEAST SQUARES FIT)**

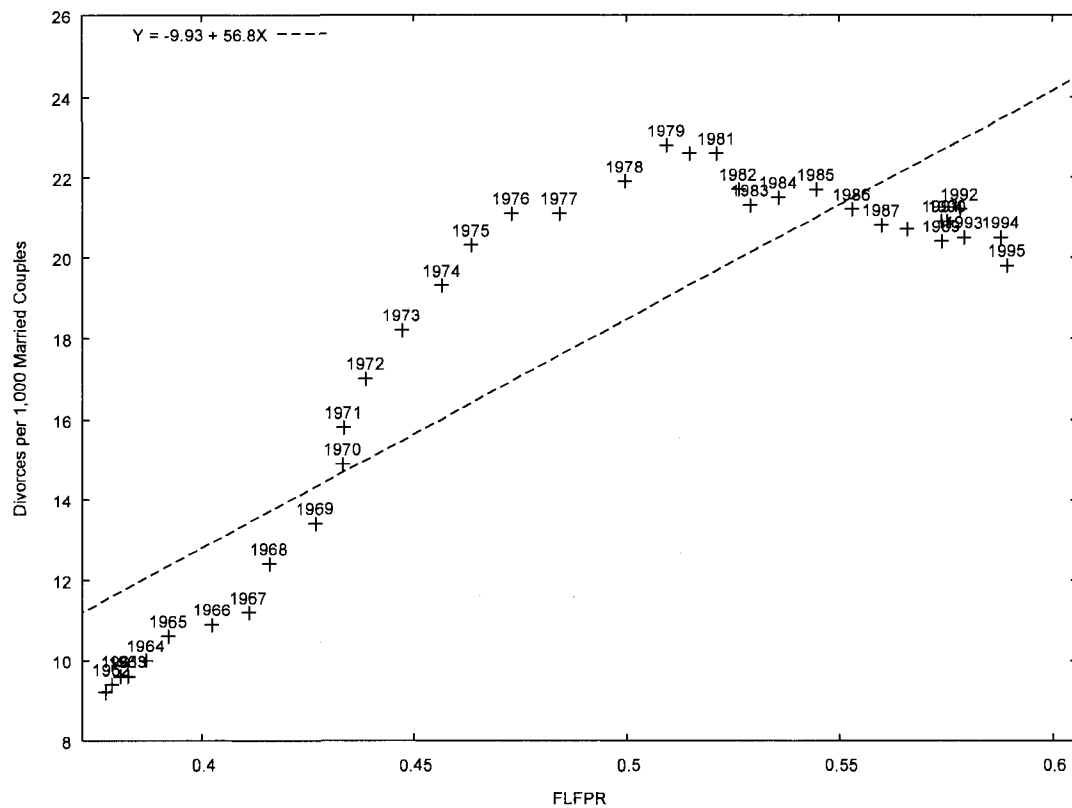


FIGURE 4—SCATTERPLOT OF THE DIVORCE RATE AND FEMALE PARTICIPATION IN HIGHER EDUCATION (1929-2006, LEAST SQUARES FIT)

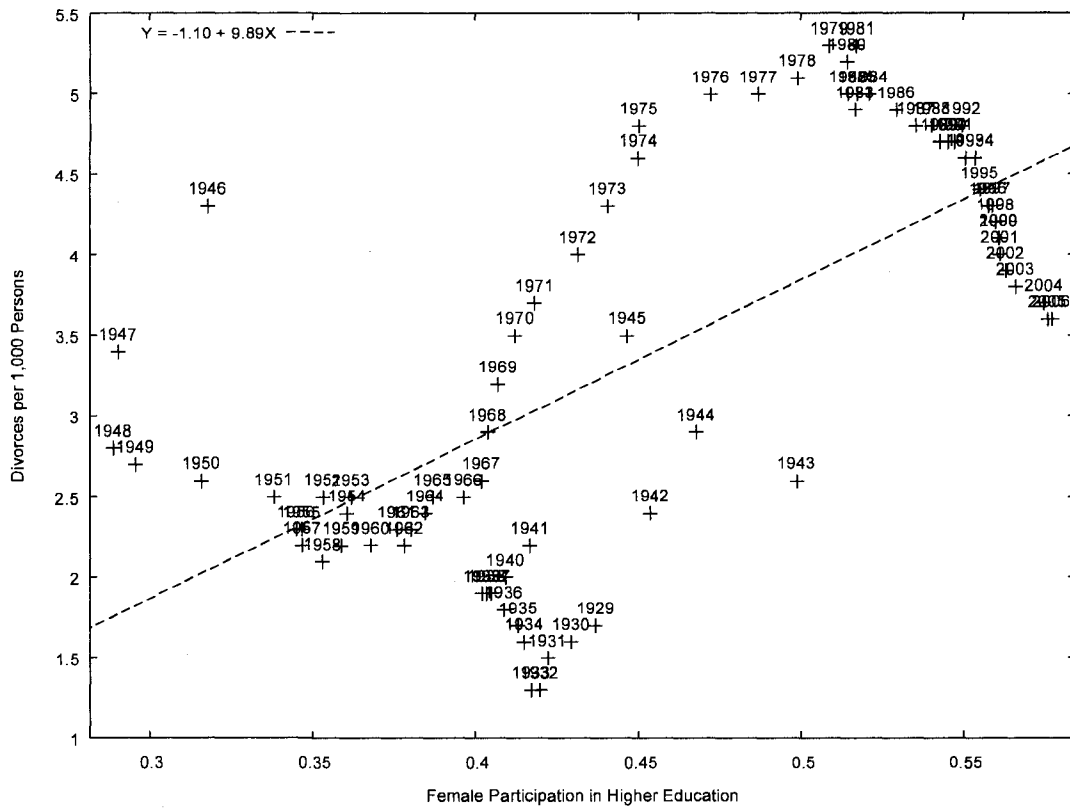
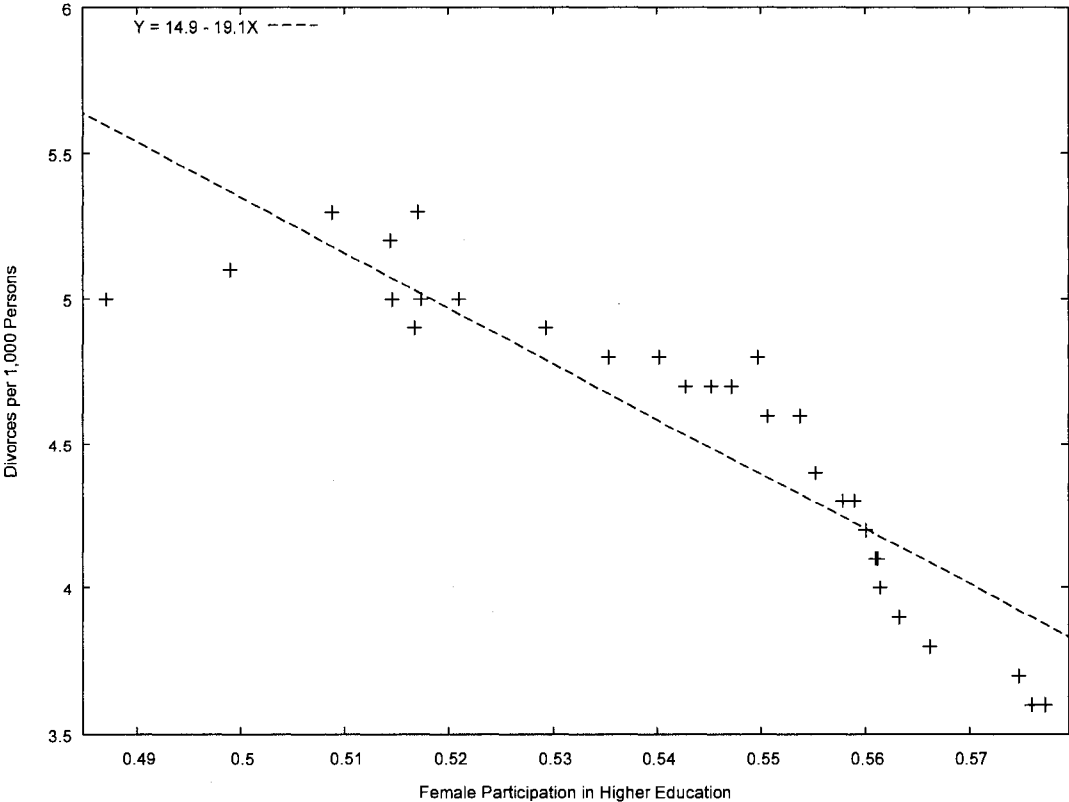
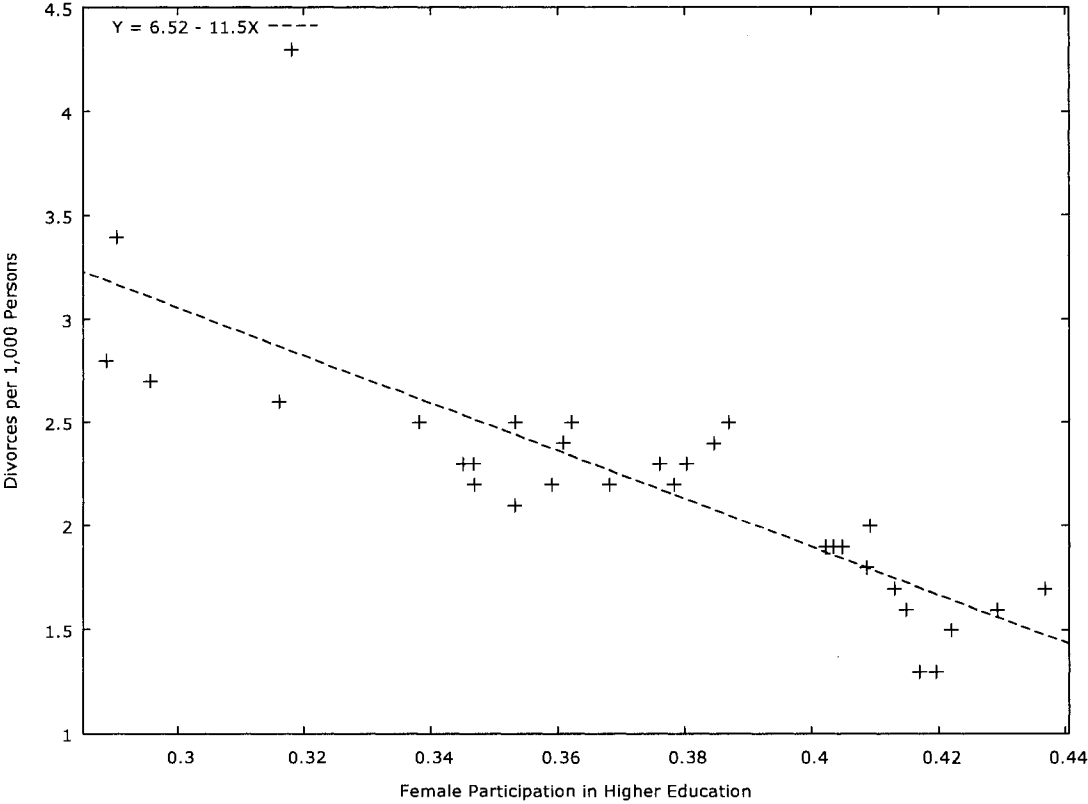


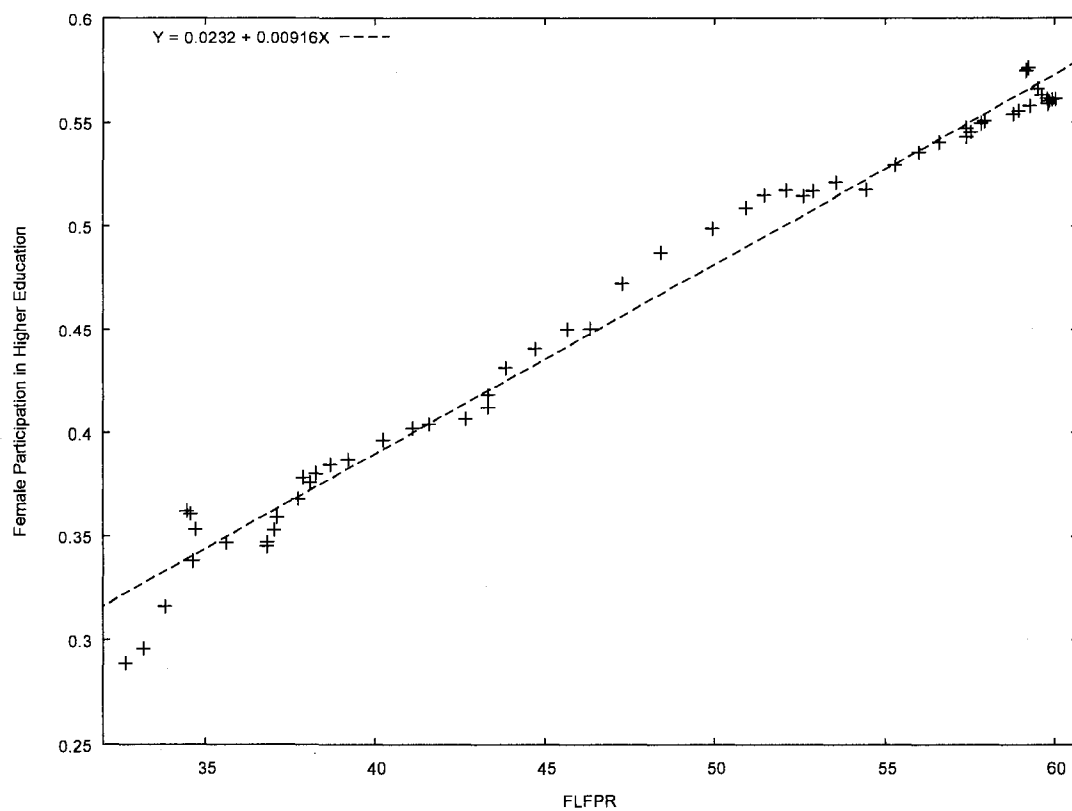
FIGURE 5—SCATTERPLOT OF THE DIVORCE RATE AND FEMALE PARTICIPATION IN HIGHER EDUCATION (1977-2006, LEAST SQUARES FIT)



**FIGURE 6—SCATTERPLOT OF THE DIVORCE RATE
FEMALE PARTICIPATION IN HIGHER EDUCATION
(1929-1965, EXCEPT 1941-1945, LEAST SQUARES FIT)**



