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THE U.S. DIVORCE RATE: THE 1960S SURGE VERSUS ITS LONG-RUN DETERMINANTS

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This study investigates the determinants of the U.S. divorce rate from 1929 to 2006, with particular emphasis on explaining its surge in the mid-1960s. The main finding is that the divorce rate and female labor-force participation, or equivalently female participation in higher education, are endogenous variables that are linked by a negative, long-run relationship. The availability of oral contraception shifted this negative relationship to a new, higher level of divorce rates during the late-1960s and early-1970s. The Vietnam War also contributed to the rise in the divorce rate at that time. The results are very robust to different estimation methodologies.

JEL Categories: J11, J12

Key words: divorce rate, female labor-force participation, female participation in higher education, oral contraceptives, unilateral divorce laws, Vietnam War

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1 Introduction

The steady rise in the United States (U.S.) divorce rate from the mid-1960s to the mid-1970s has been a topic of much debate among demographers and economists (Michael 1978; Ruggles 1997; Friedberg 1998; Goldstein 1999; Wolfers 2006). Researchers have focused on explaining the evolution of the divorce rate primarily by changes in the female labor-force participation rate (FLFPR) and in divorce laws. However, there is sufficient evidence that causality may run from divorce to a rise in the FLFPR as opposed to the other way around (Johnson and Skinner 1986; Sen 2002) and recent research on unilateral divorce laws indicates only a small, transitory impact on divorce rates (Wolfers 2006). Based on the most recent evidence, it appears that the rather drastic changes in the observed divorce rate over time are still awaiting an explanation.

The purpose of this study is to provide some new empirical evidence on the likely causes of the surge in the divorce rate during the 1960s by extending the literature in a number of ways. First and foremost, we argue, similar to Wolfers (2006), that it is impossible to understand the mid-60s surge in the divorce rate without looking significantly beyond this time period. The long-run driving forces of the divorce rate need to be captured before one can reasonably debate the causes of sudden changes, such as those experienced in the mid-1960s. For that purpose, we extend the analysis back to 1929. Second, we bring to bear on the data a number of likely causal factors for the 1960s surge in divorce rates that have been discussed in separate studies but have apparently not been combined in one study. In particular, we consider the legal availability of oral contraceptives, divorce law changes, the likely impact of the Vietnam War, and macroeconomic factors.

Our key result pertains to the long-run relationship between the FLFPR, or our proxy for it, female participation in higher education, and the divorce rate. We show that the uncertainty about the direction of causality is likely the result of discounting the possibility that the FLFPR and divorce rates are jointly determined. More importantly, we demonstrate that the commonly accepted notion that the divorce rate and the FLFPR are positively related stems from the fact that previous studies have been too narrowly focused on the years of interest, that is, the 1960s and 1970s. By 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, 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.

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. However, by relying on the evidence presented in previous research, *inter alia* Smith (1997) and Wolfers (2006), we conclude that the availability of oral contraception is the more likely causal factor for the rise in the U.S. divorce rate than divorce law changes. The Vietnam War is shown to also have had a very significant impact on the U.S. divorce rate.

The remainder of this study is organized as follows. Section 2 provides background information on the key variables used in the analysis and their time-series behavior. 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 the econometric methodology, respectively. Section 5 presents our findings. Section 6 provides a brief summary and some concluding remarks.

2 Institutional Framework and Empirical Regularities

2.1 Female Participation in the Labor-Force and in Higher Education

A number of researchers have analyzed the impact of the rising economic independence of women on the rise in the divorce rate (e.g., Bremmer and Kesselring 2004; Nunley 2008). Economic independence is typically associated with increases in the FLFPR or in the participation of females in higher education.

Using the FLFPR as a proxy for women's rising economic independence may not be ideal because until the late-1960s 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.[1] 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. 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.

[1] The percentage of female participation in higher education rose from the late-1930s to the early WWII years, but it fell substantially following the war's end, when many war veterans began attending college because of incentives created by the GI bill (Figure 2) (Goldin et al. 2006).

The theoretical relationship between female participation in higher education and the divorce rate is not clear. Becker's (1973, 1974) 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, for example market work, the traditional family model predicts a decline in the returns from marriage. 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 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, which could reduce divorce rates.

Previous empirical work on the relationship between the divorce rate and the FLFPR or female participation in higher education centers on the 1960s and 1970s and finds a positive association (e.g., Bremmer and Kesselring 2004; Nunley 2008). However, interpreting this as causation may be problematic because a comparison of Figures 1 and 2 suggests that both the FLFPR and female participation in higher education began to increase well before the dramatic rise in the divorce rate in the mid-1960s and 1970s.[2] Figure 3, which covers the years 1929 to 2006, illustrates why a positive, long-run relationship can easily be found between female participation in higher education and the divorce rate: the least-squares-regression line through the scatter plot of Figure 3 has a distinct, positive slope. However, Figure 3 also indicates two strong, negative relationships between the divorce rate and female participation in higher

[2] Likewise, identifying the effect of the FLFPR on the divorce rate has also proven to be difficult because there is evidence suggesting that the two variables may be simultaneously determined (Johnson and Skinner 1986; Sen 2002; Bremmer and Kesselring 2004).

education, one prior to 1965 (excluding the WWII years) and one from the mid-1970s onward. The negative relationships for these time spans are presented separately in Figures 4 and 5, with each figure containing a least-squares-regression line. Since Figures 4 and 5 contain together the vast majority of the data points of the 1929 to 2006 time period, they are highly suggestive of a negative, long-run relationship between the divorce rate and female participation in higher education or in the labor force.

The time period from the mid-1960s to the mid-1970s appears akin to a transitional period in Figure 3: a period in which the divorce rate permanently shifted upward to a higher level.[3] What appears clear from Figure 3 is that the rise in the divorce rate over this time period cannot be traced to the increase in female participation in higher education or in the labor force. Other factors that affected family life over this period are more likely to be the driving forces behind the sharp increase in the divorce rate. They are discussed next.

2.2 The Transitional Period: the Mid-1960s to the Mid-1970s

Over the period of rising divorce rates, a number of legal, social, and economic changes took effect: general access to oral contraception, divorce law changes, the Vietnam War, and increased variability in standard macroeconomic variables. Although their effects on the incidence of divorce have been studied extensively, we include them in our analysis to help uncover the influence of female participation in higher education.

Goldin and Katz (2000, 2002) find that access to the pill increased women's age at marriage, which can improve marital sorting through a reduction in the opportunity costs of postponing marriage. Increases in marriage-match quality have the potential to reduce divorce rates. In fact,

[3] We note that WWII also shifted the divorce rate to a higher level. This is visible from Figure 1. However, this shift was clearly temporary and, therefore, different from the events from the mid-1960s to mid-1970s.

Goldin and Katz (2002) find a negative effect of access to oral contraceptives on the divorce probabilities of college-educated women. However, there is also credible empirical evidence of a positive effect of access to the pill on divorce rates. Using time-series data from England and Wales, Smith (1997) shows that access to oral contraceptives increased the divorce rate in both regions.

Access to the pill could have a positive effect on the divorce rate for several reasons. First, the advent of the pill allowed women to participate in market work at higher rates (Bailey 2006) and permitted college-educated women to enter professional fields (Goldin and Katz 2000, 2002). In the traditional family model, an increase in market work for wives, with no change in husbands' time allocated to the labor market, leads to a decline in the returns from marriage. Second, the pill reduces 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. Prospective spouses may sort into riskier marriages because the costs of divorce are lower when fertility can be controlled.

Figure 6 provides a scatterplot of the percentage of the population affected by access to oral contraceptives and the divorce rate. It shows 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. 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.

Studies that examine the impact of major wars on the divorce rate use time-specific, indicator variables to capture these effects (South 1985; Anderson and Little 1999).[4] South (1985) finds that the divorce rate increased during the Vietnam War period, but no statistical evidence links 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. Anderson and Little (1999) also establish a statistically significant, positive effect of the WWII years on the divorce rate.

The period of the rising divorce rate also coincided with the adoption of unilateral divorce laws, which provided unrestricted access to divorce for either spouse.[5] The impact of divorce-law changes on divorce rates has been studied extensively (e.g., see Friedberg 1998; Gruber 2004; Wolfers 2006). The majority of studies identify a small, transitory, positive effect of unilateral divorce laws on divorce rates.

Concomitant with the rise in divorce, there was increased variability in macroeconomic conditions. The literature seems to have reached 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 Nunley 2008). In the context of the traditional marriage model, economic growth may create an incentive for spouses who specialize in home production to participate in market work. As a result, the returns from marriage decline, which could increase the divorce rate.

Another macroeconomic indicator that has been shown to have sizable, persistent, and positive effects on the divorce rate is the inflation rate (Nunley 2008). An increase in the

[4] 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).

[5] Proof of marital wrongdoing or mutual consent was required by courts to grant divorces prior to the adoption of unilateral divorce.

inflation rate worsens the terms of trade between spouses, which reduces the returns from marriage. This is because spouses who specialize in home production may be forced to enter the labor market in order to achieve pre-inflation consumption levels, which lowers the returns from marriage in the traditional family model. Regardless of whether one uses the traditional family model or one in which consumption complementarities define the returns from marriage, inflation reduces these returns, because it acts as a tax on the household.[6]

3 Data

This study uses time-series data from 1929 to 2006. Compared to cross-section or panel data studies (e.g., Becker et al. 1977; Weiss and Willis 1997; Charles and Stephens 2004), the much longer time horizon makes it possible to identify the uniqueness of the mid-1960s to mid-1970s period with its surge in the divorce rate. Using this long sample requires some solutions to apparent data problems, such as the unavailability of some variables back to 1929. This 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 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 illustrates that the FLFPR and female participation in higher education display very similar trends from 1948 to 2006.

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

[6] It is possible for wage increases to offset rising prices. However, 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.

collecting data on divorces per 1,000 married couples in 1997. By creating an update of the latter divorce measure from generally available data sources,[7] we can show that the two divorce measures display similar behavior over the complete sample used in this study. In fact, a scatterplot of divorces per 1,000 persons and divorces per 1,000 married couples reveals a clean, linear relationship between the two variables (Figure 7). This suggests that the estimated effects of our explanatory variables would be similar regardless of which divorce measure is used.[8]

Variables for access to oral contraception and unilateral divorce are constructed by dividing cumulated state populations that have adopted 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. We use Bailey's (2006) coding for the oral contraceptive laws, and we use Friedberg's (1998) and Gruber's (2004) codings for unilateral divorce reform. We also check the sensitivity of the effect on the divorce rate of unilateral divorce reform to alternative law codings, but the results are largely robust across different classification schemes.[9]

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. More specifically, the variable that proxies for the "stress" of the Vietnam War is defined as U.S.

