Club classification of US divorce rates

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Abstract: In this paper, we study the evolution of US divorce rates across states, from 1956 to 1998. Using a cluster algorithm, we identify different groups of states that converge on each other in the evolution of their divorce rates. We find strong empirical support for the club classification. For the whole 1956–1998 period, we find evidence of convergence in divorce rates for the majority of the states, but when we split the sample, a different pattern emerges. In the pre-unilateral divorce reform period (1956–1972), we find that most of the states had converging divorce rates within several convergence clubs, but the club classification of states significantly changed in the post-reform period (1973–1998). Finally, we analyze the explanatory factors of the club classification of states in both sub-periods, using geographical, socio-economic, and demographic variables.

Keywords: Divorce rate, convergence club.

JEL: C12, C22, J12.

1. Introduction

In an influential article, Friedberg (1998) analyzed the effect of the no-fault unilateral divorce laws of the 1970s on the evolution of the US divorce rate. Using a state-based panel of divorce data from 1968 to 1988, Friedberg found that the adoption of unilateral divorce laws had a permanent influence on divorces, accounting for almost one-sixth of the rise in the divorce rate since the late 1960s. Later, Wolfers (2006) replicated Friedberg's exercise, extending the data period (from 1956 to 1998) and adding variables that explicitly model the dynamic response of divorce. Wolfers' results show that the no-fault unilateral divorce reforms had a positive effect on the divorce rate, but the effect was transitory; after a decade, no effect on the divorce rate could be discerned. These findings have been widely accepted, and this methodology has been used by others to analyze the dynamic response of divorce rates in other countries (for instance, González and Viitanen (2009) study the effect of divorce laws on a sample of European divorce rates).[1] In this paper, we do not pretend to explore the transitory or permanent impact of divorce law reforms on divorce rates; rather, we examine whether the transition to more liberal divorce laws implied a convergence of divorce across US states' divorce rates.

The liberalization of divorce laws leads to a decrease in the costs associated with divorce, which can make divorce more feasible. From a theoretical point of view, the number of divorces, and thus the divorce rate, would increase, since a greater number of couples value the now less-expensive divorce over marriage, although, as Becker (1981) argued, divorce law reforms may not affect the probability of marriage breakdown, since they only affect property rights. Following the Coase Theorem, under mutual consent divorce, the party who wants to divorce has to compensate their spouse, in such a way that mutual consent gives considerable power to spouses who do not want a divorce. The change to a unilateral system transfers the right to divorce to the spouse most wanting a divorce. In this case, it is the party who wants to continue married who must compensate the spouse who wishes to leave. When the re-assignment of property rights between spouses is accompanied by transfers between them, no changes in the divorce rate should be observed. Then, if the divorce rates across several states do not

[1] Of course, there are more determinants of divorce; e. g., child custody and child support laws (González-Val and Marcén, 2012), economic growth (South, 1985), price stability (Nunley, 2010), unemployment (Amato and Beattie, 2011), the birth-control pill (Marcén, 2014), and culture (Furtado et al., 2013).

converge in the pre-reform period, the introduction of divorce law reforms would not generate divorce convergence. If the divorce rates in the pre-reform period do converge across some states, then, after the law changes, we should observe the same convergence. Nonetheless, the aggregate divorce rate will change when compensation between spouses is not possible (see, for example, Allen, 2002). In this case, if divorce rates converge in the pre-reform period for a group of states, it would be possible to observe non-convergence in the post-reform period, since divorce rates vary after the law changes, or there may be convergence in different groups, depending on the law reforms passed. A similar response may be possible if the divorce rates do not converge in the pre-reform period.

The study of the evolution of divorce rates is considered important in the literature since divorce has been suggested as having an impact on both women and children. The possibility of divorce may increase female labor force participation, (Michael, 1985; Johnson and Skinner, 1986; Peters, 1986; Parkman, 1992), but it can also affect the economic status of divorced women and their children (Jarvis and Jenkins, 1999; Bedard and Deschênes, 2005). Divorce may also have a negative effect on the psychological well-being of children (Seltzer, 1994; Amato, 2000; Gruber, 2004; Gähler and Palmtag, 2014).