**FIGURE 7—FEMALE PARTICIPATION IN HIGHER EDUCATION
AND FEMALE LABOR-FORCE PARTICIPATION (FLFPR)
(LEAST SQUARES FIT)**



**FIGURE 8—PERCENT OF POPULATION WITH
INCREASED ACCESS TO ORAL CONTRACEPTIVES**

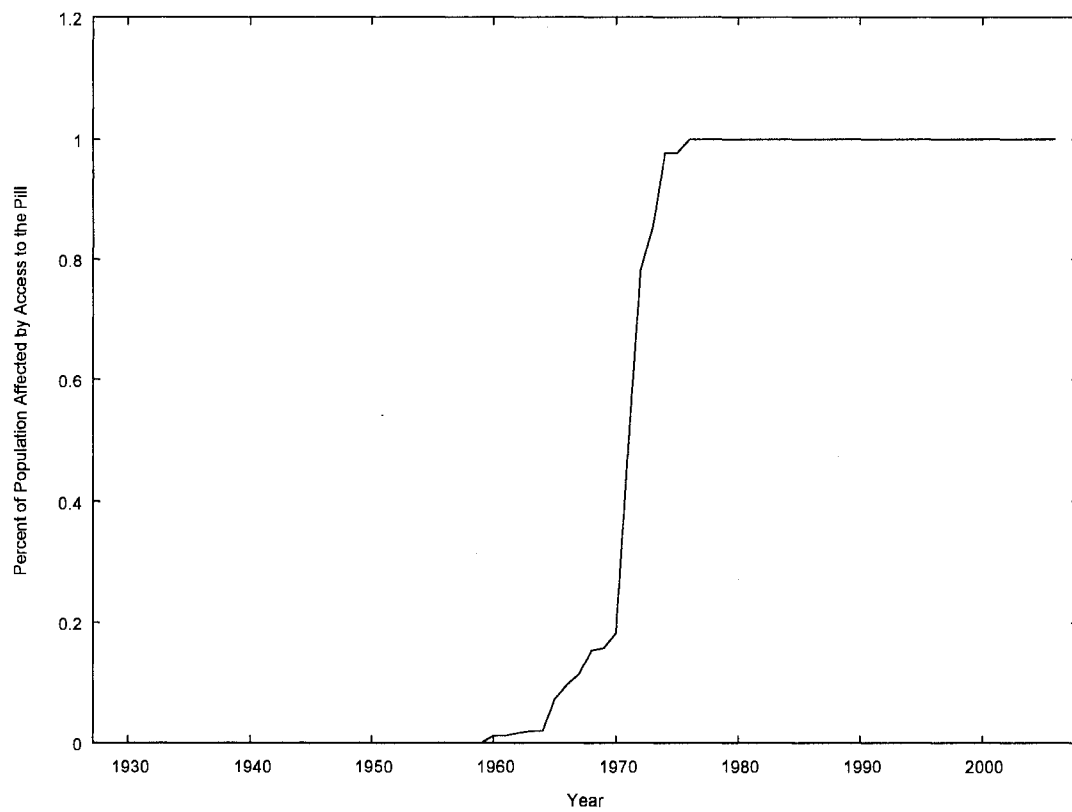


FIGURE 9—SCATTERPLOT OF THE DIVORCE RATE AND THE PERCENTAGE OF THE POPULATION WITH INCREASED ACCESS TO ORAL CONTRACEPTIVES (1929-2006, LEAST SQUARES FIT)

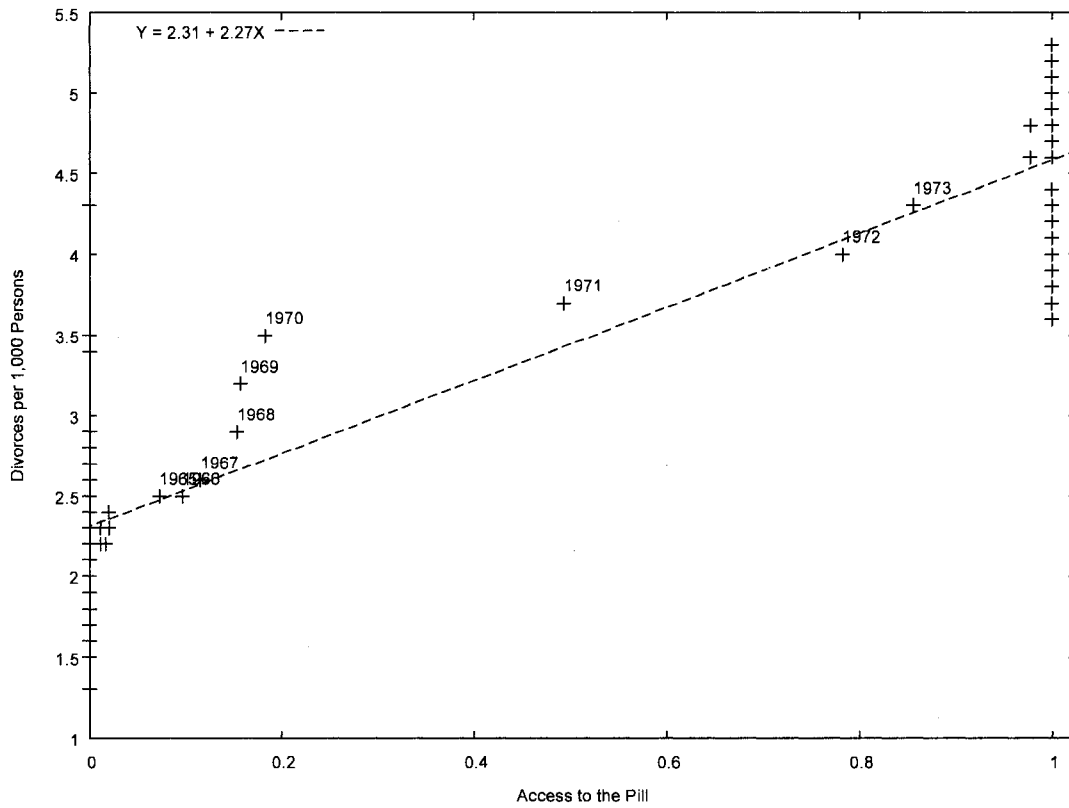
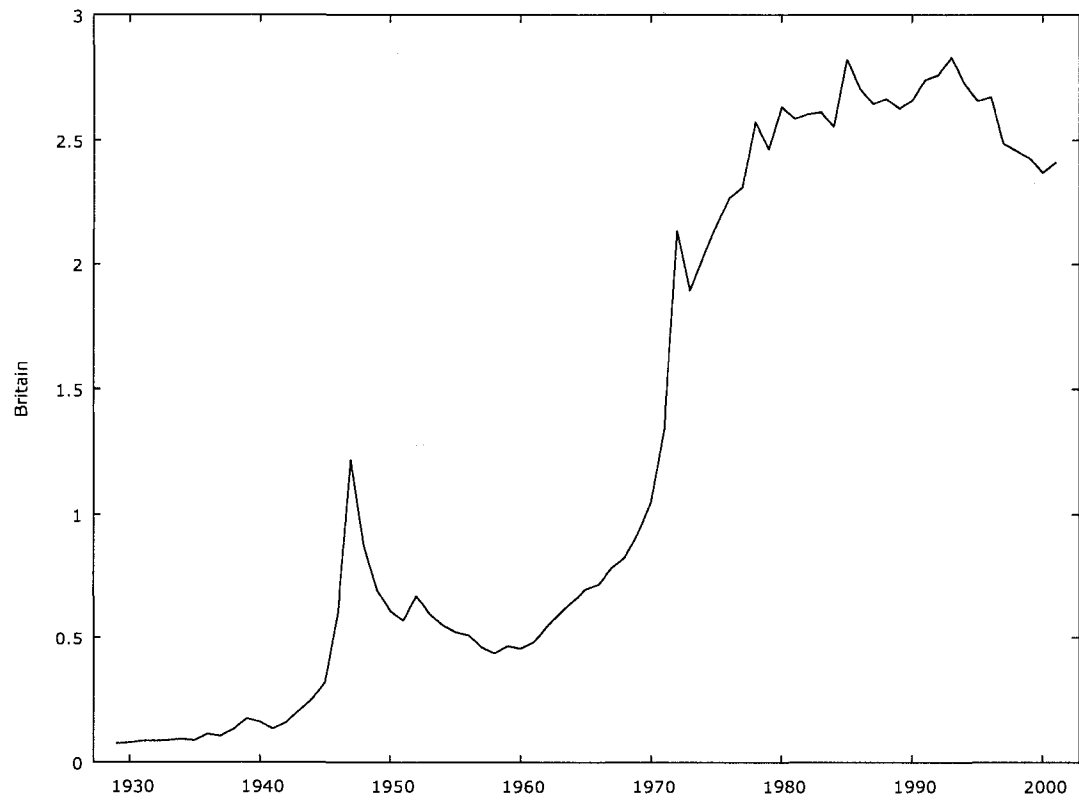
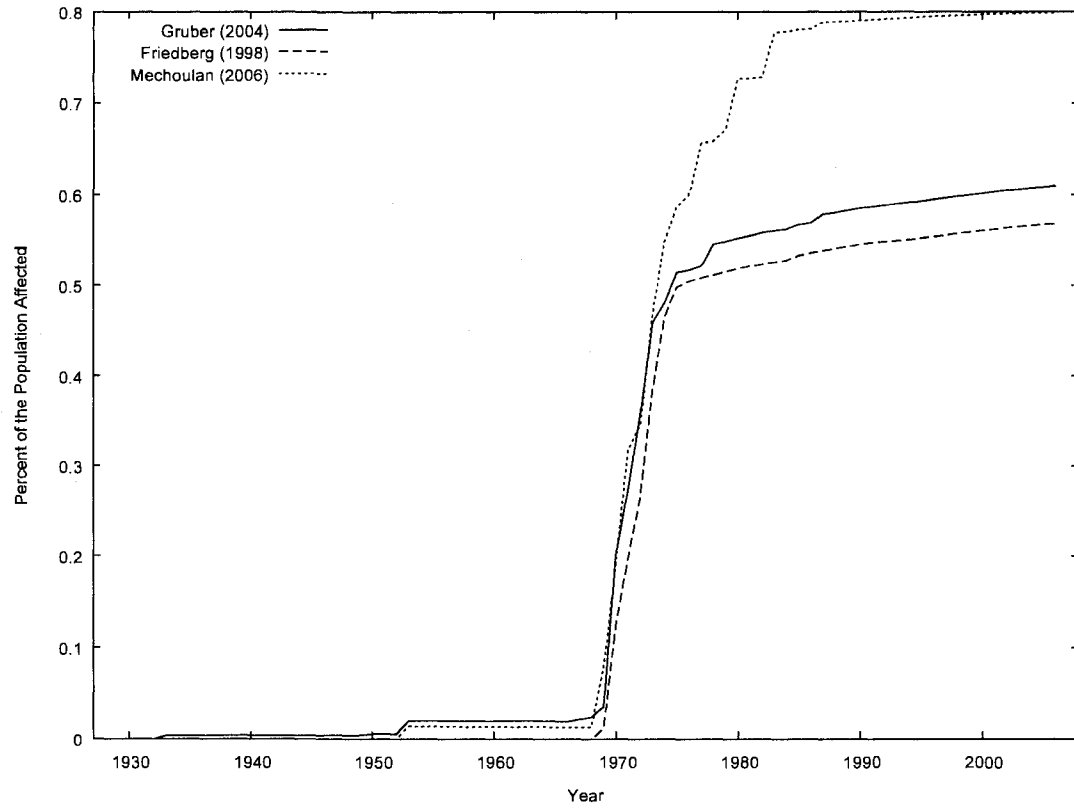


FIGURE 10—DIVORCE RATE PER CAPITA IN GREAT BRITAIN (BRITAIN)

**FIGURE 11—PERCENTAGE OF THE POPULATION
AFFECTED BY UNILATERAL DIVORCE REFORM**



**FIGURE 12—TIME PLOTS OF DIVORCE PER 1,000 PERSONS
AND DIVORCES PER 1,000 MARRIED COUPLES**

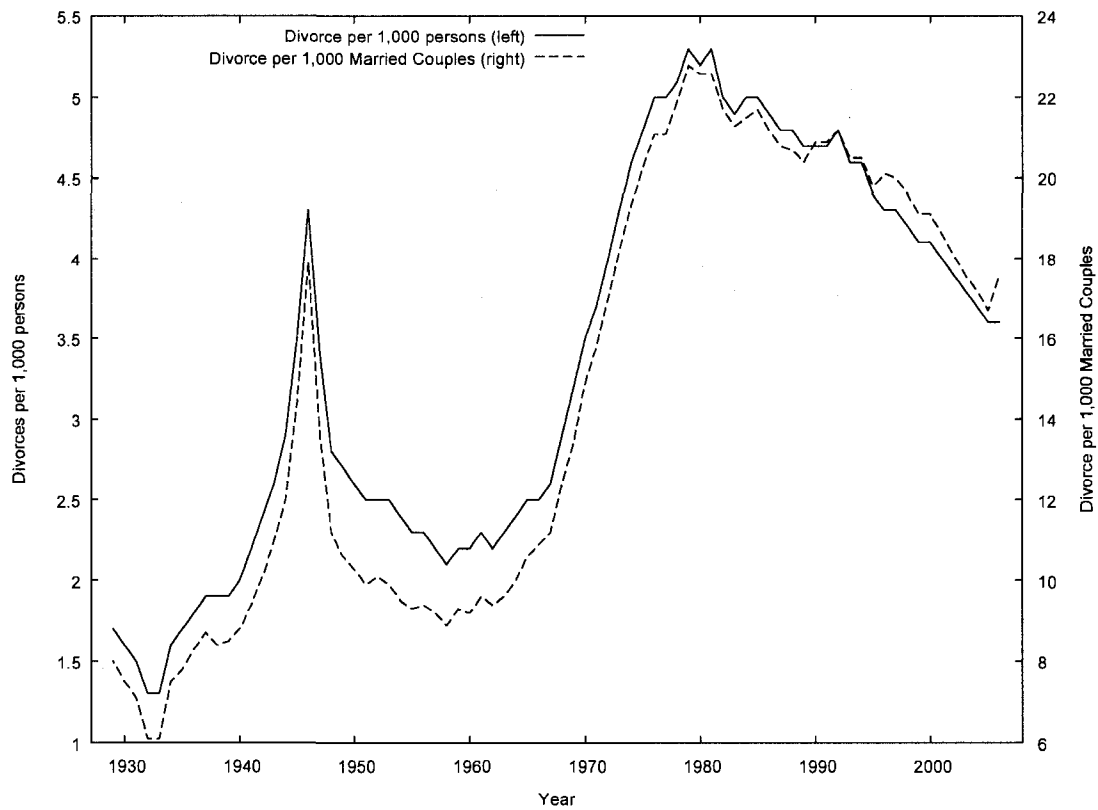


FIGURE 13—SCATTERPLOT OF THE DIVORCES PER 1,000 PERSONS AND DIVORCES PER 1,000 MARRIED COUPLES RATE (1929-2006, LEAST SQUARES FIT)