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[7] 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 draws on 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.

[8] We also estimate all models with divorces per 1,000 married couples as the dependent variable. The estimates from these models are not materially different from the models that use divorces per 1,000 persons. See Appendix.

[9] Historical accounts of divorce-law reform and reforms allowing access to oral contraceptives suggest that they are plausibly exogenous. See Jacob (1988) for details on divorce-law reform and Goldin and Katz (2002) and Bailey (2006) for a discussion of state-level reforms allowing young, unmarried women access to oral contraceptives.

military personnel deaths due to the Vietnam War as a fraction of U.S. military deployments in East Asia. The data series for the Korean War and WWII are derived in analogy to those of the Vietnam War, except troop deployments for WWII are culled from Matloff (1990).[10]

We also include three standard macroeconomic variables: the inflation rate, economic growth, and changes in the unemployment rate. An interaction term between economic growth and an indicator variable for whether the economy is in a(n) recessionary or expansionary period allows economic growth to have asymmetric effects on the divorce rate in recessionary and expansionary periods. Including squared terms of some explanatory variables allows for nonlinear responses. These variables are required in some models because Ramsey's RESET tests reveal functional form misspecification in their absence.

Table 1 provides variable names, definitions, and sources, while Table 2 presents summary statistics for 1929-2006, 1929-1948, and 1949-2006. Because of the apparent structural break for some of the macroeconomic variables 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 indicate that there is far less volatility in the variables from 1949-2006.

We test each of the variables used in our analysis for the presence of a unit root and stationarity (Table 3). This indicates 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 dummy variables 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*, and *friedberg* are also nominally I(1). However, they are not following random walks but are the

[10] See Appendix for sources and more detail on the construction of the war variables.

result of a stable diffusion process. The variables *inflation*, *ygrowth*, *ygrowth*^{2}, *das***ygrowth*, and Δ*unemp* are I(0).

4 Econometric Methodology

To encompass earlier results provided in the literature, 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. As an additional guard against spurious results, especially the omission of important variables, we check whether the single-equation models contain any stochastic or non-stochastic trend components that are not captured by the 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). As this methodology is not in common use, it is briefly described in the next few paragraphs.

Checking for unobserved stochastic or non-stochastic trends is a way to identify whether a nonstationary dependent variable, in our case the divorce rate, is driven by the presence of a trend in addition to the included regressor variables. If a trend can be identified by the UCM, one may surmise that important variables are omitted from the regression and that an OLS regression that ignores the trend generates potentially meaningless results. To allow for unobserved trend components, we specify the UCM in the form

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

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:

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

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

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.[11] 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 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).

[11] 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).

As we shall see, single-equation regression 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 new variables that we add to the set of standard explanatory variables of the divorce rate eliminate any sign of or need for a stochastic trend. OLS would therefore appear to generate meaningful results in the sense that no important variables are missing.

However, 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 the divorce rate on the left side of the regression and female participation in higher education on the right. We, therefore, investigate the need for a systems estimator, one that allows both variables to be endogenous and nonstationary. The empirical evidence indicates the need for a systems approach to cointegration as suggested by Johansen and Juselius (1990). As this methodology is well established and amply described in textbooks (e.g. Juselius 2006; Lütkepohl 2007), only a few brief comments are needed. In particular, the trace test is used to test for the cointegration rank. A constant is always included in the cointegration space. In addition to the two endogenous variables *divorce* and *fem-ratio*, a number of models also contain weakly exogenous variables in the cointegration space (*pill*, *friedberg*/*gruber*). Some models also contain dummy variables or variables, such as the war variables, that are treated in this manner in the vector error correction model (VECM). These variables never enter the cointegration space. Whenever dummy variables are added to the VECM, the critical values of the cointegration rank test are based on simulated values with 2,500 replications and random walks of length 400.[12] Bartlett's small sample corrections (Johansen 2002) are employed for the cointegration tests whenever they are available. Cointegration rank tests are conducted on VECMs whose appropriate lag length is verified by a lag length reduction test, which starts with five lags for all models and examines whether fewer lags are supported by the data.

5 Estimation Results

5.1 Single-Equation Estimates

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. 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.[13] 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) 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 autocorrelation in Models 1, 2, and 3, some autocorrelation is present in Model 4.

[12] Details are discussed in the CATS in RATS manual (Dennis et al. 2005).

[13] 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.

The UCM estimation results of Table 5 provide additional evidence that Models 2, 3, and 4 capture the DGP: no deterministic or stochastic trend remains; the estimated variances of the level and slope components are zero or very close to zero in the case of Model 3. By contrast, the OLS specification of Model 1 does not fully capture the DGP. This model contains a deterministic trend although not a stochastic trend when estimated as an UCM. One may note that a strong, stochastic trend emerges for Model 1 if it is estimated without the variables *pill* and *Vietnam*. This trend is depicted in Figure 8 along with the original divorce series. The stochastic trend 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. As is apparent from Model 1 of Table 5, the inclusion of the nontraditional variables, *pill* and *Vietnam*, remove all traces of a stochastic trend although there remains a small deterministic trend and some functional form issues. However, these problems are removed in Model 2 of Table 4 by the inclusion of second powers of *fem_ratio*, and *ygrowth*.

The above discussion suggests two conclusions: first, the nonstandard covariates *pill* and *Vietnam* play a key role in capturing the heretofore unexplained upward trend in the divorce rate from the mid-1960s to the mid-1970s; second, the fact that a relatively simple model, such as Model 2, shows neither a statistical specification problem in Table 4 nor any trace of an unobserved stochastic trend in Table 5 indicates, in fairly strong terms, that no essential explanatory variable is missing from the model.[14]

Table 6 provides the implied long-run marginal effects and elasticities of the variables and models presented in Table 4. 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

[14] This would include variables that are used in some panel data studies of divorce rates, such as changes in welfare payments (e.g., Hoffman and Duncan 1995).

two variables have positive effects on the divorce rate. The long-run impact of inflation on the divorce rate is positive, as suggested by Nunley (2008), but it is only marginally significant. Its short-run impact, by contrast, is largely statistically significant.

Model 4 of Table 4 includes a variable for the Korean War and a variable (*das*ygrowth*) that allows economic growth to have asymmetric effects on the divorce rate inside and outside of recessionary periods. The estimation results suggest that economic growth has a large, statistically significant, positive effect on the divorce rate in recessionary periods but none outside of recessions. This result may provide some impetus to reassess the impact of changes in the macroeconomy on divorce rates. Previous work finds a positive relationship between economic growth and the divorce rate (e.g., see Nunley 2008), but does not distinguish between growth during recessionary and expansionary periods.

Next, we extend the sample from 1949 back to 1929. Table 7 shows least-squares estimates for the sample period 1929-2006. Extending the sample in this way necessitates the inclusion of variables for WWII.[15] 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. The structural change appears associated mainly with the coefficients of the macroeconomic variables. For example, if one compares Tables 4 and 7, *inflation* is no longer statistically significant once the sample is extended back to 1929. Also, *ygrowth* has the opposite directional effect in the longer sample.

[15] Model 7 includes an observation specific dummy variable for the year 1947 (*d47*).

We also estimate UCM models for the time period 1929 to 2006. Table 8 reveals that no unobserved stochastic or deterministic trend is present in any of the three models. As for Models 2 to 4 of Table 5, this suggests that no important variables are omitted from the models.

Because the AIC, SBC, and HQC improve as additional covariates enter the base model of Table 7, we select Model 7 as the preferred model specification of Table 7. The estimates of Model 7 are largely consistent with the results of Table 4 for the variables *fem_ratio*, *pill*, and *Vietnam*. The statistically significant effects of the WWII variables are not surprising given the large change in the divorce rate during those years (Figure 1). The variable for the Korea War, by contrast, has no statistically significant effect in Model 7. This is consistent with previous work (South 1985; Anderson and Little 1999) and Figure 1. The long-run marginal effects and elasticities implied by the estimates of Table 7 are presented in Table 9. The values for the variables *fem_ratio*, *pill*, and *Vietnam* are similar in terms of size and statistical significance to those of Table 6.

The overall conclusion from the single-equation models is that increased access to the pill has a 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 number of divorces per capita by 2.12 in the long run. For the longer sample from 1929 to 2006, the equivalent decrease is estimated to be 1.73 divorces per capita. 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 new 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. 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 is statistically significant only during recessions and it has a negative impact on the divorce rate during those times.

5.2 Systems Estimation

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. Few researchers account for the endogeneity of female participation in higher education or in the labor force. Bremmer and Kesselring (2004) is an exception. They treat the divorce rate and the FLFPR as jointly endogenous variables in their empirical work, but do not identify a statistically significant long-run relationship between the divorce rate and the FLFPR.

Table 10 presents the cointegrating vectors and the underlying tests for cointegration based on the Johansen/Juselius vector error correction (VECM) approach. It is noteworthy that the cointegration rank tests suggest exactly one cointegrating vector regardless of variations in the model 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 include the exogenous variable, *pill*, in the cointegration space of Models 1 to 4 and one of two alternative measures for the diffusion of unilateral divorce laws in Models 5 and 6.

Models 2 to 6 of Table 10 add several unrestricted exogenous variables to the VECM, in particular the war variables *WWII*, *Korea*, *Vietnam*, and two observation-specific dummy variables, *d41* and *d47*. These variables do not have a large 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*.[16]

The coefficients of the cointegrating vectors (CIVs) of Models 1 through 4 of Table 10 largely match the long-run marginal effects estimated for the single-equation models of Tables 6 and 9. For example, the long-run marginal effect of *fem_ratio* is -21.20 for Model 2 of Table 6 and -16.01 for Model 6 of Table 9, while the long-run CIV estimate for *fem_ratio* is -22.71 for Model 4 of Table 10.[17] 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 10.

[16]Adding 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.

[17] The cointegration results as reported in Table 10 are not materially different if the sample is limited to the time period 1949 to 2006. This applies both to the cointegration tests and the estimated elements of the

Models 5 and 6 consider the long-run relationship between the divorce rate and the diffusion of unilateral divorce laws. Signs and statistical significance and fit are very similar between the two models and between Models 5 and 6 and Model 4. In fact, based on statistical grounds, there is little reason to prefer one over the other two models. The similarity of the variables for oral contraception and unilateral divorce law diffusion also do not allow selecting a preferred variable by including both variables simultaneously in a model. It is, therefore, necessary to rely on other work, such as the findings of Smith (1997) and Wolfers (2006), to decide which of the two, oral conceptives or divorce law changes, had more of an impact on the divorce rate during the 1960s and 1970s.