To analyze divorce convergence in the US, we utilize the Phillips and Sul (2007) panel convergence method on a sample of fifty US states and the District of Columbia, for the period 1956 to 1998. This methodology does not require any specific assumptions concerning the stationarity of the divorce rate, our variable of interest, and/or the existence of common factors. This cluster algorithm has been extensively used in the economic literature, for example, to explore convergence in the cost of living across US cities (Phillips and Sul, 2007), price convergence (Fischer, 2012), the historical population convergence of the US cities (González-Val and Lanaspa, 2014), the income convergence of member states of the European Union (Fritsche and Kuzin, 2011; Bartkowska and Riedl, 2012), outcome convergence within the US (Choi and Wang 2015), and even the happiness club convergence in Europe (Apergis and Georgellis, 2014), among others. We add to this literature by exploring whether divorce rates converge across US states.

Our findings suggest that there was convergence in divorce rates across most of the US states. We find empirical evidence of convergence clubs when we use the whole

sample in the analysis, and most of the states (35) are classified within the same convergence club. However, after dividing the sample into two periods, pre-reform (1956-1972) and post-reform (1973-1998), our results show that, in the pre-reform period, most of the states were converging in divorce rates within several convergence clubs, but the club classification significantly changed in the post-reform period. Then, even if the effects of divorce law reform were transitory, and disappeared after a decade, as suggested by Wolfers (2006), these results would point to the reforms as a main determinant of changes in the club classification. Finally, by using a logit model, we analyze the factors affecting the likelihood of two states to belong to the same club, using geographical, socio-economic, and demographic variables.

Section 2 presents the data used. In Section 3, we apply the cluster algorithm to identify different groups of states that converge with each other, and we explore potential determinants of the test results. Section 4 concludes.

2. Data

We use the same dataset as Wolfers (2006), testing the dynamic response of the divorce rate to a change in the legal regime that governs divorce, using data on the states' divorce rate between 1956 and 1998. The divorce rate (DR) is defined as the annual absolute number of divorces per thousand inhabitants in each state (the source is the *Vital Statistics of the United States*).

This is known as the crude divorce rate and represents the standard measure of the level of, and changes in, divorce. Nevertheless, this rate could be affected by the marital status structure of the populations to it relates. Divorce rates may be low either because marriage rates are low, or because marriages are less likely to end in divorce. To examine this issue, we could have used another measure of divorce rates, defined as the annual number of divorces per 1000 married population, but this analysis would have been less reliable due to the scarcity of data on the total number of marriages, which is only available when each census is collected, normally every 10 years (see Furtado et al., 2013; Marcén, 2014).

Table 1 incorporates information on the year in which no-fault unilateral reforms were passed. Since 1968, 31 states introduced no-fault unilateral reforms, with most of those reforms taking place during the late 1960s and 1970s, with only two exceptions of reforms implemented in the 1980s, following Gruber's (2004) classification.

Unfortunately, information on the divorce rate is not available for all states during all the period considered (see Table 1). As we explain in detail below, for three states, (California, Indiana, and Louisiana) the analysis cannot be carried out because of data limitations. Table 1 also incorporates a summary of statistics of the divorce rates across states. There are important variations in the average divorce rates, with 15 states having an average divorce rate greater than 5 (the highest being Nevada, at 18.6), 27 states with an average divorce rate between 3 and 5, and 9 states with divorce rates lower than 3, (the lowest was 2.3, in the states of Massachusetts and New York). The considerable dissimilarities in the gap between the minimum and maximum divorce rates in each state are also notable. For example, in the case of New Mexico, we observe a minimum divorce rate of 1.5 and a maximum of 9.1, while in Pennsylvania the minimum was 1 and the maximum 3.5.