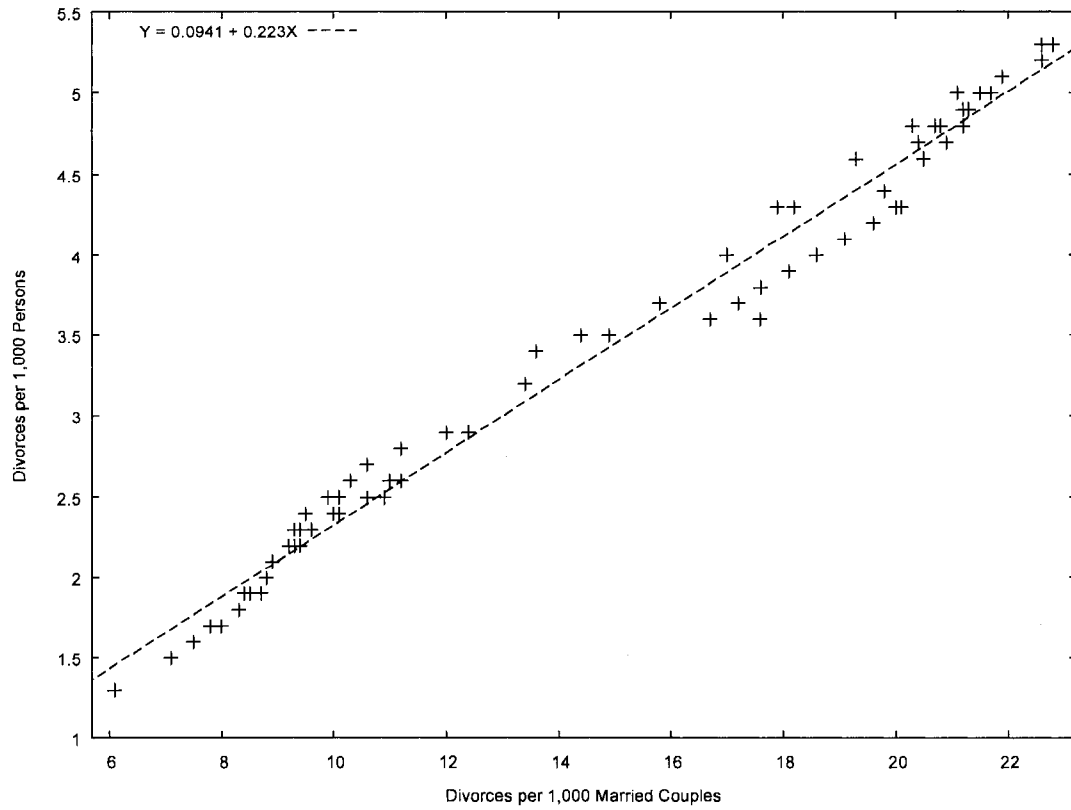


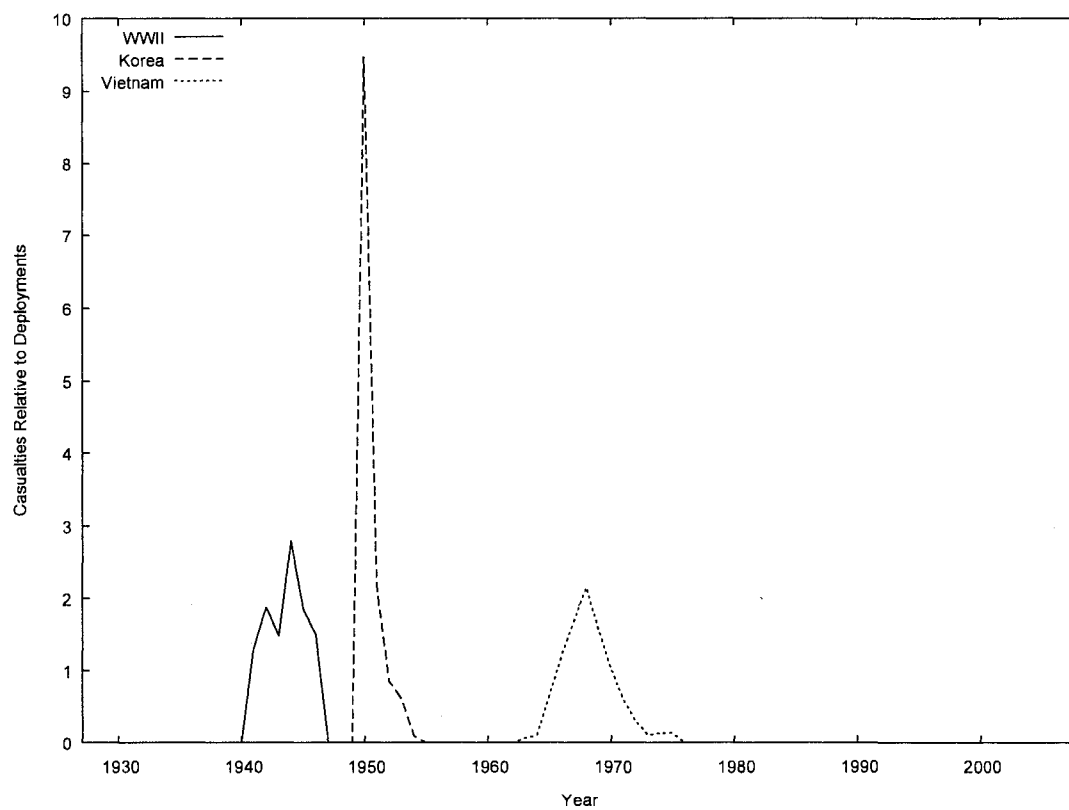
FIGURE 14—WAR VARIABLES: CASUALTIES RELATIVE TO DEPLOYMENTS

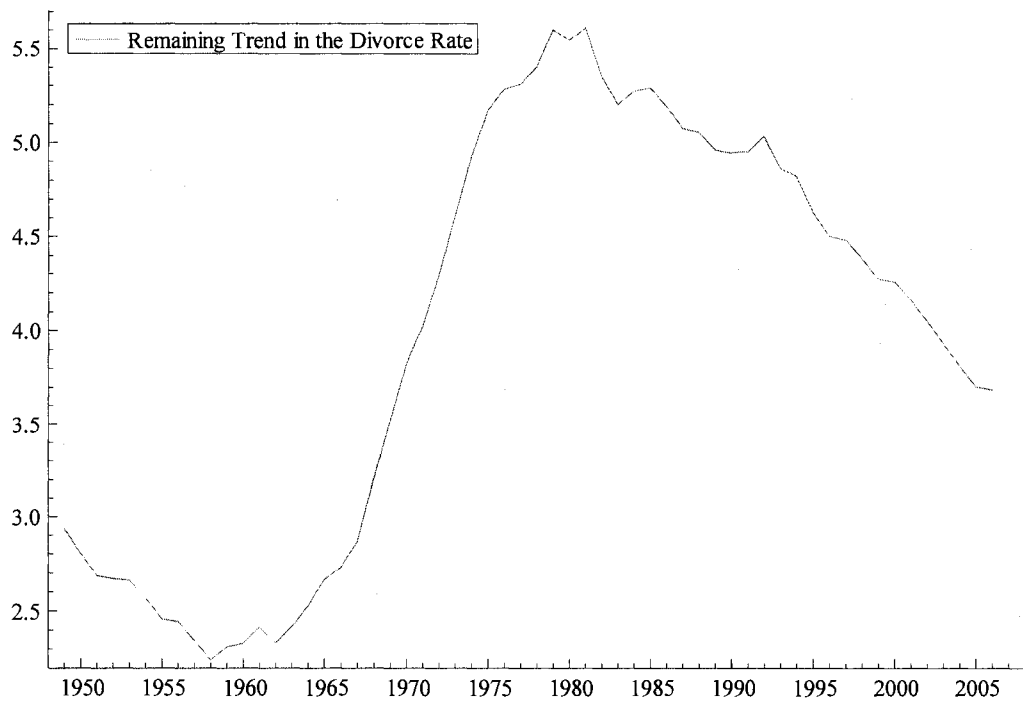
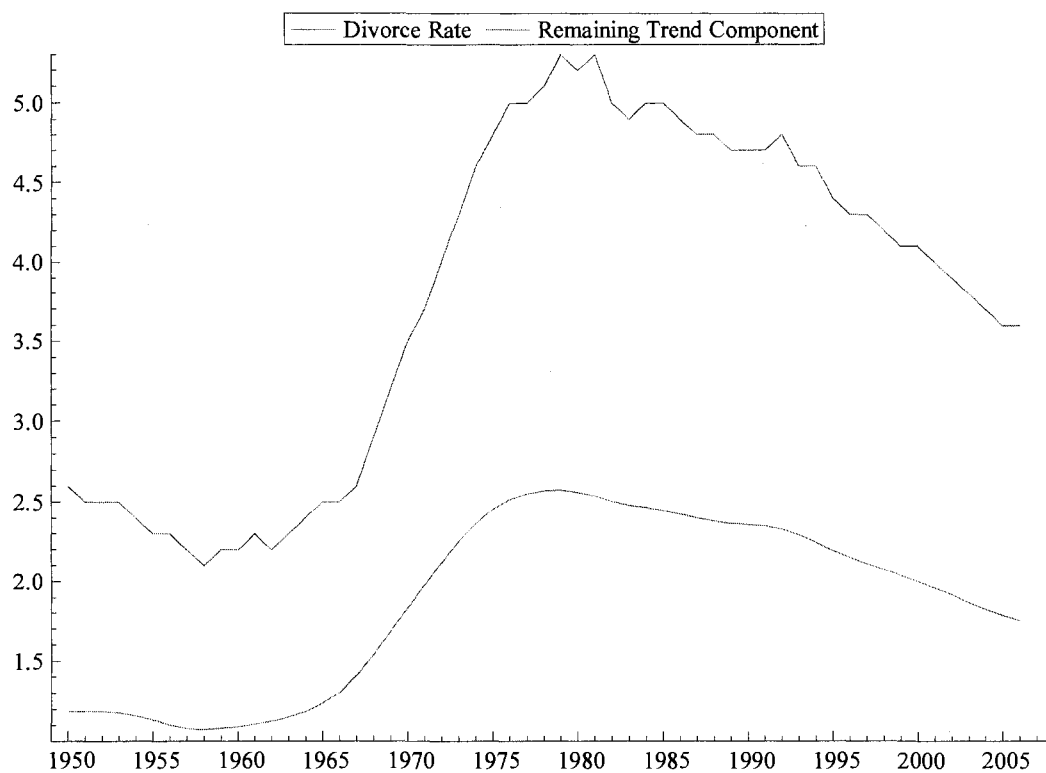
FIGURE 15—REMAINING TREND COMPONENT FOR MODEL 1 FROM TABLE 4

FIGURE 16—COMPARISON OF ACTUAL DIVORCE RATE WITH THE REMAINING TREND COMPONENT EXCLUDING THE IMPACT OF THE VIETNAM WAR AND EARLY ACCESS TO THE PILL



APPENDIX

TABLE A1: OLS ESTIMATES, 1949-2006
(SUBSTITUTING FEMALE LABOR-FORCE PARTICIPATION)

	Model 1		Model 2		Model 3		Model 4	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
<i>constant</i>	0.8662	0.001	-0.5422	0.442	-0.7411	0.270	-0.7588	0.155
<i>divorce</i> (-1)	0.8750	0.000	0.8759	0.000	0.8807	0.000	0.8987	0.000
<i>flfpr</i>	-1.7959	0.001	4.3703	0.136	4.6766	0.079	4.9681	0.027
<i>flfpr</i> ²			-6.2563	0.029	-6.3263	0.014	-6.6877	0.003
<i>pill</i>	0.6101	0.000	0.5207	0.000	0.4719	0.003	0.4588	0.000
<i>inflation</i>	0.0100	0.117	0.0078	0.153	0.0105	0.031	0.0062	0.182
<i>ygrowth</i>	0.0095	0.080	0.0330	0.040	0.0586	0.010	0.0129	0.644
<i>ygrowth</i> ²			-0.0039	0.059	-0.0048	0.025	0.0008	0.795
<i>das*ygrowth</i>							0.1808	0.001
Δ <i>unemp</i>					0.0470	0.072	0.0526	0.046
<i>Vietnam</i> (-1)	0.1940	0.000	0.1431	0.000	0.1334	0.000	0.1310	0.000
<i>Korea</i>							-0.0137	0.112

Notes: (-1) denotes a lag order of one. Δ is the first-difference operator. The models presented here are the same as those shown in Table 4, but with the FLFPR substituted for female participation in higher education.