Although not reported in Table 10, 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* is weakly exogenous. As a consequence, we expect the OLS regressions to underestimate the impact of the variables *fem_ratio* and *pill* on the divorce rate. This is borne out by the estimates: the cointegrating vectors of Table 10 all imply a larger long-run coefficient values for the variables *fem_ratio* and *pill* than the marginal effects presented in Table 9.

5.3 Plausibility of Estimation Results

To assess the plausibility of the estimation results presented so far, it is helpful to first revisit Figures 3 to 5. The figures relate to the relationship over time of the divorce rate and the participation of females in higher education, but they are materially the same as those for the

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cointegrating vector and their statistical significance. For example, the CIV corresponding to Model 1 of Table 12 for the time period 1949 to 2006 is given as (1.000, 23.316, -6.671, -11.426) and the p-values for the cointegration tests are 0.000 and 0.207, respectively.

divorce rate and the more common variable FLFPR. Hence, the discussion has more general implications.

The linear regression line of Figure 3 has a positive slope. This is not consistent with any of the long-run marginal effects estimated for the variable *fem_ratio* in Tables 6, 9, or 10. As mentioned before, the positive slope hides two distinct negative relationships between the divorce rate and the variable *fem_ratio*, one at the lower left end of Figure 3 and the other at the upper right end of Figure 3. The two negative relationships are shown separately in Figures 4 and 5. The regression coefficients on *fem_ratio* in Figures 4 and 5 (-19.1 and -11.2, respectively) are very close to and, hence, fully consistent with the long-run marginal effects of the single-equation models reported in Table 6 (in particular preferred Model 2), Table 9 (in particular preferred Model 7), and the cointegration relationships of Table 10 (in particular preferred Model 4).

The plausibility of the estimation results for the variable *pill* can best be assessed by revisiting Figure 6. The regression line superimposed on the scatterplot of Figure 6 reveals a positive regression coefficient for variable *pill* (2.27). This coefficient is fully consistent in sign and in order of magnitude with the long-run marginal effects estimates of variable *pill* reported in Tables 6, 9, and 10. The estimates presented on the highly significant and positive effect of the Vietnam War variable on the divorce also appear to be well reflected in Figure 6, if one allows for a one-year lag to allow the stress of war to affect the divorce rate. Allowing for the one-year lag, the values of the divorce rate in Figure 6 tend to move above the regression line when the Vietnam War variable reaches its highest values (1966 to 1970). One may note, that the higher than predicted values of the divorce rate for 1970 and 1971 may also be related, as suggested by Wolfers (2006) to the positive impact of the change in the divorce laws, which affected a large number of states in 1970s. The results of Table 10 would certainly be consistent with this explanation.

We find that diffusion of the pill and unilateral divorce laws have qualitatively and 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. From the above discussion, it would appear reasonable to assume that access to oral contraception had the major impact on the divorce rate, with some supporting effect of the divorce law changes starting in 1970. Such an interpretation would be largely consistent with recent results presented in the literature. For example, Wolfers (2006), using state-level panel data for the U.S., concludes that the adoption of unilateral divorce laws led to a small, transitory rise in U.S. divorce rates.[18] The careful econometric evidence collected by Smith (1997) of the impact of the pill on divorce rates in England and Wales also supports the conclusion 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. Similar to Wolfers (2006) for the U.S., Smith (1997) concludes that reforms allowing easier access to divorce led only to a temporary rise in divorce rates, with no evidence of a long-run relationship. Finally, visual inspection of time plots of the divorce rates of most European countries (Gonzalez and Viitanen 2006, Figure 2) indicate a distinct rise in divorce rates in almost all countries prior to the implementation of either no-fault or unilateral divorce laws.[19]

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

[19] The manuscript can be accessed at ftp://repec.iza.org/RePEc/Discussionpaper/dp2023.pdf.

6 Summary and Conclusions

This study examines the evolution of the U.S. divorce rate from 1929 to 2006, with the primary aim of explaining the puzzling 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 access to oral contraception contributed to the rise in the divorce rate and how this compares in its impact on divorces of changes in divorce laws in the 1970s. Second, we extend the analysis of the determinants of divorce rates back to 1929 to allow for more variation in the sample observations. Third and finally, we construct objective measures for the effects of WWII, the Korean War, and the Vietnam War on the divorce rate; this contrasts with all previous research which has used observation-specific indicator variables to capture the effects of wars on divorce rates.

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. By limiting the analysis to a sample that is overweighted by observations from the 1960s and 1970s, it is easy to miss that there is a very distinct negative relationship between the divorce rate and female participation in the labor force and higher education: this holds in general except for the ten-year time horizon from the mid-1960s to the mid-1970s.

The sharp rise in the divorce rate over that time period is the outcome of a number of transitional influences that permanently shifted the U.S. divorce rate to a higher level. The influences that are responsible for this marked shift in the level of the divorce rate are, in our assessment, mainly associated with increased access to oral contraception. The ramifications of the Vietnam War and divorce-law reforms also played a role, although it appears that their impact may have been more of a temporary nature.

References

Anderson, William L. and Derek W. Little. 1999. "All's Fair: War and Other Causes of Divorce From a Beckerian Perspective." *American Journal of Economics and Sociology*, 58(4): 901-922.

Bailey, Martha J. 2006. "More Power to the Pill: The Impact of Contraceptive Freedom on Women's Labor Supply." *Quarterly Journal of Economics*, 121(1): 289-320.

Becker, Gary S. 1973. "A Theory of Marriage: Part I." *Journal of Political Economy*, 81(4): 813-846.

_____. 1974. "A Theory of Marriage: Part II." *Journal of Political Economy*, 82(2): S11-S26.

_____, Elisabeth Landes, and Robert Michael. 1977. "An Economic Analysis of Marital Instability." *Journal of Political Economy*, 85(6): 1141-1188.

Bremmer, Dale and Randy Kesselring. 2004. "Divorce and Female Labor-Force Participation: Evidence from Time-Series Data and Cointegration." *Atlantic Economic Journal*, 32(3): 175-190.

Charles, Kerwin and Melvin Stephens. 2004. "Job Displacement, Disability, and Divorce." *Journal of Labor Economics*, 22(2): 489-522.

Christiano, Lawrence, Martin Eichenbaum, and Charles Evans. 2005. "Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy." *Journal of Political Economy*, 113(1): 1-45.

Costa, Dora. 2000. "From Mill Town to Board Room: The Rise of Women's Paid Labor." *Journal of Economic Perspectives*, 14(4): 101-122.

Dennis, Jonathan, Henrick Hansen, Søren Johansen, and Katarina Juselius. 2005. *CATS in RATS, Version 2*. Estima: Evanston, IL.

Durbin, James, and Siem J. Koopman. 2001. *Time Series Analysis by State Space Methods.* Oxford University Press: New York.

Friedberg, Leora. 1998. "Did Unilateral Raise Divorce Rates? Evidence from Panel Data." *American Economic Review*, 88(3): 608-627.

Goldin, Claudia. 2006. "The Quite Revolution That Transformed Women's Employment, Education, and Family." *American Economic Review*, 96(2): 1-21.

_____ and Lawrence F. Katz. 2000. "Career and Marriage in the Age of the Pill." *American Economic Review*: *Papers and Proceedings of the American Economic Association*, 90(2): 461-465.

_____ and _____. 2002. "The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions." *Journal of Political Economy*, 110(4): 730-770.

_____, _____, and Ilyana Kuziemko. 2006. "The Homecoming of American College Women: The Reversal of the College Gender Gap." *Journal of Economic Perspectives*, 20(4): 133-156.

Goldstein, Joshua. 1999. "The Leveling of Divorce Rates in the United States." *Demography*, 36(3): 409-414.

Gonzalez, Libertad and Tara K. Viitanen. 2006. "The Effect of Divorce Laws on Divorce Rates in Europe." *IZA Discussion Paper No. 2023*.

Gruber, Jonathan. 2004. "Is Making Divorce Easier Bad for Children? The Long-Run Implications of Unilateral Divorce." *Journal of Labor Economics*, 22(4): 799-833.

Hoffman, Saul D., and Greg J. Duncan. 1995. "The Effect of Incomes, Wages, and AFDC Benefits on Marital Disruption." *Journal of Human Resources,* 30(1): 19-41.

Hamilton, James D. 2008. "Macroeconomics and ARCH." *Working Paper*, University of California, San Diego.

Harvey, Andrew. 1989. *Forecasting Structural Time Series Models and the Kalman Filter*. Cambridge University Press: Cambridge.

_____. 1997. "Trends, Cycles, and Autoregressions." *Economic Journal*, 107(440): 192-201.

Johansen, Søren. 2002. "A Small Sample Correction of the Test for Cointegrating Rank in the Vector Autoregressive Model." Econometrica, 70(5): 1929-1961.

_____ and Katarina Juselius. 1990. "Maximum Likelihood Estimation and Inference on Cointegration with Application to the Demand for Money." *Oxford Bulletin of Economics and Statistics*, 52(2): 169-210.

Johnson, William and Jonathan Skinner. 1986. "Labor Supply and Marital Separation." *American Economic Review*, 76(3): 455-469.

Juselius, Katarina. 2006. *The Cointegrated VAR Model: Methodology and Applications*. Oxford University Press: Oxford and New York.

Koopman, Siem, Andrew Harvey, Jurgen Doornik. and Neil Shephard. 2000. *Stamp*: *Structural Time Series Analyser*, *Modeller*, *and Predictor.* Timberlake Consultants Press: London.

Lütkepohl, Helmut. 2007. *New Introduction to Multiple Time Series Analysis*. Corr. 2nd printing, Springer: Berlin, Heidelberg, New York.

Matloff, Maurice. 1990. *Strategic Planning for Coalition Warfare, 1943-1944*. Center of Military History, United States Army, Washington, D.C.; available online at http://www.history.army.mil/books/wwii/sp1943-44/index.htm.

Michael, Robert. 1978. "The Rise in Divorce Rates, 1960-1974: Age-Specific Components." *Demography*, 15(2): 177-182.

Nunley, John. 2008. "Inflation and Other Aggregate Determinants of the Trend in U.S. Divorce Rates since the 1960s." *Applied Economics*, forthcoming.