TABLE A2: OLS ESTIMATES, 1949-2006
(DEPENDENT VARIABLE: DIVORCES PER 1,000 MARRIED COUPLES)

	Model 1		Model 2		Model 3		Model 4	
	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>	<i>coeff.</i>	<i>p-value</i>
constant	3.0559	0.000	-3.5968	0.203	-3.5447	0.191	-3.7998	0.188
<i>divorce</i> (-1)	0.8408	0.000	0.8393	0.000	0.8404	0.000	0.8524	0.000
<i>fem_ratio</i>	-5.2590	0.003	25.7631	0.037	23.7353	0.042	25.5709	0.049
<i>fem_ratio</i> ²			-33.8356	0.011	-30.8815	0.015	-32.9708	0.022
<i>pill</i>	2.5036	0.000	2.3564	0.000	2.2610	0.000	2.1947	0.000
<i>inflation</i>	0.0711	0.027	0.0524	0.069	0.0620	0.031	0.0508	0.130
<i>ygrowth</i>	0.0426	0.095	0.1007	0.099	0.1867	0.038	0.0475	0.705
<i>ygrowth</i> ²			-0.0107	0.184	-0.0141	0.105	0.0030	0.828
<i>das*ygrowth</i>							0.5394	0.075
Δ <i>unemp</i>					0.1548	0.114	0.1703	0.089
<i>Vietnam</i> (-1)	0.7063	0.000	0.5152	0.000	0.5007	0.000	0.4866	0.000
<i>korea</i>							-0.0403	0.249
Unadjusted R ²	0.9946		0.9956		0.9957		0.9960	
Adjusted R ²	0.9939		0.9948		0.9949		0.9950	
Log-likelihood	29.922		18.157		17.057		15.078	
AIC	61.845		54.314		54.114		54.155	
SBC	76.268		72.858		74.719		78.881	
HQC	67.463		61.537		62.140		63.786	
RESET		0.027		0.426		0.414		0.548
Homoskedasticity		0.021		0.325		0.451		0.396
Normality		0.916		0.058		0.031		0.131
Autocorrelation LM(1)		0.711		0.475		0.591		0.240
Autocorrelation LM(2)		0.726		0.702		0.804		0.491
Autocorrelation LM(3)		0.339		0.296		0.522		0.729
ARCH (1)		0.335		0.388		0.294		0.247
ARCH (2)		0.531		0.695		0.545		0.529
ARCH (3)		0.826		0.734		0.644		0.495
Harvey-Collier (cusum)		0.016		0.951		0.619		0.687

Notes: (-1) denotes a lag order of one. AIC stands for the Akaike Information Criterion, SBC for the Schwarz-Bayesian Criterion, HQC for the Hannan-Quinn Criterion; RESET is Ramsey's test for correct functional form; Homoskedasticity is White's test; Normality is a test for normality of the residuals; the null of no autocorrelation at various lag lengths is tested with the Breusch-Godfrey test; ARCH tests the null of no relationship between the current error variance and its past values; Quandt LR tests for the lack of structural breaks (Stock and Watson 2007); and Harvey-Collier tests parameter stability using cumulated recursive residuals.

**TABLE A3: LEAST SQUARES ESTIMATES, 1929-2006, HAC COVARIANCE MATRIX
(DEPENDENT VARIABLE: DIVORCES PER 1,000 MARRIED COUPLES)**

	<i>coeff.</i>	<i>P-value</i>	<i>coeff.</i>	<i>P-value</i>	<i>coeff.</i>	<i>P-value</i>
constant	-7.5267	0.005	-6.7731	0.013	-5.8445	0.010
<i>divorce</i> (-1)	0.7422	0.000	0.7153	0.000	0.8004	0.000
<i>fem_ratio</i>	52.0614	0.000	45.5483	0.000	37.8883	0.000
<i>fem_ratio</i> ²	-65.3729	0.000	-56.7339	0.000	-46.8438	0.000
<i>pill</i>	3.9724	0.005	4.0264	0.003	2.7190	0.000
<i>inflation</i>	-0.0079	0.817	0.0330	0.367	0.0527	0.084
<i>ygrowth</i>	-0.0759	0.145	0.1459	0.049	0.0755	0.318
<i>ygrowth</i> ²	-0.0056	0.006	-0.0179	0.000	-0.0138	0.001
<i>das*ygrowth</i>			-0.3924	0.016	-0.3147	0.013
Δ <i>unemp</i>	-0.2205	0.019	-0.1173	0.036	-0.1782	0.030
<i>WWII</i>	3.8986	0.000	4.6660	0.000	3.6386	0.000
<i>WWII</i> ²	-1.1333	0.000	-1.4346	0.000	1.0950	0.000
<i>Vietnam</i> (-1)	0.1029	0.001	0.5100	0.000	0.5348	0.000
<i>Korea</i>	0.7422	0.000	0.0720	0.014	0.0464	0.200
<i>d47</i>					-2.9651	0.000
Unadjusted R ²	0.9904		0.9918		0.9939	
Adjusted R ²	0.9886		0.9901		0.9926	
Log-likelihood	59.851		53.718		42.164	
Akaike	145.703		135.435		114.328	
Schwarz Bayesian	176.172		168.248		149.485	
Hannan-Quinn	157.890		148.560		128.391	
RESET		0.191		0.054		0.069
White Heteroskedasticity		0.000		0.000		0.024
Normality		0.000		0.047		0.249
Autocorrelation LM(1)		0.079		0.135		0.034
Autocorrelation LM(2)		0.198		0.290		0.210
Autocorrelation LM(3)		0.063		0.024		0.032
ARCH (1)		0.001		0.011		0.997
ARCH (2)		0.001		0.028		0.493
ARCH (3)		0.003		0.054		0.674
Harvey-Collier (CUSUM)		0.940		0.465		0.482

Notes: (-1) denotes a lag order of one. AIC stands for the Akaike Information Criterion, SBC for the Schwarz-Bayesian Criterion, HQC for the Hannan-Quinn Criterion; RESET is Ramsey's test for correct functional form; Homoskedasticity is White's test; Normality is a test for normality of the residuals; the null of no autocorrelation at various lag lengths is tested with the Breusch-Godfrey test; ARCH tests the null of no relationship between the current error variance and its past values; Quandt LR tests for the lack of structural breaks (Stock and Watson 2007); and Harvey-Collier tests parameter stability using cumulated recursive residuals.

TABLE A4: LEAST ABSOLUTE DEVIATIONS ESTIMATES, 1929-2006
(DEPENDENT VARIABLE: DIVORCES PER 1,000 MARRIED COUPLES)

	<i>coeff.</i>	<i>p-</i> <i>value</i>	<i>coeff.</i>	<i>p-</i> <i>value</i>	<i>coeff.</i>	<i>p-</i> <i>value</i>
<i>constant</i>	-2.1035	0.268	-3.7553	0.051	-5.3811	0.014
<i>divorce (-1)</i>	0.8342	0.000	0.8325	0.000	0.8074	0.000
<i>fem_ratio</i>	20.0027	0.016	26.4521	0.018	33.8394	0.000
<i>fem_ratio</i> ²	-29.5321	0.001	-35.8924	0.000	42.9335	0.000
<i>pill</i>	2.5809	0.000	2.5505	0.000	2.7624	0.000
<i>inflation</i>	0.0549	0.014	0.0520	0.011	0.0575	0.008
<i>ygrowth</i>	0.1608	0.005	0.1996	0.000	0.1949	0.000
<i>ygrowth</i> ²	-0.0160	0.000	-0.0176	0.000	-0.0177	0.000
<i>das*ygrowth</i>	-0.3880	0.000	-0.4066	0.000	-0.4032	0.000
Δ <i>unemp</i>	-0.0764	0.252				
<i>WWII</i>	3.8596	0.000	3.9937	0.000	4.0080	0.000
<i>WWII</i> ²	-1.1711	0.000	-1.2324	0.000	-1.2346	0.000
<i>Vietnam (-1)</i>	0.3977	0.002	0.4087	0.003	0.4316	0.002
<i>Korea</i>	-0.0125	0.818				
<i>d47</i>	-3.5531	0.000	-3.1161	0.000		
Sum of Absolute Residuals	22.3727		22.7019		25.795	
ARCH (1)		0.493		0.429		0.000
ARCH (2)		0.792		0.275		0.001
ARCH (3)		0.916		0.462		0.002

Notes: (-1) denotes a lag order of one. Δ is the first-difference operator. *d47* is an indicator variables for year 1947. We begin with the final model from Table 7, and test the model down to its most parsimonious form. We only report tests for ARCH effects, as the LAD estimator allows for outliers in the data. Model 9 provides the best fit to the data, as it is more parsimonious than Model 8 and has no ARCH effects present.

TABLE A5: COINTEGRATION AND VECM RESULTS, 1929 TO 2006
(DEPENDENT VARIABLE: DIVORCES PER 1,000 MARRIED COUPLES)

	Model 1	Model 2	Model 3
<i>Implicit CIV Normalized on Divorce</i>			
constant	60.933 (8.521)	57.382 (7.822)	53.105 (9.314)
<i>fem_ratio</i>	-85.091 (-6.739)	-131.665 (-6.744)	-118.041 (-7.807)
<i>pill</i>	27.488 (12.336)		
<i>gruber</i>		65.044 (10.431)	
<i>friedberg</i>			62.762 (12.275)
<i>Cointegration Trace Test</i>			
p-values:			
null of rank = 0	0.000	0.001	0.000
null of rank = 1	0.139	0.284	0.321
<i>Statistics and Specification Tests of VECM</i>			
Restricted Log-likelihood	369.549	370.292	372.119
SBC	-8.303	-8.323	-8.373
HQC	-8.888	-8.908	-8.958
R ² divorce equation	0.660	0.688	0.702
R ² second equation	0.601	0.575	0.582
P-values of system tests:			
ARCH (1)	0.000	0.000	0.000
ARCH (2)	0.000	0.000	0.000
Ljung Box Q (18)	0.599	0.508	0.495
Autocorrelation (1)	0.266	0.093	0.042
Autocorrelation (2)	0.381	0.581	0.447
Normality	0.000	0.000	0.000

Note: T-values are in parenthesis. 5 lags are used. The p-values of the cointegration rank test are based on simulated values with 2,500 replications and random walks of length 400. The effective sample is 1934 to 2006. Calculations are done in CATS in RATS, version 2 (Dennis et al. 2005).

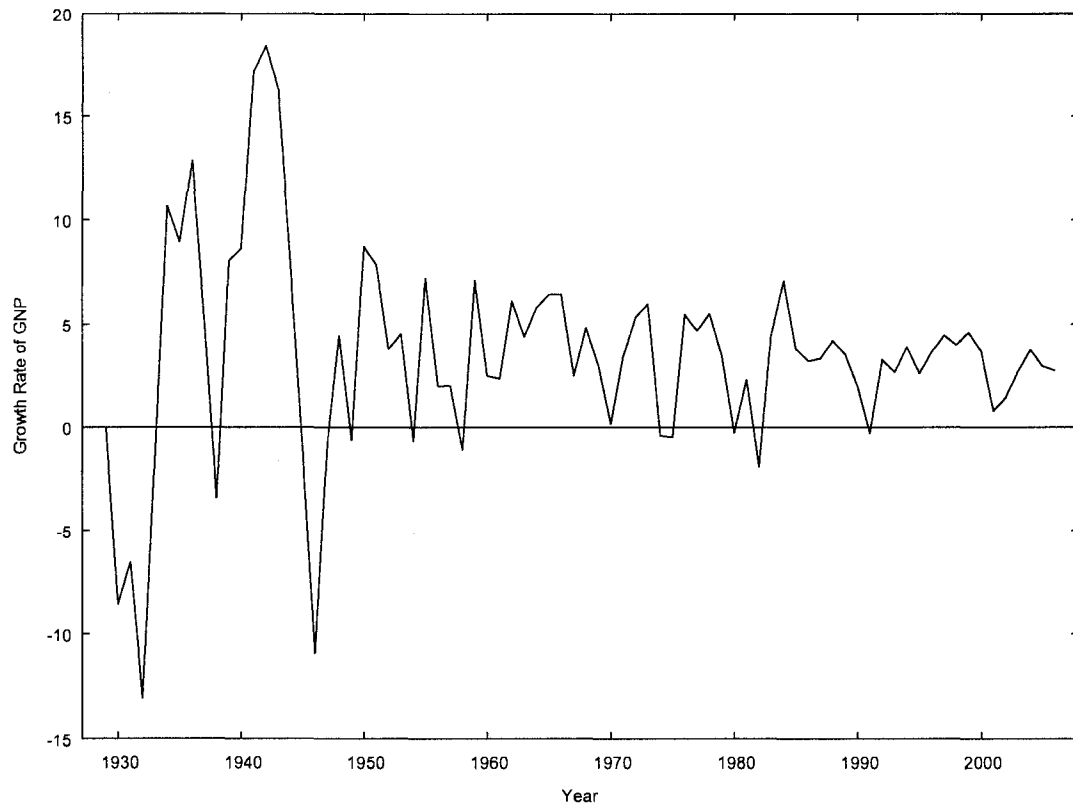
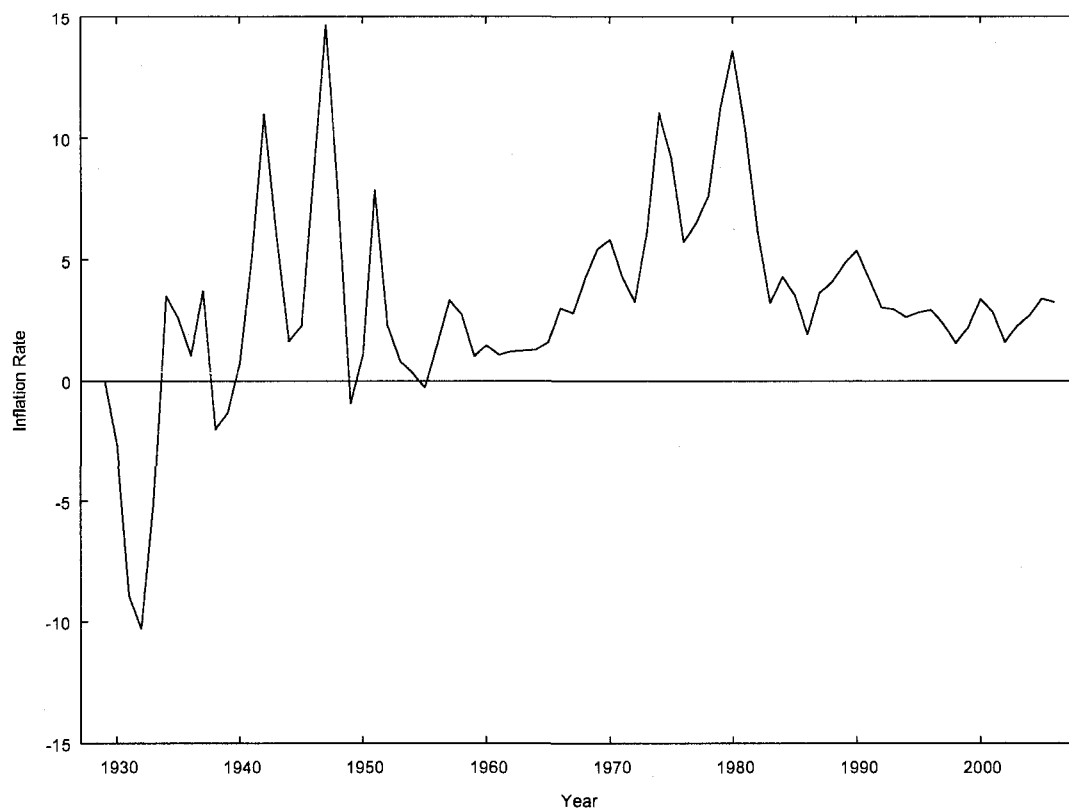
FIGURE A1—GROWTH RATE OF GROSS NATIONAL PRODUCT (GNP) OVER TIME

FIGURE A2—INFLATION RATE OVER TIME



CHAPTER 5

CHILD-CUSTODY REFORM AND MARRIAGE-SPECIFIC INVESTMENT IN CHILDREN

(with Alan Seals)

5.1. INTRODUCTION

Child custody is one of the most common and controversial issues in family court. During the first half of the 20th century, courts in the U.S. typically favored the mother in child-custody cases (Jacob 1988, Ch. 8; Brinig and Buckley 1998a). In the 1960s, states began to remove the explicit preference for mothers so that a parent's gender was no longer the basis for child-custody awards. Even after the move away from maternal preference, most courts continued to award sole custody to mothers (Cancian and Meyer 1998). In the 1970s, several states made explicit provisions in their laws favoring joint custody or revealed their preference indirectly by ruling in favor of joint custody (Brinig and Buckley 1998a). Citing the best interests of children as the impetus for legislative change, the majority of states followed with legal provisions for joint custody by the mid-1980s (Brinig and Buckley 1998a; Cancian and Meyer 1998). Although custody reform became a nation-wide phenomenon, the debate over joint custody's costs and benefits

was carried out by a relatively small group of politically active supporters, with little empirical evidence to support their claims (Jacob 1988, Ch. 8).

Analyzing the effects of family-law reform on marital investment in children provides an opportunity to study the bargaining behavior of spouses. Changes in child-custody laws may significantly alter the amount of time each parent spends with their children after divorce occurs. Since the allocation of child custody could affect the value of child-specific investment to parents in the event of divorce, it is useful to investigate the impact joint-custody laws have on outcomes of intact families. A Coasian analysis suggests that when bargaining costs are absent married couples will reach an efficient child-investment outcome, regardless of the initial assignment of custodial rights. However, there is evidence that changes in family laws shape the bargaining process over the course of marriage (e.g., see Browning et al. 1994; Gray 1998; Lundberg et al. 1997; Duflo 2003; Genadek et al. 2007; Stevenson 2007, 2008; Ward-Batts 2008). If child-custody reform alters the distribution of the marital surplus after divorce occurs, bargaining models of family behavior predict that married couples will change their child-specific investment behavior.¹

This is the first study to examine the impact of joint-custody laws on marriage-specific investment in children. We exploit the timing of child-custody laws across states, with data from the 1980 and 1990 U.S. Population Censuses, to identify the effect of joint-custody reform on married couples' investment in child quality.² Married

¹ See Manser and Brown (1980), McElroy and Horney (1981), and Rasul (2006). See Bergstrom (1996) and Lundberg and Pollak (1996) for review of various family models.

² We use Brinig and Buckley's (1998a) coding for the child-custody laws. See TABLE I.

couples with children in states that change their laws to favor joint custody between 1980 and 1990 constitute the treatment group in a natural experiment. Married couples with children who live in states that had either instituted joint-custody reform before 1980 or that did not institute joint-custody reform before 1990 represent the comparison group.

The dependent variable is children's private school attendance—a verifiable marriage-specific investment in child quality. Although most children in the U.S. attend public school and private school represents only one of the many investments parents can make in child quality, we feel that private school proxies well for overall parental investment.³ Since the financial costs of private school warrant long-run planning by parents, any observed differences in private school attendance resulting from joint-custody reform could be extrapolated to other forms of child investment.

Considering property-division laws in conjunction with joint-custody reform may provide a clearer picture of household bargaining. Gray (1998) suggests the effects of unilateral divorce laws on marital bargaining are conditional on the underlying property-division laws across states, with bargaining power shifting to women in community-property and equitable-division states and to men in common-law states. To further examine marital investment in children and other assets as an outcome of spousal bargaining, we consider the impact of child-custody reform by the property-division laws in place across states and the consideration of marital fault in the division of marital assets.

³ Cáceres-Depliano (2006) and Conley and Glauber (2006) also use children's private school attendance and grade retention as proxies for parental investment in child quality. Unfortunately, our data set does not allow us to examine the impact of joint-custody laws on grade retention.

Because socioeconomic status (SES) is likely to play a role in the decision for married couples to send their children to private school, we also partition the sample by mother's education in supplemental models. We choose mother's education as a proxy for SES because household income is a likely endogenous variable.

We find a marginally statistically significant, 1.2 percentage point decline in the probability of children's private school attendance in states that enact joint custody. The effects of joint-custody reform on children's private school attendance are larger in common-law and community-property states, with 3.0 and a 2.5 percentage point declines, respectively. These findings suggest that married couples bargain over child-specific investment and other marital assets, with negative consequences for children in states that enact joint custody with property-division laws that favor one spouse over another.

We also find that child-investment behavior differs between married couples of varying SES. For the lowest SES group, joint-custody reform reduces the probability of children's private school attendance by 5.2 percentage points. For the other SES groups, we find no evidence that joint custody reform affects marital investment in children's private school attendance. However, we find that the effects of child-custody reform for the higher SES groups depend on the type of property-division regime in place. For the highest SES group, the probability of children's private school attendance declines by 6.2 and 3.4 percent in common-law states and states that do not consider marital fault in the divorce settlement that enact joint custody, respectively. The size and statistical significance of the estimated effects are highly sensitive to the addition of time-varying,

state-level controls. In particular, the estimated effects appear to be most sensitive to the inclusion of the state-level, welfare policy variables (e.g., AFDC and Food Stamp benefits) that appear to be correlated with the adoption of joint custody.

5.2. THEORETICAL FRAMEWORK

5.2.1. Legal Background

The authority of family-court judges to exercise wider discretion and institute joint-custody arrangements is a relatively recent legal innovation. Although child welfare was cited as the primary basis for child-custody reform, the passage of joint custody went against the widely held view among psychologists that sole custody was optimal (Goldstein et al. 1984). However, challenges to the sole-custody standard were issued at this time on the basis that it was an impulsion for post-marital conflict and therefore contrary to the best interests of the child (Stack 1976). As a result, when most states began enacting joint-custody legislation there was no consensus on the appropriate custodial arrangement (Jacob 1988, Ch. 8).

There were many underlying causes of child-custody reform. Women's increasing labor-force participation and the more prominent role of men in child rearing were both key demographic changes that helped facilitate joint-custody reform (Jacob 1988, Ch. 8). The preponderance of "dead-beat" parents (primarily fathers), who were in arrears of child-support payments, also generated political incentives to alter child-custody laws

(Jacob 1988, Ch. 8).⁴ Contrary to previous family-law reforms, such as no-fault divorce, expert opinion was relatively absent and personal experiences were more often cited in the legislative discourse on joint custody (Jacob 1988, Ch. 8). Because it was difficult to show the negative consequences for children and the potential gains came at a low cost to the public, joint-custody reform was discussed by a small group of proponents and passed legislatures in relative obscurity (Jacob 1988, Ch. 8).

A joint-custody provision relegates courts to handle only those custody disputes which cannot be settled privately (Buehler and Gerard 1995).⁵ In the event that child custody must be decided in court, judges have discretion to rule in favor of joint custody if it conforms to the best interests of the child.⁶ Depending on family-specific circumstances, joint custody can fall under a protocol of (i) joint legal custody in which parents share in the decisions of child upbringing but the child's primary residence is with one of the parents or (ii) joint physical custody in which both parents share in child-rearing decisions and also share physical custody of the child. Under either joint-custody settlement, courts expect divorced parents to maintain a cooperative relationship while raising their children.

Divorce settlements also depend on a state's specific property-division laws. In the event of divorce, courts distribute marital assets according to three property-division

⁴ Since welfare payments were a federal issue, by 1984, this political activity also included a U.S. Congress mandate that funds would be withheld from paychecks and federal tax returns to pay delinquent child-support (Jacob 1988, pp. 132). By simply granting greater custodial rights, joint custody could also have been a low-cost (for the state) incentive for fathers to pay child support. Brinig and Buckley (1998a) find a positive effect of joint-custody reform on child-support payments.

⁵ Many states mandate third-party mediation in child-custody cases (Fineman 1988).

⁶ See Buehler and Gerard (1995) for a discussion of the Best Interests of the Child (BIOC) standard and how the standard varies by state.

regimes: (i) equitable-division, (ii) common-law, or (iii) community property. Equitable-division property laws authorize judges to divide marital property as they see fit, which generally protects the most damaged individual in the event of divorce (Gray 1998). In common-law states, the spouse who holds legal title receives control of the property in the event of divorce. By contrast, community-property states attempt to divide marital assets evenly between spouses. Assuming the husband is the breadwinner, community-property states generally reward the wife a larger share of the marriage-specific assets following divorce, while common-law property states generally reward the husband a larger share of the marital assets (Gray 1998).

The consideration of marital wrongdoing in the division of property creates potential economic disadvantages for many spouses. Prior to no-fault divorce reform in the 1960s and 1970s, marital fault had to be proven or acknowledged by one party in order to be granted a divorce. During the 1970s and 1980s, many states also removed marital misconduct as a consideration in the division of marital assets (Ellman and Lohr 1998; Brinig and Buckley 1998b). This reform was largely implemented to expedite divorce proceedings and to prevent one spouse from extorting a greater (punitive) share of the marital assets from the at-fault spouse (Gruber 2004).

5.2.2. Child Investment in Models of Intrahousehold Distribution

Becker (1974, 1991) assumes that families pool all sources of income and an altruistic member makes decisions based on what is best for each family member.⁷ Becker's model predicts that changes in the laws governing child custody and marriage-specific assets would have the same effect on within-family distribution regardless of which spouse benefits.⁸ If prospective spouses can make costless, binding, distributional agreements, bargaining does not occur over the course of marriage. However, Becker's "common-preference" assumption may not be realistic in the context of child custody, since courts in the U.S. have been reluctant to recognize marital contracts specifying child custody in the event of divorce (Francesconi and Muthoo 2003).

Manser and Brown (1980) and McElroy and Horney (1981) develop models in which couples bargain over the marital surplus and divorce represents an external threat point or outside option. The key feature of these models is the role of environmental factors (e.g., laws governing the division of marriage-specific property and child custody), which determine the threat point in the bargaining game.⁹ A change in a state's child-custody

⁷ Pollak (1985) specifies Becker's model in the context of a two-stage bargaining game in which the altruist moves first and makes take-it-or-leave-it offers to household members. The difference in Pollak's and Becker's models is the former assumes that it is not altruism but his/her bargaining position within the family that determines intrahousehold distribution.

⁸ The main obstacle in testing Becker's model is finding an exogenous factor that affects the control of resources within a family. A number of studies have used nonlabor income to test the income-pooling assumption in Becker's model (e.g., Thomas (1990) and Schultz (1990)). Supporters of the income-pooling hypothesis typically conclude that nonlabor income may be endogenous; suggesting that the changes in within-family distribution relates to unobserved heterogeneity. Lundberg and Pollak (1996) and Behrman (1997) show that nonlabor income is not entirely exogenous, which provides support for those who favor the income-pooling model. Lundberg et al. (1999) and Ward-Batts (2002) provide evidence rejecting the altruist model. Although evidence rejecting the income-pooling hypothesis is sparse, statistical evidence supporting the income-pooling hypothesis is non-existent.

⁹ Lundberg and Pollak (1993) contend that a within-marriage outcome is more reasonable since the costs associated with divorce are often high. They assume the existence of traditional gender roles that determine internal threat points (e.g., sleeping on the couch, burnt toast, or "silent treatment") in a Nash-bargaining framework, which have distributional and efficiency consequences. In their separate-spheres

laws may alter the value of each spouse's options outside of marriage, which could have consequences for within-marriage investment.¹⁰

The return on investment in children is non-rival within marriage; however, the return is rival outside of marriage. Therefore, laws which govern the allocation of child custody could alter the expected value of divorce. Joint-custody reform may lower divorce costs for men because they expect to lose less of their return from child investment. Compared to a maternal preference custody regime, joint-custody reform may cause divorce costs to rise for women because they expect to receive less of a return on their investment in children.¹¹ If joint-custody reform shifts bargaining power to men, we expect household investments to reflect the preferences of men to a greater extent. Men may choose to invest more in children within marriage because they expect a greater return on their investment following divorce. Women may also have additional incentive to invest in their children in order to bind the marriage. Conversely, joint-custody reform could negatively affect child investment. Women may choose to invest less if the expected post-divorce return is lower. Men could also choose to invest less in the children because there is less incentive to bind the marriage.

Rasul (2006) develops a model of within-family bargaining in which child quality is a public good and married couples contract an *ex ante* allocation of child custody should

model, there is no outside option (i.e. divorce) and they make no assumptions regarding the efficiency of the equilibrium outcome. Since their model has no external threat point, joint-custody reform should have no consequences for intrafamily distribution.

¹⁰ Divorce-threat bargaining models assume a Pareto-efficient outcome.

¹¹ Brinig and Buckley (1998) find a negative relationship between joint-custody laws and state divorce rates.

divorce occur.¹² If spouses have homogenous preferences for child quality, joint custody is the optimal post-divorce custody allocation because it maximizes investment in the public good during marriage. However, if spouses have heterogeneous preferences for child quality, sole custody with the high-valuation spouse is the optimal child-custody allocation. If both spouses have an equally high valuation of child quality, we may observe a rise in the probability that a child attends private school when a state moves to joint custody. Alternatively, if one spouse values child quality more, we may either observe a positive or negative impact of the joint-custody laws. If the reform shifts bargaining power to the low-valuation spouse, the probability of children's private school attendance may decline. However, if bargaining power shifts to the high-valuation spouse, the rate at which parents invest in private school for their children may increase.

5.2.3. Joint Custody, Property Division, and Household Bargaining

Brinig and Allen (2000) find a significant increase (decrease) in the propensity of women to file for divorce when they (do not) expect to receive sole custody. Brinig and Allen's results suggest that the expectation of child custody is the most important factor in women's decisions to file for divorce. In the context of Brinig and Allen's findings, joint-custody reform should unambiguously shift bargaining power to fathers because the value of divorce would decrease for mothers. Hence, post-reform marital investment in

¹² Weiss and Willis (1985, 1993) also examine the allocation of child custody in the event of divorce; however, they only consider sole custody as a post-divorce child-custody allocation. Francesconi and Muthoo (2003) consider joint custody as an option and examine the child-specific investment behavior of parents. Their paper differs from Rasul's in several ways. For example, they consider cases in which the allocation of sole child custody to the low-valuation parent is optimal and divorce cannot occur in their model.

children may reflect the preferences of fathers to a greater extent. However, it is likely that the effects of joint-custody laws on marital investments in children are predicated on the underlying property-division laws, which may benefit either fathers or mothers.