Ruggles, Steven. 1997. "The Rise of Divorce and Separation in the United States, 1880-1990." *Demography*, 34(4): 455-466.

Sen, Bisakha. 2000. "How Important is Anticipation of Divorce in Married Women's Labor Supply Decisions? An Intercohort Comparison Using NLS Data." *Economic Letters*, 67(2): 209-216.

Smith, Ian. 1997. "Explaining the Growth of Divorce in Great Britain." *Scottish Journal of Political Economy*, 44(5): 519-544.

South, Scott. 1985. "Economic Conditions and the Divorce Rate: A Time-Series Analysis of Postwar United States." *Journal of Marriage and the Family*, 47(1): 31-41.

Stevenson, Betsey and Justin Wolfers. 2007. "Marriage and Divorce: Changes and their Driving Forces." *Journal of Economic Perspectives*, 21(2): 27-52.

Stock, James and Mark W. Watson. 2003. *Introduction to Econometrics*. Pearson Addison Wesley: Boston, San Francisco, New York, London.

Weiss, Yoram and Robert Willis. 1997. "Match Quality, New Information, and Marital Dissolution." *Journal of Labor Economics*, 15(1): S293-S329.

Wolfers, Justin. 2006. "Did Unilateral Divorce Laws Raise Divorce Rates? A Reconciliation and New Results." *American Economic Review*, 96(5): 1802-1820.

Table 1: Variable Names, Definitions, and Sources

Name	Definition	Source
divorce	Number of divorces per 1,000 persons	*Historical Statistics of the United States*: *Millennium Edition* and various U.S. Statistical Abstracts.
WWII	U.S. military personnel deaths due to WWII as a fraction of U.S. military deployments for WWII.	See Appendix
Korea	U.S. military personnel deaths as a fraction of U.S. military deployments in East Asia.	See Appendix
Vietnam	U.S. military personnel deaths due to the Vietnam War as a fraction of U.S. military deployments in East Asia.	See Appendix
pill	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 the *Historical Statistics of the United States*: *Millennium Edition* and U.S. Statistical Abstracts.
friedberg/ gruber	Diffusion function, measured as the percentage of the U.S. population affected by unilateral divorce law reform.	Law codings are from (1998) and Gruber (2004); population data by state and for the U.S. are from the *Historical Statistics of the United States*: *Millennium Edition* and U.S. Statistical Abstracts.
inflation	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
ygrowth	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
*das*ygrowth*	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
fem_ratio	Percentage of women enrolled in higher education relative to the total population enrolled in higher education	*Historical Statistics of the United States*: *Millennium Edition* and various U.S. Statistical Abstracts.
Δ*unemp*	Annual change in the unemployment rate, measured as the percentage of the population who is unemployed but actively pursuing employment.	1929 to 1941: U.S. Bureau of the Census, Historical Statistics of the United States, Colonial Times to 1957 (Washington, D.C., 1960), p.70; 1942 to 1947: BLS statistics at www.bls.gov/cps/cpsaat1.pdf; for 1948 to 2006: FRED St. Louis, monthly data averaged.

Table 2: Summary Statistics

Variable	Mean	Med.	Min.	Max.	St. Dev.	C.V.	Skew.	Kurt.
1929-2006								
divorce	3.36	3.45	1.30	5.30	1.22	0.36	0.04	-1.44
fem_ratio	0.45	0.43	0.29	0.58	0.08	0.18	-0.04	-1.19
pill	0.46	0.13	0.00	1.00	0.48	1.05	0.18	-1.93
inflation	3.30	2.95	-10.30	14.65	4.06	1.23	-0.13	2.34
ygrowth	3.47	3.60	-13.05	18.43	5.08	1.46	-0.13	2.58
*das*ygrowth*	-0.66	0.00	-13.05	0.00	2.26	3.44	-4.18	17.11
Δ *unemp*	0.02	-0.25	-4.96	7.71	2.08	109.1	1.16	3.64
WWII	0.14	0.00	0.00	2.80	0.50	3.62	3.66	12.81
Korea	0.17	0.00	0.00	9.48	1.10	6.49	7.96	64.12
Vietnam	0.12	0.00	0.00	2.16	0.40	3.18	3.59	12.46
1929-1948								
divorce	2.22	1.90	1.30	4.30	0.81	0.36	1.09	0.40
fem_ratio	0.41	0.42	0.29	0.50	0.05	0.13	-1.05	0.84
inflation	1.90	1.96	-10.30	14.65	6.17	3.24	-0.05	-0.10
ygrowth	3.66	4.85	-13.05	18.43	9.37	2.56	-0.13	-0.99
*das*ygrowth*	-2.29	0.00	-13.05	0.00	4.08	1.78	-1.64	1.23
Δ *unemp*	0.03	-0.70	-4.96	7.71	3.84	119.5	0.73	-0.56
WWII	0.54	0.00	0.00	2.80	0.89	1.65	1.23	0.11
1949-2006								
divorce	3.75	4.00	2.10	5.30	1.09	0.29	-0.22	-1.51
fem_ratio	0.47	0.49	0.30	0.58	0.09	0.19	-0.26	-1.46
pill	0.62	1.00	0.00	1.00	0.47	0.75	-0.46	-1.73
inflation	3.78	3.02	-0.95	13.58	2.94	0.78	1.36	1.77
ygrowth	3.41	3.51	-1.87	8.74	2.37	0.69	-0.15	-0.29
*das*ygrowth*	-0.09	0.00	-1.87	0.00	0.30	3.25	-4.25	19.73
Δ *unemp*	0.01	-0.25	-2.09	2.83	1.06	71.35	0.92	0.73
Korea	0.23	0.00	0.00	9.48	1.28	5.59	6.81	46.48
Vietnam	0.17	0.00	0.00	2.16	0.45	2.70	2.98	8.13

Table 3: Unit Root and Stationarity Tests
(1929-2007 or maximum available)

	ADF – H0: I (1)			KPSS – H0: I(0)	
Variable	Constant	Constant with Trend	Lag Order	Trend	No Trend
Continuous Variables:					
divorce	0.445	0.852	1	1.441	2.797
fem_ratio	0.943	0.576	4	0.283	1.245
fem_ratio^2	0.941	0.542	3	1.586	1.586
pill	0.790	0.563	1	1.785	2.500
friedberg	0.860	0.573	4	0.204	1.427
gruber	0.871	0.592	4	0.261	1.455
mechoulan	0.869	0.622	5	0.193	1.226
inflation	0.001	0.018	2	0.369	0.353
ygrowth	0.000	0.000	3	0.063	0.063
ygrowth^2	0.088	0.000	3/2	0.843	0.843
*das*ygrowth*	0.000	0.000	0	0.624	1.141
Δ *unemp*	0.000	0.000	3	0.064	0.064
Quasi Dummy Variables:					
WWII	0.005	0.010	2	0.329	0.401
WWII^2	0.006	0.015	2	0.301	0.358
Korea	0.000	0.000	0	0.167	0.240
Vietnam	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 follow 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	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
constant	0.8753	0.000	-0.7652	0.236	-0.7529	0.237	-0.8318	0.233
divorce (-1)	0.8761	0.000	0.8705	0.000	0.8732	0.000	0.8881	0.000
fem_ratio	-1.9070	0.000	5.8625	0.045	5.3304	0.066	5.8634	0.072
fem_ratio^2			-8.6085	0.006	-7.8314	0.011	-8.3935	0.016
pill	0.5912	0.000	0.5738	0.000	0.5453	0.000	0.5238	0.000
inflation	0.0143	0.019	0.0101	0.073	0.0125	0.032	0.0091	0.122
ygrowth	0.0112	0.042	0.0343	0.011	0.0562	0.005	0.0145	0.566
$ygrowth^2$			-0.0040	0.035	-0.0049	0.014	0.0002	0.938
*das*ygrowth*							0.1624	0.018
d_unemp					0.0393	0.130	0.0443	0.086
Vietnam (-1)	0.1850	0.000	0.1319	0.000	0.1284	0.000	0.1247	0.000
Korea							-0.0120	0.334
Unadjusted R^2	0.9935		0.9952		0.9955		0.9960	
Adjusted R^2	0.9928		0.9945		0.9946		0.9950	
Log-likelihood	59.408		68.305		69.704		73.351	
AIC	-104.815		-118.610		-119.408		-122.702	
SBC	-90.392		-100.066		-98.804		-97.977	
HQC	-99.197		-111.386		-111.382		-113.071	
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)		0.831		0.277		0.241		0.050
Autocorrelation LM(2)		0.798		0.378		0.213		0.033
Autocorrelation LM(3)		0.849		0.344		0.313		0.077
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
Quandt LR (max)		<0.01		>0.10		>0.10		>0.10
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 lenths 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 2003); 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	0.0068	1.0000	0.0065	1.0000	0.0064	1.0000	0.0059	1.0000
σ_{η}^2	0.0000	0.0000	0.0000	0.0000	0.0001	0.0116	0.0000	0.0005
σ_{ξ}^2	0.0000	0.0007	0.0000	0.0004	0.0000	0.0000	0.0000	0.0000
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
μ	-0.745	0.323	-0,846	0.295	-0.932	0.243	-1.256	0.133
β	-0.017	0. 028	-0.001	0.865	-0.003	0.705	-0.008	0.351

Notes: Only parameters that relate to stochastic or deterministic trends are provided, not the coefficients of the fixed regressors. The first three lines report the estimated variances for the level and slope components when they are allowed to be stochastic. The Q Ratio is the ratio of the estimated variances relative to the variance of the irregular component. Zero variances and Q values imply the lack of a stochastic trend. The last two lines present the coefficient estimates for the level and slope parameters when they are restricted to be fixed (non-stochastic); μ equals the regression constant in this case and β the coefficient of a deterministic trend. The p-values for the slope coefficients in Models 2, 3, and 4 suggest the absence of a deterministic trend. The results imply that these models fully capture the trend in the divorce rate.