In divorce settlements, Gray (1998) contends that men benefit more in common-law states and women benefit more in community-property states. In common-law states that enact joint custody, fathers may have additional incentive to invest in assets because the corresponding return has increased relative to child investment.¹³ Mothers in common-law states may have less incentive to invest in children because joint-custody reform has lowered their expected return to child quality relative to the return on investment in marital assets. Alternatively, mothers in these states may have additional incentive to invest in children to further bind their marriages rather than incur a loss of both assets and child custody.

Fathers in community-property states that enact joint custody may be more likely to send their children to private school because they will reap more benefits from child investment relative to their return on marital assets. On the other hand, joint-custody reform in community-property states could lower the return on child investment for mothers relative to the return on marital assets, as they may expect less time with their children after divorce.

¹³ Under a joint-custody regime, fathers may expect a larger share of child custody in the event of divorce. This could result in lower paternal investment in children, in which that investment was previously directed toward binding the marriage so as not to lose time with children. Under these circumstances, a law change which grants more custody to fathers in a state with property-division laws favoring fathers would likely increase the relative return to investment in marital assets.

In equitable-division states that adopt joint custody, the ways in which bargaining power shifts are less clear. However, equitable-division-property laws typically favor the spouse who is damaged the most by the divorce (Gray 1998). Weitzman (1985) and Hoffman and Duncan (1988) find that the economic well-being of divorced women falls by 73 percent and 30 percent, respectively. Assuming a decline in the well-being of women in the event of divorce, equitable-division property laws favor mothers. If joint-custody reform shifts bargaining power to fathers, children's private school attendance may increase in states with equitable-division property because fathers expect a greater return on child investment relative to that of assets. However, mothers may have additional incentive to invest in assets, as the expected return to child quality is diminished.

Joint-custody enactment in states with no-fault property division laws could shift bargaining power to the spouse who would otherwise be at fault. We may observe a decrease in children's private school attendance in these states because fathers who would otherwise be at fault have an additional incentive to invest in marital assets. However, if the mother were the otherwise at-fault parent, an increase in child investment, in order to further bind the marriage, may result because she stands to lose time with her children in the event of divorce. In fault-based states that adopt joint custody, at-fault spouses may have additional incentive to invest in children, as the return on investment in marital assets is relatively less.

5.3. DATA AND ECONOMETRIC METHODOLOGY

We use data from the 1980 and 1990 five-percent Integrated Public Use Microdata Series (IPUMS) from the U.S. Population Census and the child-custody law coding from Brinig and Buckley (1998a, see TABLE 1) to examine the effect of joint-custody laws on marriage-specific investments in children. We compare the child-specific investment behavior of households who live in states that change their custody laws between 1980 and 1990 with those from states that did not change their custody laws during this time. Married couples with children in states that change their child-custody laws to favor joint custody between 1980 and 1990 are the treatment group. Married couples with children who live in states that had either instituted joint-custody reform before 1980 or that did not institute joint-custody reform before 1990 are the comparison group. Because investment decisions made for children in blended families would most likely be the product of decisions made by their biological parents, one of whom is absent, we only examine in-tact households with own children present.

The dependent variable is children's private school attendance. The units of observation are households with children less than nine-years-old in grades one through four. IPUMS collapses the education variable such that children in kindergarten and those not enrolled in school are in the same category. As a result, we do not consider children in grades lower than first. Unfortunately, IPUMS also uses a collapsed variable for grades one through four in 1990. Because of this data limitation, we eliminate states that enact joint custody after 1983 so that the oldest child in our sample in 1990 would be

only one-year-old at the time of the last enactment.¹⁴ A sample of younger children allows us to create a treatment group of households with children who are not attending elementary school prior to the implementation of a joint-custody regime. However, selecting the sample based on the timing of school attendance and legal reform can also be problematic, as parents could decide on private school well before children are of school age. Estimates of joint-custody's effect on investment in children may be biased if parents make investment decisions prior to the law change. Hence, restricting the sample such that the youngest children are unborn and the oldest are one-year-old as of the last joint-custody reform also allows us to minimize the potential bias stemming from the long-run planning of parents.

With 21 states enacting joint custody between 1980 and 1984, there is substantial variation in the population who are affected by child-custody reform. In 1980, 32 percent of children in our sample (observations = 148,714) live in states with joint custody as the preferred custody allocation. The percentage of children in our sample (observations = 159,008) living in a state with joint custody as the preferred option in 1990 is 84 percent. In 1980, the percentage of children attending private school in states that have yet to enact joint custody is 14.11. By contrast, in 1990, children's private school attendance in states that enact joint custody declines slightly to 13.99 percent.¹⁵

¹⁴ We delete observations from Illinois, Maryland, Oregon, South Dakota, Tennessee, Texas, Utah, and Virginia from all empirical specifications. Since these states would have been treatment states, deleting these observations has no effect on the control states.

¹⁵ According to the 2000 U.S. Census (Table 247) approximately 2.7 million children (9.2 percent of the total population of elementary school students) attended private elementary school in 1990. Although we show a higher percentage of children attending private school, the difference likely comes from our sample consisting of only in-tact households with own children.

With census data from 1980 and 1990, we can control for time-invariant, unobserved heterogeneity at the state level. However, omitted time-varying, state-specific variables correlated with the passage of joint custody could bias estimates.¹⁶ For example, child custody was at the forefront of legislative agendas in the late 1970s and early 1980s because increased welfare receipts were attributed to delinquent child-support payments (Jacob 1988, Ch. 8). The spread of joint-custody reform could also be related to changing societal preferences for child-rearing responsibilities, as evidenced by the numerous fathers' rights groups who were politically prominent during this time (Jacob 1988, Ch. 8). To separate the effect of joint-custody reform from the influence of other aggregate variables, we control for time-varying, state-level demographic, economic, political, and welfare policy variables.

The empirical specification takes the probit functional form. We estimate the following equation:

$$Private_{i,s,t} = \beta_0 + \beta_1 Joint_{s,t} + \beta_2 C_{i,s,t} + \beta_3 P_{i,s,t} + \beta_4 S_{s,t} + \sum_s \eta_s + \sum_t \tau_t + \varepsilon_{i,s,t}. \quad (1)$$

The terms i , s , and t index children, states, and time, respectively. *Private* is an indicator variable that equals one if a child attends private school and equals zero if the child attends public school;¹⁷ *Joint* is an indicator variable that equals one if a state explicitly codifies or reveals its preference for joint custody in any year prior to the 1990 Census

¹⁶ Wolfers (2006) and Stevenson and Wolfers (2006) control for time-varying, state controls, finding that the effects of unilateral divorce on divorce rates and measures of family distress do not depend on other time-varying, state-level covariates. Alternatively, Stevenson (2008) shows that Gray's (1998) results, who examines the effects of unilateral divorce by the underlying property-division laws, are sensitive to the inclusion of time-varying, state-level controls.

¹⁷ We are unable to distinguish between various religious or parochial private schools because the U.S. Population Census survey questions on school type are not consistent across the two decennial periods.

year and zero otherwise; \mathbf{C} is a vector of child-specific controls; \mathbf{P} is a vector of parental controls; \mathbf{S} is a vector of time-varying, state-level controls; η and τ are state and time fixed effects; ε is an error term; and the β_i are parameters to be estimated. The variables in \mathbf{C} are the child's age, a squared term of their age, race, gender, and whether the child lives in a city, and those in \mathbf{P} are parents' age, race, and education level. The variables in \mathbf{S} include the age and racial composition of the population, the unemployment rate, real per-capita income, a measure of the extent to which a state's congressional delegation cast liberal votes, political party of the governor, the consideration of marital fault in the divorce settlement, the Aid to Families with Dependent Children (AFDC) participation rate and maximum benefit, the value of Food Stamp outlays, and the Supplemental Security Income (SSI) participation rate.¹⁸ The inclusion of \mathbf{S} allows us to minimize the potential bias from a spurious correlation between joint-custody reform and other state-level characteristics.¹⁹ Summary statistics and formal variable definitions for the controls in \mathbf{C} and \mathbf{P} are shown in TABLE 2 and those in \mathbf{S} are shown in TABLE 3.

We also examine the potential tradeoff between child investment and other marriage-specific assets. To estimate these effects, we estimate the effects of joint-custody reform by the type of property-division laws in place across states. We use Gray's (1998) property-division law coding found in TABLE 1. Because there is limited variation in

¹⁸ See Appendix for sources of the state-level controls.

¹⁹ It is not necessary to control for the property-division laws, as they are time invariant. We capture the effects of the property-division laws by including state fixed effects. It may also be necessary to control for unilateral divorce laws, as they could be correlated with joint-custody reform. Only South Dakota and Utah changed their divorce laws to favor unilateral divorce between 1980 and 1990. Both South Dakota and Utah also began to favor joint custody in 1987; therefore, these states would be excluded. Therefore, the potential impact of the divorce-law regime in place is removed by including state fixed effects.

joint-custody reform among states with certain property-division laws, the estimated interaction effects are more likely to capture a spurious relationship between time-varying, state-level variables and child-custody reform. This may not present a problem for the interaction between joint-custody reform and equitable-division property, because 13 of the 28 equitable-division states enact joint custody between 1980 and 1984. However, of the 14 common-law states, only four states enact joint custody between 1980 and 1984. Identification of these effects also relies on S to control for state-level changes correlated with the passage of joint custody and/or the underlying property-division laws.

The equation testing the potential tradeoff between investment in children and other marital assets is

$$\begin{aligned} Private_{i,s,t} = & \alpha_0 + \alpha_1 Joint\ with\ Community_{s,t} + \alpha_2 Joint\ with\ Commom_{s,t} & (2) \\ & + \alpha_3 Joint\ with\ Equitable_{s,t} + \alpha_4 C_{i,s,t} + \alpha_5 P_{i,s,t} + \alpha_6 S_{s,t} \\ & + \sum_s \eta_s + \sum_t \tau_t + \varepsilon_{i,s,t}. \end{aligned}$$

We define the terms i , s , and t , and variables $Private$, C , P , S , η , τ , and ε above. The variables $Joint\ with\ Community$, $Joint\ with\ Common$, and $Joint\ with\ Equitable$ equal one when community-property, common-law, and equitable-division states adopt joint custody and zero otherwise, respectively. The α_i are parameters to be estimated.

In the final specification, we estimate separate effects for the adoption of joint custody with fault-based property division and the adoption of joint custody without

fault-based property division. We use Ellman and Lohr's (1998) no-fault property coding found in TABLE 1.²⁰ This equation is

$$\begin{aligned} Private_{i,s,t} = & \gamma_0 + \gamma_1 Joint\ with\ Fault_{s,t} + \gamma_2 Joint\ with\ No\ Fault_{s,t} \\ & + \gamma_3 C_{i,s,t} + \gamma_4 P_{i,s,t} + \gamma_5 S_{s,t} + \sum_s \eta_s + \sum_t \tau_t + \varepsilon_{i,s,t}. \end{aligned} \quad (3)$$

We define the terms i , s , and t and variables C , P , S , η , τ , and ε above. The variables *Joint with Fault* and *Joint without Fault* equal one when a fault-based property-division and no-fault property-division state enacts joint custody and zero otherwise, respectively. The γ_i are parameters to be estimated.

In order to estimate the full effect of joint-custody reform, we do not control for covariates that may be affected by the reform, as in Stevenson (2007). For example, household income could be affected by child-custody reform because the reform may change spousal labor-supply decisions. Because the reform may change investments in children, it could also alter fertility decisions. If the reform affects household income and/or the number of children, then at least a portion of the effect of the joint-custody laws would be removed because its impact would also be captured in the estimates for household income and/or the number of children (Lee 2005).

5.4. RESULTS

In each empirical specification, we successively add controls to check the sensitivity of the estimates. We estimate six specifications. Our first specification includes state and

²⁰ There are 25 states that removed fault as a consideration in divorce settlements, with only four changing their laws between 1980 and 1990. Our sample only has two states that change from fault-based to no-fault-property division because we delete Florida and Utah. Both Florida and Utah enact joint custody after 1983. Of the 25 no-fault-property states, 11 enact joint custody between 1980 and 1984.

time fixed effects and child-specific controls for sex, race, age, age-squared, and whether the child lives in a city. In the second specification, we add controls for parents' age, race, and education level. For the third and subsequent specifications, we add controls for time-varying, state-specific effects.²¹ In our third specification, we include demographic controls including the age and racial composition of the respondent's state. The controls for age include the percentage of the population under 5, 5-17, 18-24, 25-44, 45-64, and 65 and over. The racial variables include the percentage of the population that is white, black, and other. The fourth specification adds controls for changes in economic conditions including the state unemployment rate and real per-capita income. Next, we add controls for state-specific, political characteristics including whether the governor is a democrat and measures for the degree to which state's congressional delegation casts liberal votes. Our final specification adds controls for state-level welfare policy changes. The state-policy variables are the maximum Aid to Families with Dependent Children (AFDC) benefit paid to families of four, the AFDC participation rate, the value of Food Stamp payments, the Supplemental Security Income (SSI) participation rate, and whether fault is a consideration in the divorce settlement. For equation (3), we control for no-fault property division. In these models, the omitted category is fault-based property division.

²¹ Stevenson (2008) shows that the inclusion of time-varying, state-level controls may be important when considering the effects of family-law reform. In fact, Stevenson (2008) shows that Gray's (1998) findings are sensitive to the inclusion of time-varying, state-level controls. Gray finds no effect of unilateral divorce on female labor-force participation; however, when Gray considers the underlying property-division laws, the effects become statistically significant, and the directional impact of the effects depends on the property-division law in place. Stevenson (2008) finds that unilateral divorce reform increases both married and unmarried women's labor-force participation after controlling for a variety of time-varying, state-level variables that seem to be correlated with the adoption of unilateral divorce and/or the property-division laws.

Our first model takes the form of equation (1) and considers the impact of the joint-custody laws on children's private school attendance. TABLE 4 shows these estimates. In Models 1 through 5, the effect of joint-custody reform on children's private school attendance is not statistically significant. However, the directional impact of the estimated effects changes once we successively include additional state-level controls. For example, when we control for state-level, demographic and economic variables the estimated effect becomes positive, but then becomes negative once we control for state-level, political variables. The reform's effect on children's private school attendance remains negative but becomes marginally statistically significant in Model 6 when we include state-level, and policy changes to welfare, which indicates a 9.50 percent (a 1.17 percentage point decrease) decline in the probability of children's private school attendance.