Table 6: Implied Long-Run Marginal Effects and Elasticities For Table 4

	Model 1		Model 2		Model 3		Model 4	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
Long-Run Implied Marginal Effects:								
fem_ratio	-15.390	0.004	-21.201	0.000	-19.725	0.001	-22.600	0.002
pill	4.7714	0.000	4.4299	0.000	4.3010	0.000	4.6794	0.000
inflation	0.1151	0.028	0.0782	0.080	0.0987	0.044	0.0809	0.133
ygrowth (if positive)	0.0905	0.090	0.2335	0.044	0.4044	0.035	0.1319	0.528
ygrowth (if negative)	0.0905	0.090	0.2335	0.053	0.4044	0.035	1.5825	0.036
Vietnam	1.4933	0.006	1.0185	0.007	1.0124	0.008	1.1140	0.013
Δ *unemp*					0.3097	0.168	0.3953	0.141
Korea							-0.1069	0.378
Long-Run Implied Elasticities:								
fem_ratio	-0.2365		-0.3442		-0.3135		-0.3171	
pill	0.0976		0.0947		0.0900		0.0865	
inflation	0.1435		0.0102		0.0126		0.0091	
ygrowth (if positive)	0.0102		0.0275		0.0466		0.0135	
ygrowth (if negative)	0.0102		0.0275		0.0466		0.1611	
Vietnam	0.0083		0.0059		0.0057		0.0056	
Δ *unemp*					0.0002		0.0002	
Korea							-0.0007	

Notes: The long-run marginal effects are calculated using the `nlcom` command in STATA.

Table 7: Least Squares Estimates, 1929-2006, HAC Covariance Matrix

	Model 5		Model 6		Model 7	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
constant	-2.1962	0.000	-2.0716	0.000	-1.7562	0.000
divorce (-1)	0.7953	0.000	0.7778	0.000	0.8259	0.000
fem_ratio	14.0684	0.000	12.8679	0.000	10.8408	0.000
fem_ratio^2	-18.0201	0.000	-16.4577	0.000	-13.8258	0.000
pill	0.8432	0.001	0.8460	0.001	0.6390	0.000
inflation	-0.0038	0.626	0.0040	0.648	0.0093	0.199
ygrowth	-0.0169	0.141	0.0275	0.086	0.0177	0.298
ygrowth^2	-0.0013	0.002	-0.0037	0.000	-0.0031	0.001
*das*ygrowth*			-0.0770	0.021	-0.0668	0.021
Δ *unemp*	-0.0517	0.019	-0.0298	0.039	-0.0364	0.035
WWII	0.9095	0.000	1.0661	0.000	0.8877	0.000
WWII^2	-0.2539	0.000	-0.3152	0.000	-0.2568	0.000
Vietnam (-1)	0.1368	0.000	0.1394	0.000	0.1289	0.000
Korea	0.0232	0.002	0.0171	0.014	0.0112	0.162
d47					-0.5498	0.000
Unadjusted R^2	0.9911		0.9922		0.9937	
Adjusted R^2	0.9894		0.9906		0.9923	
Log-likelihood	57.994		62.935		71.461	
AIC	-89.988		-97.871		-112.921	
SBC	-59.518		-65.057		-77.764	
HQC	-77.800		-84.746		-98.859	
RESET		0.535		0.437		0.177
Homoskedasticity		0.000		0.000		0.004
Normality		0.009		0.162		0.025
Autocorrelation LM(1)		0.048		0.095		0.018
Autocorrelation LM(2)		0.076		0.112		0.065
Autocorrelation LM(3)		0.030		0.018		0.017
ARCH (1)		0.001		0.002		0.233
ARCH (2)		0.004		0.004		0.383
ARCH (3)		0.012		0.009		0.599
Quandt LR (max)		<0.01		<0.01		<0.01
Harvey-Collier (CUSUM)		0.663		0.315		0.481

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 2003); 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}	0.0152	1.0000	0.0123	1.0000	0.0115	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	0.0016	0.0000	0.0000
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
μ	-2.472	0.001	-2.409	0.001	-1.902	0.003
β	0.002	0.400	0.002	0.267	0.001	0.624

Notes: Only parameters that relate to stochastic or deterministic trends are provided, not the coefficients of the fixed regressors. The first three lines report the estimated variances for the level and slope components when they are allowed to be stochastic. The Q Ratio is the ratio of the estimated variances relative to the variance of the irregular component. Zero variances and Q values imply the lack of a stochastic trend. The last two lines present the coefficient estimates for the level and slope parameters when they are restricted to be fixed (non-stochastic); μ equals the regression constant in this case and β the coefficient of a deterministic trend. The p-values for the slope coefficients suggest the absence of a deterministic trend. The results imply that the models fully capture the trend in the divorce rate.

Table 9: Implied Long-Run Marginal Effects and Elasticities For Table 7

	Model 5		Model 6		Model 7	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
Long-Run Implied Marginal Effects:						
fem_ratio	-19.303	0.000	-16.152	0.000	-17.146	0.000
pill	4.1189	0.000	3.8064	0.000	3.6702	0.000
inflation	-0.0184	0.577	0.0181	0.618	0.0531	0.203
ygrowth (if negative)	-0.0885	0.226	-0.2396	0.024	-0.2998	0.010
ygrowth (if positive)	-0.0885	0.226	0.1071	0.167	0.0892	0.380
Vietnam	0.6683	0.002	0.6270	0.003	0.7405	0.005
Korea	0.1134	0.014	0.0771	0.112	0.0642	0.385
WWII	3.2025	0.000	3.3786	0.000	3.6241	0.000
Δ *unemp*	-0.2524	0.070	-0.1343	0.173	-0.2090	0.059
Long-Run Implied Elasticities:						
fem_ratio	-2.5852		-2.1632		-2.2963	
pill	0.5639		0.5211		0.5025	
inflation	-0.0181		0.0178		0.0522	
ygrowth (if negative)	-0.0914		-0.2474		-0.3096	
ygrowth (if positive)	-0.0914		0.1106		0.0921	
Vietnam	0.0239		0.0224		0.0264	
Korea	0.0057		0.0039		0.0032	
WWII	0.1334		0.1408		0.1510	
Δ *unemp*	-0.0015		-0.0008		-0.0012	

Notes: The long-run marginal effects are calculated using the `nlcom` command in STATA.

Table 10: Cointegration and VECM Results, 1929 to 2006

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Implicit CIV Normalized on Divorce						
constant	-11.450	-11.333	-11.263	-10.685	-14.627	-14.172
	(-10.01)	(-9.97)	(-9.49)	(-13.22)	(-9.96)	(-11.03)
fem_ratio	24.839	24.908	24.0230	22.710	33.865	33.167
	(8.232)	(8.02)	(7.64)	(10.68)	(8.68)	(9.49))
pill	-7.156	-6.673	-6.7230	-6.589		
	(-13.21)	(-12.50)	(-12.14)	(-17.58)		
gruber					-15.847	
					(-12.93)	
friedberg						-15.857
						(-14.11)
Unrestricted		*WWII*	*WWII*	*WWII*	*WWII*	*WWII*
Exogenous Variables		*Vietnam*	*WWII*^{2}	*WWII*^{2}	*WWII*^{2}	*WWII*^{2}
		Korea	*Vietnam*	*Vietnam*	*Vietnam*	*Vietnam*
			Korea	*Korea*	*Korea*	*Korea*
				d41	*d41*	*d41*
				d46	*d46*	*d46*
Cointegration Trace Test						
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
Statistics and Specification Tests of VECM						
Restricted						
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—Time Plots of Divorces per 1,000 Persons and per 1,000 Married Couples

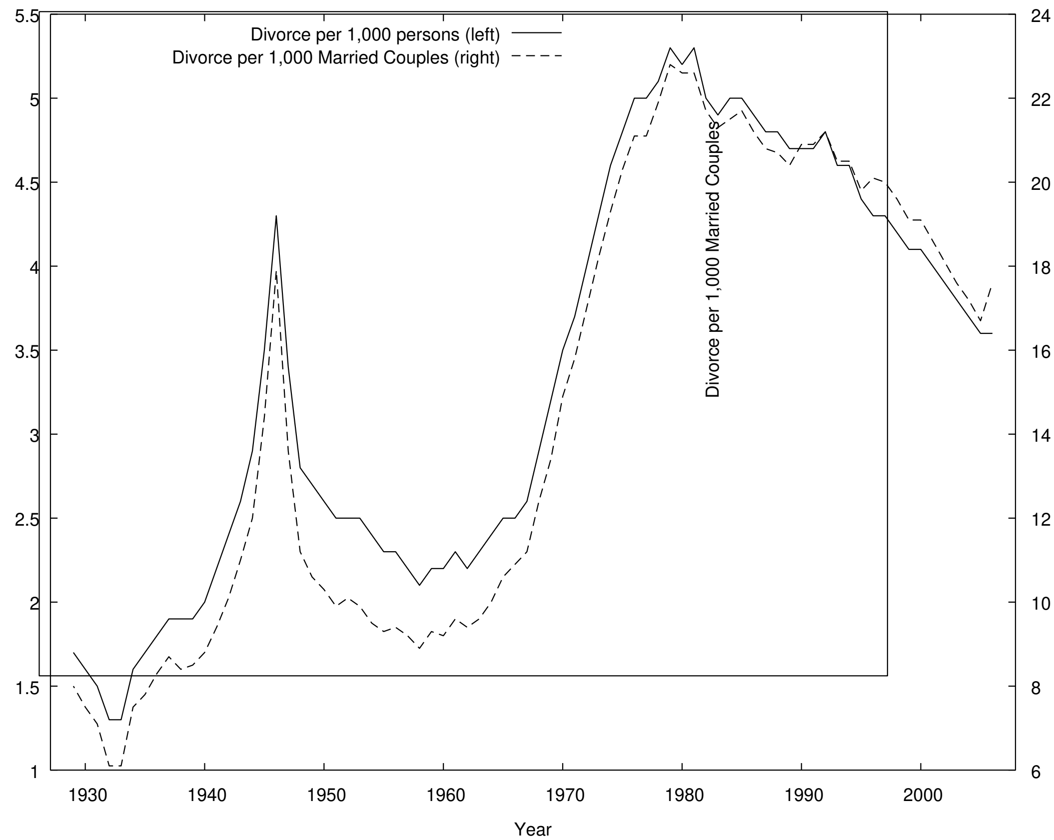


Figure 2—Female Participation in Higher Education and the Female Labor-Force Participation Rate (FLFPR)

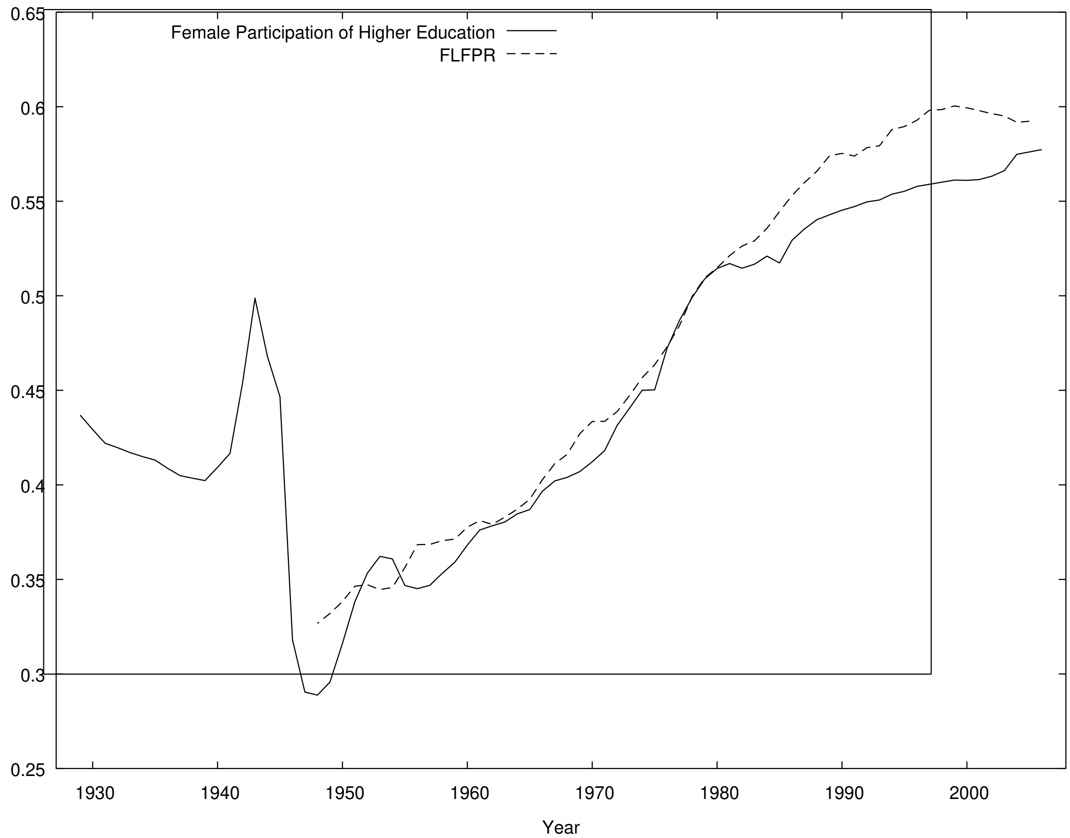


Figure 3—Scatterplot of the Divorce Rate and Female Participation in Higher Education (1929-2006, Least Squares Fit)

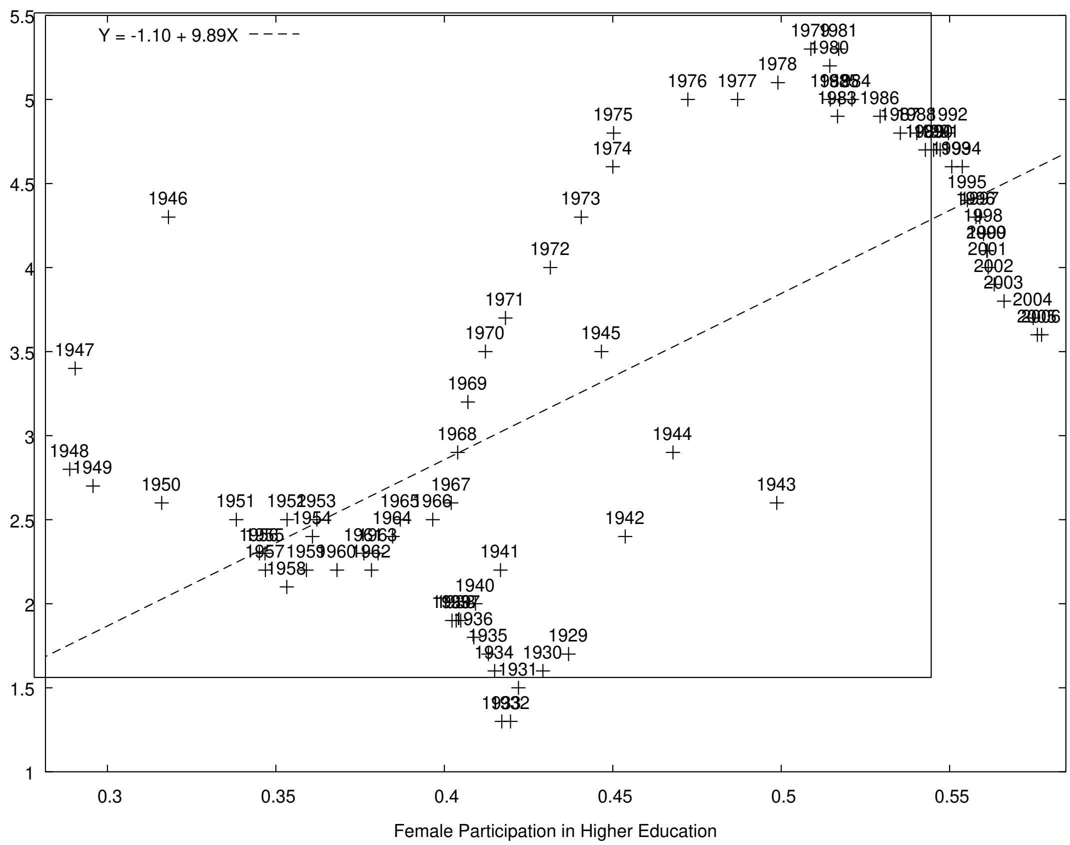


Figure 4—Scatterplot of the Divorce Rate and Female Participation in Higher Education (1977-2006, Least Squares Fit)

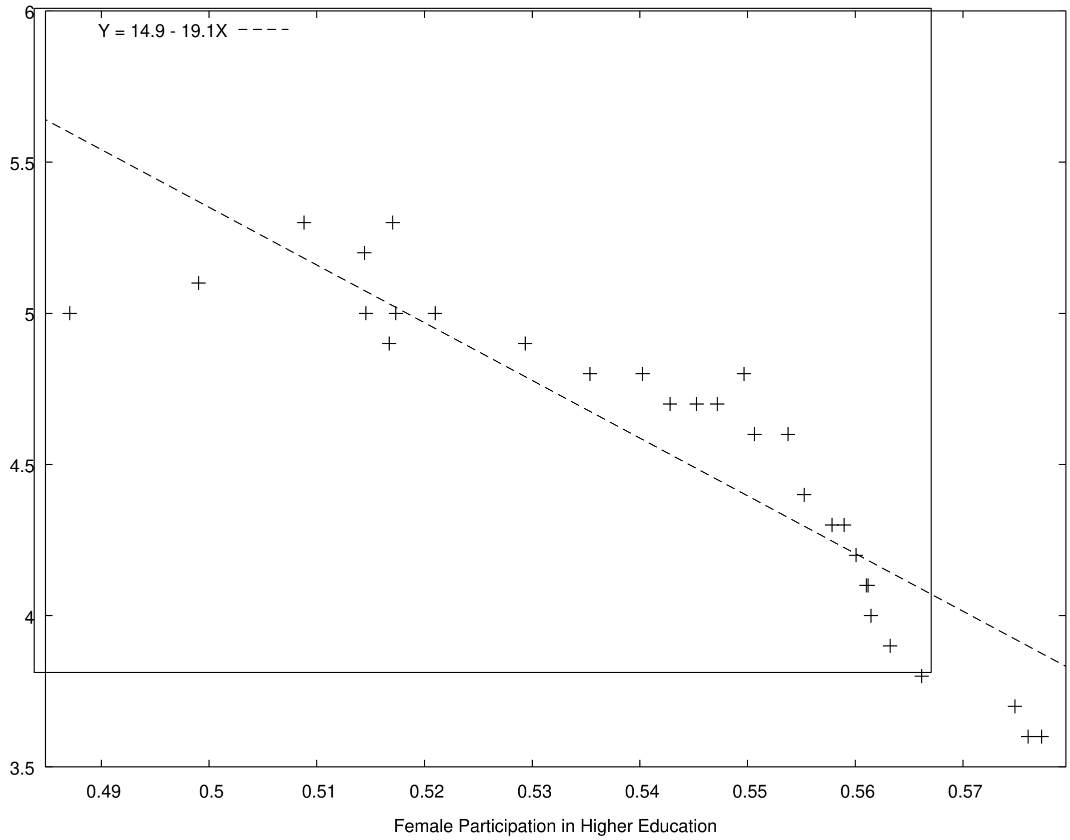


Figure 5—Scatterplot of the Divorce Rate and Female Participation in Higher Education (1929-1965, except 1941-1945, Least Squares Fit)

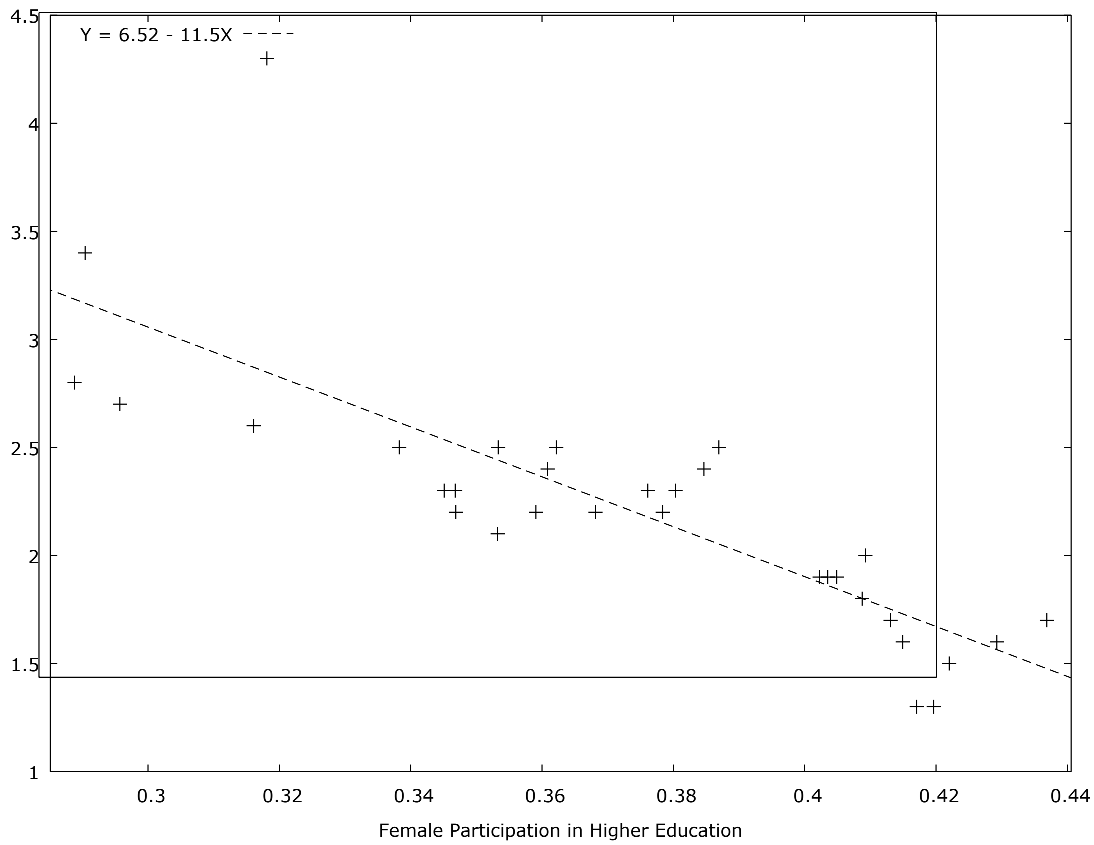


Figure 6—Scatterplot of the Divorce Rate and the Percentage of the Population with Access to Oral Contraceptives (*pill*) (1929-2006, Least Squares Fit)