Our next specification, which takes the form of equation (2), considers the effects of joint-custody reform by the type of property-division law in place across states. This model tests for a tradeoff between child investment and other marital investments. TABLE 5 presents these results. We find that the effects of joint-custody reform in both common-law and community-property states are negative and statistically significant at the one-percent confidence level in Model 6 only. The directional impact and statistical significance of joint-custody reform by property division are highly sensitive to the inclusion of time-varying, state-level controls, with the exception of common-law states. For example, in Model 1, joint-custody reform in community-property states is positive and marginally statically significant, with the effect remaining positive in Model 2 but

becoming statistically insignificant. In Model 3, the estimated effect becomes negative and remains negative in subsequent models. However, the effect only becomes statistically significant in Model 6, which suggests that the estimates are especially responsive to the inclusion of welfare policy controls. In community-property and common-law states that enact joint custody, the probability of children's private school attendance declines by 20.64 and 18.25 percent (2.48 and 2.96 percentage point decreases), respectively. The effect of joint-custody reform in equitable-division states is not statistically discernable from zero. The marginal effects of joint-custody reform in Model 6 of TABLE 5 indicate negative consequences for children in states with property-division laws that are consistently favorable to one spouse. Because men benefit in common-law states and women benefit in community-property states, the sizeable, negative effects found in community-property and common-law states suggest that spouses invest more in marital assets when the property-division laws favor one spouse over the other.

Our final specification, taking the form of equation (3), examines the impact of joint-custody reform for states that do and do not consider marital wrongdoing in the division of marital assets. TABLE 6 shows these results. As in TABLES 4 and 5, these estimates are also highly sensitive to additional state-level control variables. In Model 3, joint-custody reform in fault-based states has a statistically significant, positive effect on the probability of children's private school attendance; however, the effect remains positive, but becomes statistically insignificant in Models 4, 5 and 6. For children in no-fault states, the estimated effect is positive but statistically insignificant in Models 1 and 2,

with negative effects found in subsequent models. In Models 4 and 6, the estimated negative effect is statistically significant at the ten-percent confidence level, both with similar sized coefficient estimates. In Model 6, the probability of children's private school attendance declines by 12.72 percent (a 1.57 percent decrease) in no-fault states that adopt joint custody. The negative effect found for joint-custody reform in no-fault states suggests that the otherwise at-fault spouse has additional incentive to invest in assets instead of child quality.

We also partition the sample by the household's SES, as child-custody reform may affect families of varying SES differently. We use mother's educational attainment as a measure of SES. A comparison of the sample's descriptive statistics suggests that children's private school attendance varies greatly by mother's education: 7.37 percent for mothers who are high school dropouts; 13.45 percent for those who are high school graduates and those with some college; and 19.74 percent for those who are college graduates. We compare descriptive statistics on children's private school attendance by mother's education between states with and without joint custody in TABLE 7. Children's private school attendance is greater in joint-custody states for children of mothers who are high-school and/or college graduates. For children of the lowest SES, private school attendance is lower in joint-custody states.

TABLE 8 presents estimates by SES for the final specifications shown in TABLES 4, 5, and 6.²² For the lowest SES group, the impact of joint-custody reform on children's

²² The specifications for Models 1, 4, and 7, Models 2, 5, and 8, and Models 3, 6, and 9 are analogous to the final models in Tables 4, 5, and 6, respectively.

private school attendance is negative and statistically significant at the one-percent confidence level in Model 1, which indicates a 45.19 percent decrease (a 3.7 percentage point decline). In Model 2, we investigate the impact of joint-custody reform by the property-division regime in place across states (i.e. equation (2)), and find 70.82, 31.14, and 31.26 percent decreases (5.2, 2.2, and 2.4 percentage point declines) in community-property, common-law, and equitable-division states, respectively. Model 3 presents the results from equation (3) for the lowest SES group. The effects of joint-custody reform are negative and statistically significant, regardless of whether or not marital fault is a consideration in the divorce settlement. However, the percentage decrease in children's private school attendance is slightly larger in fault-based states than no-fault states, with a percent decline of 49.04 and 47.91 (3.3 and 3.1 percentage point decreases), respectively. The results from Models 1, 2, and 3 for the lowest SES group indicate that joint-custody reform negatively affects the probability of children's private school attendance regardless of how the state divides marital property.

For children of high-school graduates and those with some college, joint-custody reform is not statistically significant in Model 4. However, the effects of joint-custody reform are negative and statistically significant in community-property and common-law states. In community-property and common-law states that enact joint custody, the probability of children's private school attendance declines by 14.01 and 22.43 percent (2.4 and 2.5 percentage point declines), respectively. For this SES group, the estimated effects are not statistically discernable from zero when we examine the impact of joint-

custody reform in states with and without the consideration of marital fault in the divorce settlement.

In Model 7, the effect of joint-custody reform on the probability of children's private school attendance is not statistically significant. However, in Model 8, we find a 21.28 percent decrease (a 6.18 percentage point decline) in the probability of children's private school attendance in common-law states that enact joint custody. We also find a 16.75 percent decline (a 3.4 percentage point decline) in the probability of children's private school attendance in no-fault states that enact joint custody.

Because there is some debate on the classification of states' property-division laws, we check the robustness of our results to alternative property-law codings. Gray (1998, pp. 632, footnote 3) suggests that five states could be characterized as having equitable-division property laws: Arizona, Idaho, Nevada, Texas, and Washington. Similarly, Bring and Buckley (1998b)'s law coding for the consideration of marital fault in the divorce settlement differs substantially from Ellman and Lohr's (1998). We find that the results shown in TABLES 5, 6, and 8 are largely robust to these alternative property-law codings. However, in a few cases, we find slight discrepancies in the size and statistical significance of the estimated effects.²³ The robustness checks are shown in TABLES A1-A3 in the Appendix.

²³ In Model 6 of Table A1, the estimated effect for joint-custody reform in community-property states is statistically significant at the five-percent level. By contrast, in Table 5, we find that the effect is statistically significant at the one-percent level. In Model 6 of Table A2, we find that joint-custody reform in no-fault states is statistically significant at the five-percent level; however, in Model 6 of Table 6 we find the effect is statistically significant at the ten-percent level. In the models partitioned by SES, we only find one minor discrepancy. The effect of joint-custody reform in community-property states is statistically significant at the one-percent level in Table A3, with the estimated effect being statistically significant at

5.5. CONCLUSIONS

This is the first study to investigate the effects of joint or shared child-custody laws on marriage-specific investment in child quality. We use variation in child-custody laws across states and time to identify the effects of joint-custody reform on children's private school attendance. Although most children in the U.S. do not attend private school, observed differences in private school attendance can likely be generalized to other forms of child investment. We find a marginally statistically significant, 9.50 percent decrease in the probability that a child attends private school in states that enact joint custody. When we partition the sample by SES, we only find a negative, statistically significant effect of joint-custody reform for the lowest SES group.

We also consider the potential tradeoff between investing in marital assets and child quality. In both community-property and common-law states that enact joint custody, the probability of children's private school attendance declines by 20.64 and 18.25 percent, respectively. An economic explanation of these sizable negative effects is that spouses invest less in their children when they stand to gain more of the marital surplus in the event of divorce. Dividing the sample by SES for these models yields many interesting results. For the lowest SES group, we observe a decrease in the probability of children's private school attendance regardless of the underlying property-division laws in states

the five-percent level in Table 8. The estimated marginal effect is also slightly larger. Although the statistical significance differs slightly with different law codings, the overall effects are similar.

that enact joint custody. In common-law states, joint-custody reform reduces the probability of children's private school attendance for all SES groups.

We also consider the effects of joint-custody reform in states that do and do not consider marital fault in the consideration of divorce settlements. For the full sample, we find some statistical evidence of a decline in the probability of children's private school attendance in no-fault states that enact joint custody. For the highest SES group, the effect of joint-custody reform on private school attendance in no-fault states is negative and statistically significant.

Judges in the U.S. are directed to consider the best interests of the child in the adjudication of child-custody cases. To that end, joint-custody reform may lessen the impact for children of losing regular contact with one of their parents. However, the prospect of post-divorce cooperation under a joint-custody regime may have negative within-marriage consequences regarding child investment. The incentives to invest in children and other marital assets could be predicated on the potential return to those investments in the event of divorce. According to our results, the effect of joint-custody reform on marital investment in children also depends on state laws specifying the division of marital assets. Further consideration of how joint-custody laws alter child-investment incentives within married households could help avoid negative, albeit unintended, consequences for children.

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TABLE 1—YEAR OF INTRODUCTION OF JOINT-CUSTODY LAWS AND THE PREVAILING PROPERTY-DIVISION LAWS BY STATE

State	Joint Custody	Property Division	No-Fault Property	State	Joint Custody	Property Division	No-Fault Property
Alabama		Common		Montana	1981	Equitable	1975
Alaska	1982	Equitable	1974	Nebraska	1983	Equitable	1972
Arizona		Community	1973	Nevada	1981	Community	1973
Arkansas		Equitable	1979	New Hampshire	1974	Equitable	
California	1979	Community	1970	New Jersey	1981	Equitable	1980
Colorado	1983	Equitable	1971	New Mexico	1982	Community	1976
Connecticut	1981	Equitable		New York	1981	Common	
Delaware	1981	Equitable	1974	North Carolina	1979	Common	
Florida	1979	Common	1986	North Dakota		Equitable	
Georgia	1990	Common		Ohio	1981	Common	
Hawaii	1980	Equitable	1960	Oklahoma	1990	Equitable	1975
Idaho	1982	Community	1990	Oregon	1987	Equitable	1971
Illinois	1986	Equitable	1977	Pennsylvania	1981	Common	
Indiana	1973	Equitable	1973	Rhode Island		Common	
Iowa	1977	Equitable	1972	South Carolina		Common	
Kansas	1979	Equitable	1990	South Dakota	1989	Equitable	
Kentucky	1979	Equitable		Tennessee	1986	Common	
Louisiana	1981	Community		Texas	1987	Community	
Maine	1981	Equitable	1985	Utah	1988	Equitable	1987
Maryland	1984	Common		Vermont		Equitable	
Massachusetts	1983	Equitable		Virginia	1987	Common	
Michigan	1981	Equitable		Washington		Community	1973
Minnesota	1981	Equitable	1974	West Virginia		Common	
Mississippi	1983	Common		Wisconsin	1979	Equitable	1977
Missouri	1983	Equitable		Wyoming		Equitable	

Notes: Data for the child-custody laws are from Brinig and Buckley (1998a). Property-division laws are from Gray (1998) and the no-fault property-division laws are from Ellman and Lohr (1998).

TABLE 2—SUMMARY STATISTICS FOR CHILD, MOTHER, AND FATHER CONTROLS

Variable	Variable Description	Mean	Std. Dev.
Private School Attendance	= 1 if child attends private school	0.1340	0.3341
<i>Child Covariates:</i>			
Gender	= 1 if child is a male	0.5004	0.5000
Hispanic	= 1 if child is Hispanic	0.0911	0.2878
Black	= 1 if child is black	0.0754	0.2641
City	= 1 if child lives in a city	0.1331	0.3397
<i>Mother Covariates:</i>			
Age	In years	31.014	5.3260
Hispanic	= 1 if mother is Hispanic	0.0872	0.2822
Black	= 1 if mother is black	0.0732	0.2605
High School	= 1 if mother has only a high-school degree	0.4291	0.4949
Some College	= 1 if mother has attended college with no degree	0.2422	0.4284
College Graduate	= 1 if mother is a college graduate	0.1572	0.3640
<i>Father Covariates:</i>			
Age	In years	36.660	6.3622
Hispanic	= 1 if father is Hispanic	0.0868	0.2816
Black	= 1 if father is black	0.0760	0.2650
High School	= 1 if father has only a high-school degree	0.3506	0.4772
Some College	= 1 if father has attended college with no degree	0.2382	0.4260
College Graduate	= 1 if father is a college graduate	0.2364	0.4249

Notes: There are 341,922 observations for all variables. For mothers' and fathers' education level, those with less than a high school education are the omitted category.

TABLE 3—SUMMARY STATISTICS FOR THE TIME-VARYING STATE-LEVEL VARIABLES

Variable	Variable Description	Mean	Std. Dev.
Demographic			
Age (under 5)	Percentage of the population under 5 years of age	0.0737	0.0068
Age (5 to 17)	Percentage of the population aged 5 years to 17 years	0.1943	0.0191
Age (18 to 24)	Percentage of the population aged 18 years to 24 years	0.1196	0.0139
Age (25 to 44)	Percentage of the population aged 25 years to 44 years	0.3000	0.0288
Age (45 to 65)	Percentage of the population aged 45 years to 64 years	0.1918	0.0135
Age (over 65)	Percentage of the population over 65 years of age	0.1210	0.0203
Black	Percentage of the population who is black	0.1127	0.0783
White	Percentage of the population who is white	0.8531	0.0856
Other	Percentage of the population who are <i>not</i> black or white	0.0342	0.0571
Economic			
Unemployment Rate	Percentage of the population unemployed	6.4349	1.5672
Per Capita Income	Average real personal income	14,679	5,149
Political			
Liberal Quotient (House)	The degree to which a state's House of Representatives casts liberal votes	0.4932	0.1599
Liberal Quotient (Senate)	The degree to which a state's Senate casts liberal votes	0.5195	0.2393
Governor	=1 if the governor is a Democrat	0.5923	0.4914
Policy			
No-Fault Property	=1 if the state does not consider marital fault in the divorce settlement	0.4213	0.4938
AFDC Benefit	Dollar amount of the maximum AFDC benefit paid to families of four	576.65	218.65
AFDC Participation	AFDC participation rate	12.069	3.4964
Food Stamp Value	Dollar amount of Food-Stamp outlays	588.67	410.82
Social Security Rate	Supplemental Security Income participation rate	1.9502	0.8144

TABLE 4—CHILD-CUSTODY LAWS AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
<i>Joint Custody</i>	-0.0035 (0.0056)	-0.0051 (0.0054)	0.0053 (0.0045)	0.0016 (0.0057)	-0.0022 (0.0055)	-0.0117* (0.0071)
<i>Controls:</i>						
Child	X	X	X	X	X	X
Parental		X	X	X	X	X
State Demographic			X	X	X	X
State Economic				X	X	X
State Political					X	X
State Policy						X
Number of Observations	307,722	307,722	307,722	307,722	307,722	307,722
Pseudo R-squared	0.0468	0.0735	0.0739	0.0740	0.0740	0.0741
Log Pseudo-Likelihood	115,513	112,288	112,232	112,228	112,217	112,207

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. Each model contains state and year fixed effects. We cluster our standard errors at the state-year level (See Bertrand et al. 2004).