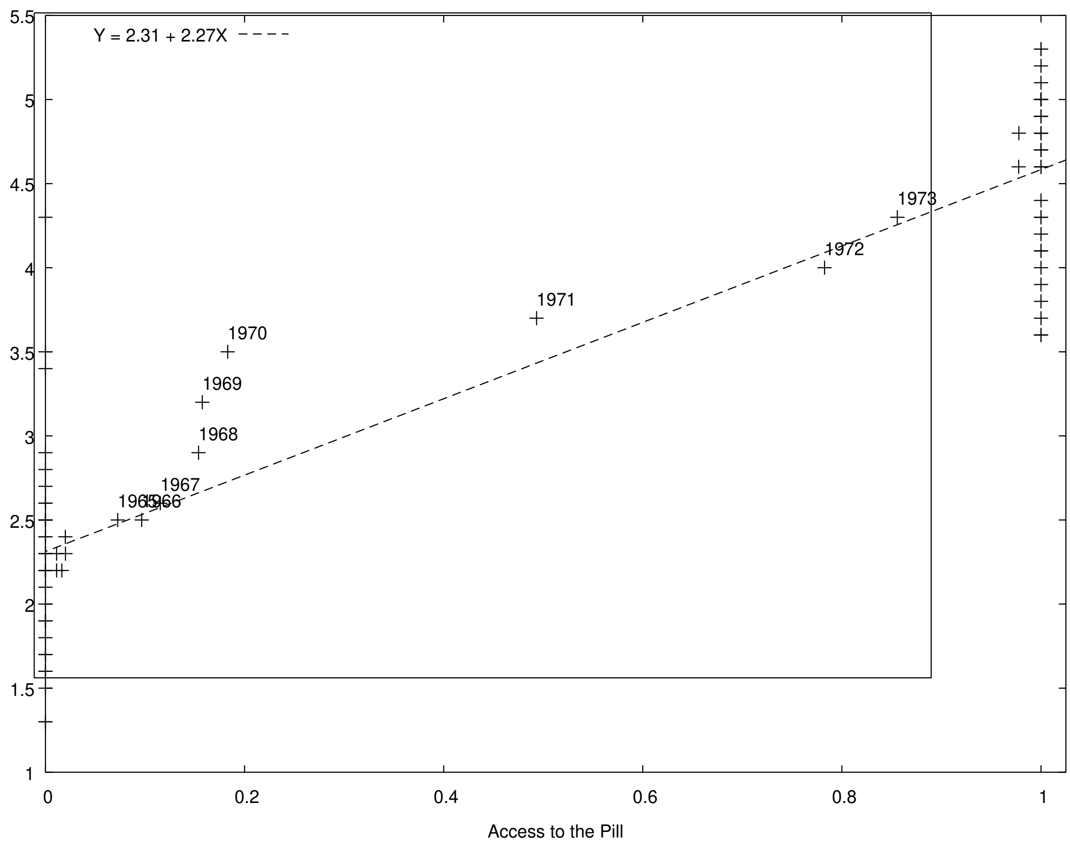


Figure 7— Scatterplot of Divorces per 1,000 Persons and Divorces per 1,000 Married Couples Rate (1929-2006, Least Squares Fit)

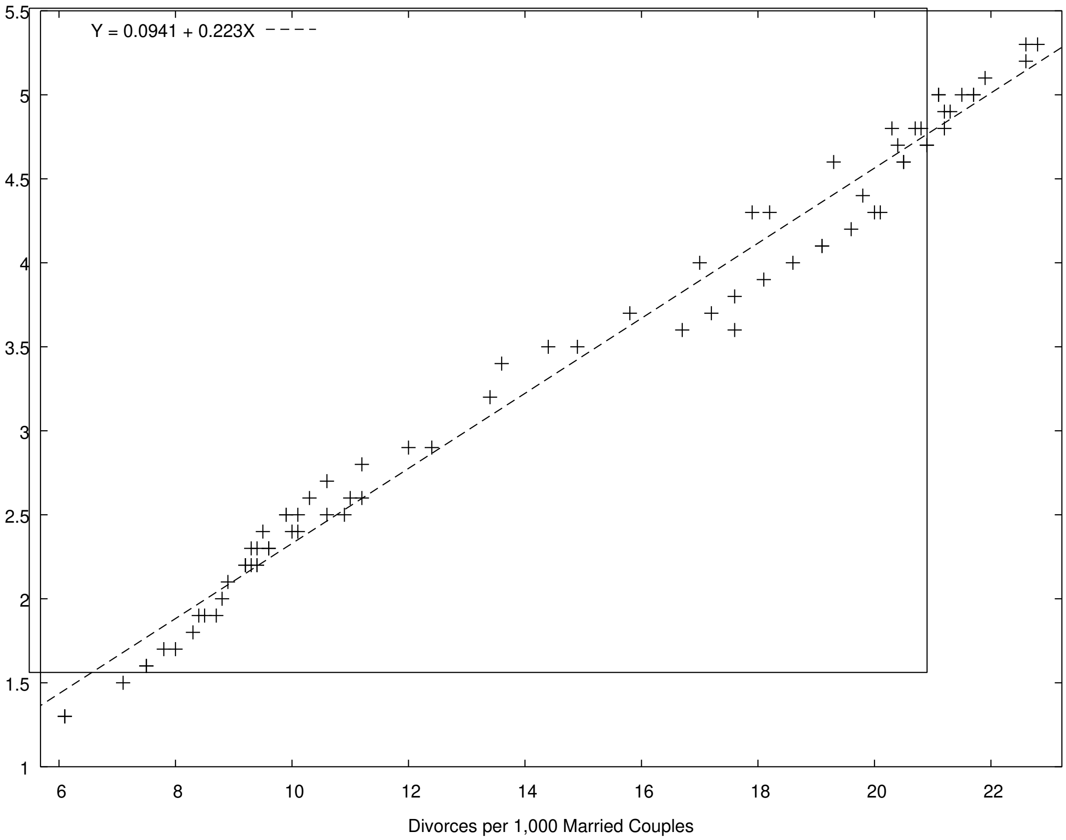
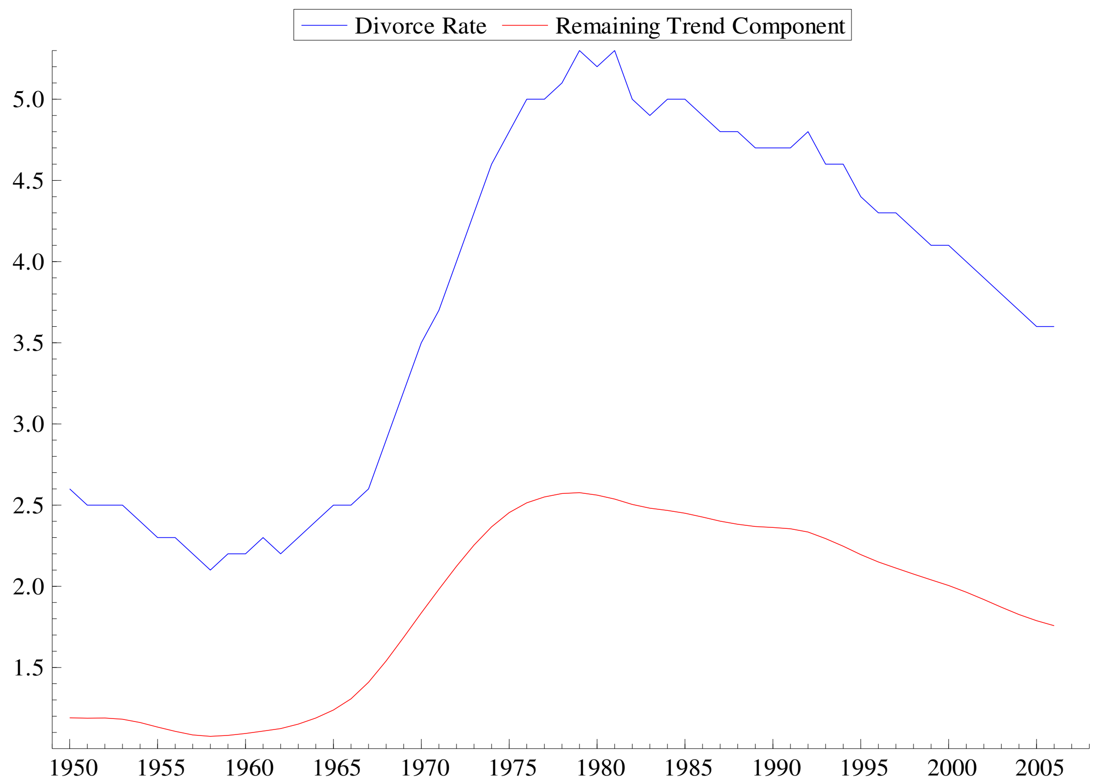


Figure 8—Comparison of Actual Divorce Rate with the Remaining Trend Component Excluding the Vietnam War and Pill Variables



NOT FOR PUBLICATION
APPENDIX FOR REVIEW PURPOSES

A.1 Sources and Construction of the War Variables

Each of the war variables are constructed as casualties relative to deployments. For the Vietnam variable, deaths include not only combat casualties but also deaths due to accidents and death from wounds suffered in Vietnam but occurring elsewhere. 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 U.S. troop deployment in all of East Asia, not just Vietnam. 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 Vietnam War death count comes from http://www.archives.gov/research/vietnam-war/casualty-statistics.html#year: **CACCF Record Counts by Year of Death or Declaration of Death (as of 12/98). The Vietnam War variable is constructed only for the years 1963 to 1975, although deaths are recorded prior to and past that time period.** 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.[20] 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

[20] http://www.archives.gov/research/korean-war/casualty-lists.