TABLE 5—CHILD-CUSTODY LAWS, PROPERTY-DIVISION LAWS, AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
<i>Joint Custody with Community-Property Division</i>	0.0071* (0.0039)	0.0058 (0.0036)	-0.0018 (0.0068)	-0.0091 (0.0078)	-0.0113 (0.0076)	-0.0248*** (0.0084)
<i>Joint Custody with Common-Law Property</i>	-0.0130* (0.0069)	-0.0140** (0.0069)	-0.0032 (0.0055)	-0.0159** (0.0077)	-0.0215*** (0.0080)	-0.0296*** (0.0065)
<i>Joint Custody with Equitable-Division Property</i>	0.0054 (0.0055)	0.0032 (0.0053)	0.0120*** (0.0043)	0.0087** (0.0043)	0.0046 (0.0042)	-0.0001 (0.0051)
<i>Controls:</i>						
Child	X	X	X	X	X	X
Parental		X	X	X	X	X
State Demographic			X	X	X	X
State Economic			X	X	X	X
State Political					X	X
State Policy						X
Number of Observations	307,722	307,722	307,722	307,722	307,722	307,722
Pseudo R-squared	0.0470	0.0736	0.0740	0.0740	0.0741	0.0742
Log Pseudo-Likelihood	115,496	112,272	112,227	112,217	112,206	112,197

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. Each model contains state and year fixed effects. We cluster our standard errors at the state-year level (See Bertrand et al. 2004).

TABLE 6—CHILD-CUSTODY LAWS, PROPERTY-DIVISION LAWS, NO-FAULT PROPERTY-DIVISION LAWS,
AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
<i>Joint Custody with Fault-Based Property Division</i>	-0.0045 (0.006)	-0.0066 (0.006)	0.0091** (0.005)	0.0058 (0.006)	0.0033 (0.006)	-0.0086 (0.008)
<i>Joint Custody with No-Fault Property Division</i>	0.0027 (0.006)	0.0043 (0.005)	-0.0082 (0.007)	-0.0158* (0.009)	-0.0123 (0.012)	-0.0157* (0.008)
<i>Controls:</i>						
Child	X	X	X	X	X	X
Parental		X	X	X	X	X
State Demographic			X	X	X	X
State Economic				X	X	X
State Political					X	X
State Policy						X
Number of Observations	307,722	307,722	307,722	307,722	307,722	307,722
Pseudo R-squared	0.0470	0.0735	0.0739	0.0740	0.0741	0.0741
Log Pseudo-Likelihood	115,512	112,284	112,229	112,222	112,211	112,207

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. Each model contains state and year fixed effects. We cluster our standard errors at the state-year level (See Bertrand et al. 2004).

**TABLE 7—SUMMARY STATISTICS FOR CHILDREN’S PRIVATE SCHOOL ATTENDANCE
BY MOTHER’S EDUCATION AND JOINT-CUSTODY REGIME**

Mother’s Education	Joint Custody=0			Joint Custody=1		
	Mean	Std. Dev.	Obs.	Mean	Std. Dev.	Obs.
High School Dropouts	0.0819	0.2742	24,137	0.0667	0.2496	28,646
High School Graduates and Some College	0.1266	0.3325	157,235	0.1328	0.3394	181,358
College Graduates	0.1833	0.3869	17,067	0.2051	0.4038	31,307

TABLE 8—CHILD-CUSTODY LAWS, PROPERTY-DIVISION LAWS, NO-FAULT PROPERTY-DIVISION LAWS, AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL BY MOTHER'S EDUCATIONAL ATTAINMENT

Variable	High School Dropout			High School Graduate and Some College			College Graduate		
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9
<i>Joint Custody</i>	-0.0371*** (0.0086)			-0.0067 (0.0082)			-0.0119 (0.0113)		
<i>Joint Custody with Community Property</i>		-0.0524*** (0.0070)			-0.0236** (0.0098)			0.0024 (0.0168)	
<i>Joint Custody with Common Law</i>		-0.0215*** (0.0068)			-0.0248*** (0.0076)			-0.0618*** (0.0125)	
<i>Joint Custody with Equitable Division</i>		-0.0240*** (0.0050)			0.0052 (0.0062)			-0.0049 (0.0121)	
<i>Joint Custody with Fault-Based Property</i>			-0.0313*** (0.0076)			-0.0041 (0.0091)			0.0045 (0.0132)
<i>Joint Custody with No-Fault Property</i>			-0.0328*** (0.0092)			-0.0102 (0.0106)			-0.0342*** (0.0139)
Number of Obs.	52,783	52,783	52,783	206,563	206,563	206,563	48,376	48,376	48,376
Pseudo R-squared	0.1090	0.1093	0.1090	0.0658	0.0659	0.0658	0.0485	0.0487	0.0485
Log Pseudo-Likelihood	12,371	12,367	12,371	76,188	76,180	76,188	22,867	22,863	22,866

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. All models include child, parental, father's education, and time-varying, state-level controls used in the final models in TABLES 4, 5, and 6. We cluster our standard errors at the state-time level.

APPENDIX

DATA SOURCES:

The *Demographic* variables come from the United States (U.S.) Census and the Center for Disease Control (CDC): <http://www.census.gov/popest/archives/1980s/>, <http://www.census.gov/popest/archives/1990s/>, and <http://wonder.cdc.gov/Census.html>.

The *Economic* variables come from the Bureau of Labor Statistics (BLS) and www.economagic.com: <http://www.census.gov/prod/99pubs/99statab/sec13.pdf> and <http://www.economagic.com/beapira.htm>.

The *Political* variables come from <http://www.adaction.org/votingrecords.htm> and U.S. Almanacs.

The *Policy* variables come from <http://aspe.hhs.gov/hsp/indicators07/apa.pdf> and the Department of Health and Human Services.

TABLE A1—CHILD-CUSTODY LAWS, PROPERTY-DIVISION LAWS, AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL (ROBUSTNESS CHECK)

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
<i>Joint Custody with Community Property</i>	0.0075* (0.0039)	0.0053 (0.0036)	0.0011 (0.0065)	-0.0066 (0.0078)	-0.0073 (0.0076)	-0.0259** (0.0096)
<i>Joint Custody with Common Law</i>	-0.0130* (0.0070)	-0.0140* (0.0069)	-0.0027 (0.0054)	-0.0159** (0.0078)	-0.0212** (0.0082)	-0.0299*** (0.0067)
<i>Joint Custody with Equitable Division</i>	0.0054 (0.0054)	0.0034 (0.0052)	0.0105** (0.0045)	0.0069 (0.0045)	0.0026 (0.0044)	-0.0017 (0.0053)
<i>Controls:</i>						
Child	X	X	X	X	X	X
Parental		X	X	X	X	X
State Demographic			X	X	X	X
State Economic				X	X	X
State Political					X	X
State Policy						X
Number of Observations	307,722	307,722	307,722	307,722	307,722	307,722
Pseudo R-squared	0.0470	0.0736	0.0739	0.0740	0.0741	0.0742
Log Pseudo-Likelihood	115,496	112,271	112,228	112,218	112,208	112,197

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. Each model contains state and year fixed effects. We cluster our standard errors at the state-year level, as Bertrand et al. (2004) suggest.

TABLE A2—CHILD-CUSTODY LAWS, PROPERTY-DIVISION LAWS, NO-FAULT PROPERTY-DIVISION LAWS,
AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL
(ROBUSTNESS CHECK)

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
<i>Joint Custody with Fault-Based Property Division</i>	-0.0041 (0.0062)	-0.0058 (0.0059)	0.0096** (0.0046)	0.0063 (0.0057)	0.0033 (0.0057)	-0.0084 (0.0078)
<i>Joint Custody with No-Fault Property Division</i>	0.0011 (0.0059)	0.0028 (0.0048)	-0.0099 (0.0072)	-0.0167* (0.0086)	-0.0146 (0.0112)	-0.0176** (0.0075)
<i>Controls:</i>						
Child	X	X	X	X	X	X
Parental		X	X	X	X	X
State Demographic			X	X	X	X
State Economic				X	X	X
State Political					X	X
State Policy					X	X
Number of Observations	307,722	307,722	307,722	307,722	307,722	307,722
Pseudo R-squared	0.0468	0.0735	0.0740	0.0740	0.0741	0.0741
Log Pseudo-Likelihood	115,512	112,285	112,227	112,221	112,211	112,206

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. Each model contains state and year fixed effects. We cluster our standard errors at the state-year level, as Bertrand et al. (2004) suggest.

TABLE A3—CHILD-CUSTODY LAWS, PROPERTY-DIVISION LAWS, NO-FAULT PROPERTY-DIVISION LAWS,
AND MARRIED COUPLES' INVESTMENT IN CHILDREN'S PRIVATE SCHOOL BY MOTHER'S EDUCATIONAL ATTAINMENT

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9
<i>Joint Custody</i>	-0.0371*** (0.0086)			-0.0067 (0.0082)			-0.0119 (0.0113)		
<i>Joint Custody with Community Property</i>		-0.0557*** (0.0074)			-0.0303*** (0.0104)			0.0258 (0.0166)	
<i>Joint Custody with Common Law</i>		-0.0216*** (0.0067)			-0.0258*** (0.0075)			-0.0601*** (0.0129)	
<i>Joint Custody with Equitable Division</i>		-0.0248*** (0.0048)			0.0046 (0.0062)			-0.0007 (0.0122)	
<i>Joint Custody with Fault-Based Property</i>			-0.0287*** (0.0070)			-0.0037 (0.0092)			0.0032 (0.0130)
<i>Joint Custody with No-Fault Property</i>			-0.0367*** (0.0094)			-0.0124 (0.0093)			-0.0357*** (0.0123)
Number of Obs.	52,783	52,783	52,783	206,563	206,563	206,569	48,376	48,376	48,376
Pseudo R-squared	0.1090	0.1093	0.1091	0.0658	0.0660	0.0659	0.0485	0.0487	0.0485
Log Psuedo- Likelihood	12,371	12,367	12,369	76,188	76,179	76,187	22,869	22,863	22,868

Notes: Standard errors are in parentheses. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively. All models include child, parental, father's education, and time-varying, state-level controls used in the final models in TABLES 4, 5, and 6. As Bertrand et al. (2004) suggest, we cluster our standard errors at the state-time level.

CHAPTER 6

CONCLUSION

The first essay of this dissertation is entitled, “The Effects of Household Income Volatility on Divorce.” While research on the impact of income and income shocks is widely investigated in the economics literature, no study has considered the interdependence of spouses' incomes. This essay considers the interdependence of spousal incomes by considering whether household income volatility affects individual-level divorce propensities. Negative shocks to household income raise the probability of divorce for both men and women, regardless of the level of household income. The impact of positive household income shocks is not consistent for men and women, and is not robust across lower- and higher-levels of household income. Positive income volatility has a statistically significant positive effect on the divorce propensities of men, while no such effect is found for women. Increases in household income volatility raise the probability of divorce for individuals in the higher-household income groups. By contrast, positive shocks to household income lower the divorce risk for the lower-household income group. Consistent across all models is the role of negative household income volatility, indicating a rise in the individual divorce propensities.

The second essay, “Inflation and Other Aggregate Determinants of the Trend in U.S. Divorce Rates since the 1960s,” focuses on a neglected macroeconomic variable's effect on the divorce rate: the inflation rate. The traditional family model predicts that the returns from

marriage when spouses specialize in market and household work. Inflation worsens the terms of trade between spouses, as time may be shifted away from leisure or household work in order to achieve pre-inflation, consumption levels. Consistent with this prediction, I find that increases in the inflation rate have persistent, positive effects on the divorce rate. I also investigate whether the unemployment rate, the growth rate of Gross Domestic Product (GDP), and changes in female participation in higher education affect the divorce rate. The results for the unemployment rate are mixed, with some models indicating a positive effect and others a negative effect. The results support previous research on the relationship between the growth rate of GDP and the divorce rate, finding a positive relationship. Increases in female participation in higher education lead to a rise in the divorce rate; however, the effect is small. Another way this essay extends empirical research on the divorce rate is by using the structural time-series methodology. Though not used often, this approach circumvents problems associated with trended variables by modeling the divorce rate as an unobserved variable. This makes for unbiased parameter estimates.

The third essay entitled, "Explaining the Evolution of the U.S. Divorce Rate." This essay extends empirical research on the divorce rate in a number of ways. First, we examine whether increased access to oral contraception in the 1960s and 1970s had a measurable impact on the divorce rate. Second, we objectively take into account the effects of major wars, including World War II (WWII), the Korean War, and the Vietnam War. We also revisit the relationship between the divorce rate and the female labor-force participation rate or its proxy, female participation in higher education. Third, we extend the analysis back to 1929. We show that research too narrowly focused on the 1960s and 1970s necessarily

identify a positive relationship between the divorce rate and female participation in the labor market or higher education. Failure to investigate a longer sample hides two distinct negative relationships: one before and after WWII and another from the late-1970s onwards. The sharp rise in the divorce rate during the 1960s and 1970s is marked by the diffusion of oral contraception, increased access to divorce, and the Vietnam War. Our results also indicate that increased access to oral contraception and the Vietnam War shifted the divorce rate to higher level from the early-1960s to the late-1970s. The longer sample period, accounting for legal changes permitting increased access to oral contraception and divorce, and including a measureable variable for the Vietnam War reveals new insights into the determinants of the rise in the divorce rate from the mid-1960s to the mid-1970s and the relationship between the divorce rate and rising economic power of women.

The fourth essay, "Child-Custody Reform and Marriage-Specific Investment in Children," examines whether the adoption of joint custody in many states in the early-1980s affected how married parents invest in their children. Prior to joint-custody reform, states had an explicit preference for mothers in child-custody cases. As such, joint-custody reform redefines the division of children in the event of divorce. Divorce-threat bargaining models posit that legal changes (or environmental factors) that alter the division of the marital surplus affect within-marriage distribution. This essay investigates whether the prospect of joint child custody alters marital investment in children, measured as children's private school attendance. The findings indicate negative consequences for children living in states that enact joint custody relative to those who live in states with a sole-custody regime. When we consider the effects of the reform by the underlying property-division laws, the effects

remain negative but become much larger in states that have property-division laws that favor one spouse over another.