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.[21] 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.[22]

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.[23] Troop deployments in East Asia are used for the Korean War. The source of the deployment figures is the same as that used for the Vietnam War.[24] Troop deployments for WWII are culled from Matloff (1990). However, the deployment figures for 1946 that are related to WWII are assumed to be one sixth of all troops stationed overseas during that year.

A.2 Sensitivity Analysis

In this section, we check the sensitivity of some of the key estimation 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 (*fem_ratio*) 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 similar to those of the variable female participation in higher

[21] http://www.history.navy.mil/library/online/ww2_statistics.htm.
[22] http://stinet.dtic.mil/oai/oai?verb=getRecord&metadataPrefix=html&identifier=ADA438106.
[23] http://www.heritage.org/Research/NationalSecurity/cda06-02.cfm.
[24] http://www.history.army.mil/books/wwii/sp1943-44/index.htm (chapter 17, Tables 4 and 5 and Appendix E).

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 apparent by relying on female participation in higher education as a proxy for female labor-force participation. This confirms Figure 10, which identifies a very strong and close relationship between the two variables.

Next, we check whether the chosen definition of the dependent variable, divorces 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 and 7 with divorces per 1,000 married couples as the dependent variable. These estimates are shown in Tables A2 and A3. While the estimated coefficients are larger than those reported in Tables 4 and 7 for the alternative measure of the divorce rate, 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 coefficient estimates in Tables A2 and A3 simply result from the fact that divorces per 1,000 married couples exceed divorces per 1,000 persons by a substantial margin. Table A4 presents the resulting cointegrating equations when divorces per 1,000 married couples is used as the dependent variable. Similar to the results of Table 10, we find only one cointegrating vector, with the same directional relationship and statistical significance for the variables *fem_ratio* and *pill*.

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 strong evidence for the presence of two cointegrating vectors. The two cointegrating equations both normalized on the divorce rate per 1,000 married couples are given as

$$divorce = -23.460 + 0.097 * flfpr \tag{4}$$

$$divorce = 115.776 - 2.037 * flfpr. \tag{5}$$

According to equation 4, the FLFPR and the divorce rate have a very small but positive relationship over time. This is similar to the findings of Bremmer and Kesselring (2004). Equation 5, by contrast, suggests the existence of an order of magnitude larger negative relationship between the divorce rate and the FLFPR, which is consistent with the results of both Tables 6, 9, and 10.

In sum, the sensitivity analyses support the results shown in Sections 5.1 and 5.2. The estimates reported in those sections are not materially affected by the chosen measure of the divorce rate. Likewise, when using the FLFPR instead of female participation in higher education, there is little difference in the estimated effects. Lastly, we can encompass the result of a small positive, long-run relationship between the divorce rate and the FLFPR found by Bremmer and Kesselring (2004) for the years 1960 to 2001. But we show that there are in fact two cointegrating equations for this time period and the dominant one of the two indicates a negative relationship between the divorce rate and the FLFPR. This result again fully supports our main conclusion that female participation in higher education or in the labor force lowers the divorce rate in the long run.

Table A1: OLS Estimates, 1949-2006
(Substituting Female Labor-Force Participation)

	Model 1		Model 2		Model 3		Model 4	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
constant	0.8662	0.001	-0.5422	0.523	-0.7412	0.378	-0.7589	0.388
divorce (-1)	0.8750	0.000	0.8759	0.000	0.8807	0.000	0.8987	0.000
flfpr	-1.7959	0.000	4.3703	0.233	4.6767	0.193	4.9683	0.196
$flfpr^2$			-6.2563	0.082	-6.3264	0.073	-6.6879	0.079
pill	0.6101	0.000	0.5207	0.000	0.4719	0.001	0.4588	0.001
inflation	0.0100	0.117	0.0078	0.211	0.0105	0.096	0.0063	0.321
ygrowth	0.0095	0.080	0.0330	0.016	0.0586	0.004	0.0130	0.611
$ygrowth^2$			-0.0039	0.049	-0.0048	0.017	0.0008	0.776
*das*ygrowth*							0.1809	0.012
Δ*unemp*					0.0470	0.086	0.0527	0.048
Vietnam (-1)	0.1940	0.000	0.1431	0.002	0.1334	0.000	0.1311	0.000
Korea							-0.0138	0.277

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	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
constant	3.0551	0.001	-3.5572	0.217	-3.5053	0.220	-3.7463	0.245
divorce (-1)	0.8400	0.000	0.8383	0.000	0.8395	0.000	0.8512	0.000
fem_ratio	-5.2380	0.010	25.6045	0.049	23.5841	0.069	25.3492	0.091
fem_ratio ^{2}			-33.6512	0.014	-30.7080	0.025	-32.7200	0.039
pill	2.5074	0.000	2.3625	0.000	2.2672	0.000	2.2038	0.000
inflation	0.0716	0.005	0.0530	0.036	0.0626	0.017	0.0516	0.058
ygrowth	0.0428	0.065	0.1014	0.087	0.1870	0.034	0.0500	0.669
ygrowth ^{2}			-0.0108	0.201	-0.0142	0.107	0.0027	0.834
*das*ygrowth*							0.5309	0.090
Δunemp					0.1542	0.183	0.1696	0.152
Vietnam (-1)	0.7056	0.000	0.5151	0.000	0.5007	0.000	0.4870	0.001
Korea							-0.0400	0.482
Unadjusted R^2	0.9945		0.9955		0.9957		0.9959	
Adjusted R^2	0.9939		0.9948		0.9949		0.9950	
Log-likelihood	-24.129		-18.439		-17.360		-15.461	
AIC	62.258		54.879		54.719		54.922	
SBC	76.681		73.443		75.324		79.647	
HQC	67.876		62.102		62.745		64.553	
RESET		0.029		0.440		0.428		0.563
Homoskedasticity		0.020		0.329		0.465		0.383
Normality		0.919		0.058		0.031		0.129
Autocorrelation LM(1)		0.667		0.447		0.529		0.232
Autocorrelation LM(2)		0.635		0.739		0.769		0.467
Autocorrelation LM(3)		0.477		0.340		0.557		0.628
ARCH (1)		0.297		0.386		0.288		0.243
ARCH (2)		0.508		0.701		0.563		0.535
ARCH (3)		0.801		0.722		0.638		0.474
Quandt LR (max)		<0.01		>0.10		>0.10		>0.05
Harvey-Collier (cusum)		0.017		0.956		0.626		0.693

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 lenths 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 2003); 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)

	Model 5		Model 6		Model 7	
	coeff.	*p-value*	*coeff.*	*p-value*	*coeff.*	*p-value*
constant	-7.7222	0.003	-6.9134	0.011	-5.9049	0.011
divorce (-1)	0.7442	0.000	0.7171	0.000	0.7978	0.000
fem_ratio	52.9385	0.000	46.2748	0.000	38.1299	0.001
fem_ratio^2	-66.2416	0.000	-57.5174	0.000	-47.0461	0.000
pill	3.9318	0.004	4.0025	0.002	2.7403	0.008
inflation	-0.0082	0.810	0.0317	0.392	0.0539	0.076
ygrowth	-0.0844	0.133	0.1347	0.065	0.0815	0.265
ygrowth^2	-0.0054	0.008	-0.0175	0.000	-0.0141	0.001
*das*ygrowth*			-0.3827	0.016	-0.3191	0.012
Δ *unemp*	-0.2413	0.020	-0.1340	0.028	-0.1679	0.029
WWII	3.8358	0.000	4.6067	0.000	3.6764	0.000
WWII^2	-1.1111	0.000	-1.4126	0.000	1.1095	0.000
Vietnam (-1)	0.5830	0.000	0.5993	0.000	0.5340	0.000
Korea	0.1058	0.001	0.0747	0.014	0.0471	0.203
d47					-2.8496	0.000
Unadjusted R^2	0.9905		0.9919		0.9939	
Adjusted R^2	0.9887		0.9902		0.9925	
Log-likelihood	-59.341		-53.546		-42.713	
Akaike	144.682		135.091		115.425	
Schwarz Bayesian	175.151		167.905		150.582	
Hannan-Quinn	156.869		148.216		129.487	
RESET		0.208		0.061		0.073
Homoskedasticity		0.000		0.001		0.017
Normality		0.001		0.077		0.155
Autocorrelation LM(1)		0.105		0.431		0.138
Autocorrelation LM(2)		0.206		0.543		0.218
Autocorrelation LM(3)		0.291		0.656		0.295
ARCH (1)		0.001		0.010		0.958
ARCH (2)		0.001		0.026		0.447
ARCH (3)		0.004		0.051		0.633
Quandt LR (max)		<0.01		<0.01		<0.01
Harvey-Collier (CUSUM)		0.903		0.468		0.438

Notes: (-1) denotes a lag order of one. Δ is the first-difference operator. AIC stands for Akaike Information Criterion; SBC stands for the Schwarz-Bayesian Criterion; HQC is the Hannan-Quinn Criterion; RESET is Ramsey's test for correct functional form; Homoskedasticity is White's test; Normality is a test for non-normality of the residuals, and uses normality as the null hypothesis; the tests for autocorrelation are the Bruesch-Godfrey tests for higher-order autocorrelation; ARCH tests for a relationship between the current error variance and its past values; Quandt LR tests for structural breaks; Harvey-Collier tests for parameter stability.

Table A4: Cointegration and VECM Results, 1929 to 2006
(Dependent Variable: Divorces per 1,000 Married Couples)

	Model 1	Model 2	Model 3
Implicit CIV Normalized on Divorce			
constant	60.933	57.382	53.105
	(8.521)	(7.822)	(9.314)
fem_ratio	-85.091	-131.665	-118.041
	(-6.739)	(-6.744)	(-7.807)
pill	27.488		
	(12.336)		
gruber		65.044	
		(10.431)	
friedberg			62.762
			(12.275)
Cointegration Trace Test			
p-values:			
null of rank = 0	0.000	0.001	0.000
null of rank = 1	0.139	0.284	0.321
Statistics and Specification Tests of VECM			
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

Notes: 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).