3. Results

3.1. Convergence clubs across states

To explore the evolution of US divorce rates across states, we apply a cluster algorithm that allows us to identify different groups of states that converge with each other in the evolution over time of their divorce rates. The cluster procedure is based on the log $t-$test (Phillips and Sul, 2007, 2009), which focuses on the evolution over time of idiosyncratic transitions in relation to the common component. Other papers have studied the evolution of divorce rates from a time series perspective, such as that of González-Val and Marcén (2012), where the path of the common component of state divorce rates is analyzed through panel unit root tests.[2] This new approach is different from that of prior empirical studies of growth convergence clubs, such as the regression tree analysis used by Durlauf and Johnson (1995) and the predictive density of data used by Canova (2004) to identify different clusters of countries or regions. The procedure of Phillips and Sul focuses on the evolution over time of divorce rate relative to the average rather than on individual divorce rate by state. Thus, their methodology enables us to identify the relative transitions that occur within subgroups, and to measure these transitions against the correlative of a common trend (Phillips and Sul, 2009). The regression model of the log $t-$test is

[2] They find that the unit root null hypothesis cannot be rejected for most states, even when one or multiple structural breaks are allowed.

$$\log \frac{H_1}{H_t} - 2\log(\log t) = \beta_0 + \beta_1 \log t + u_t, \qquad \text{for } t = T_0,...,T \tag{1}$$

where $\frac{H_1}{H_t}$ is the cross-sectional variance ratio, H_t is the transition distance, $H_t = N^{-1}\sum_{i=1}^{N}(h_{it}-1)^2$, and h_{it} is the relative transition coefficient, defined as $h_{it} = \frac{\log DR_{it}}{N^{-1}\sum_{i=1}^{N}\log DR_{it}}$ (DR_{it} is the divorce rate of state i at time t.). These relative transition coefficients exclude the common growth path (μ_t) by scaling, measuring state i's transition element relative to the cross-section average. This means that h_{it} traces out state i's individual trajectory relative to the average, so Phillips and Sul (2009) call h_{it} the 'relative transition path.' Moreover, h_{it} also measures for each state i the departure of divorce rate from the common growth path μ_t in relative terms.

Thus, Eq. (1) simply represents a time series regression; the null hypothesis is convergence across all states, and the alternatives include no convergence and partial convergence among subgroups of states. As the t-statistic of the test refers to the coefficient β_1 of the $\log t$ regressor in Eq. (1), the test is called the '$\log t$' convergence test. It is important to note that not only the sign of the coefficient β_1 of $\log t$, but also its magnitude, measures the speed of convergence, so that the higher the value of the coefficient the faster the rate of convergence.

The cluster procedure performs the $\log t$ test for each of the groups and stops when the group of remaining states does not satisfy the convergence test. First, it defines an initial core primary group, and other groups are then formed according to certain criteria that maximize the value of the t-statistic. A much more detailed explanation of the constructive steps of the procedure can be found in Phillips and Sul (2007, 2009).

Figure 1 shows the path of all the states and demonstrates that it is not easy to infer any specific pattern. However, it seems clear that around the beginning of the 1970s there is a rise in the trend of divorce rates in all the states, followed by a subsequent fall in the second half of the 1980s. Wolfers (2006) identified the 1969–1977 period as the reform period, in which 28 states adopted unilateral divorce, but in

most of these states the law changed in the first three years of the 1970s (see Table 1),[3] so the vertical red line indicates the intermediate year 1972.[4] The fall in rates in the 1980s, a decade later, coincides with the temporal duration of the effects of the unilateral reforms estimated by Wolfers (2006).[5]

Table 2 shows the results of applying the cluster algorithm to our sample of states.[6] We consider three periods: 1956–1972 (pre-reform period), 1973–1998 (post-reform period) and 1956–1998 as a whole. The "club" column shows the number of states that are members of each group. The distribution of states within groups can be found in Table 3. From 1956 to 1998, the algorithm classifies states into three groups, revealing three different steady divorce rates in the US. Three remaining states (California, Indiana, and Louisiana) are excluded because the algorithm requires a balanced panel dataset, and Wyoming is not classified into any club. In each group, the estimated coefficient $\hat{\beta}_1$ is significant, strongly supporting the club classification. The majority of the states (35) are classified into group 3, indicating convergence among the majority when the whole period is considered. Figure 2 shows the path over time of the divorce rate of the states in each convergence club.

The number of groups and their composition significantly changes if we split the sample into two periods. When we consider the pre-reform period (1956–1972) the cluster procedure identifies five groups. California, Indiana and Louisiana are excluded due to data limitations and Wyoming is not classified into any club, again. The coefficient $\hat{\beta}_1$ is significant for all groups, supporting the club classification. The results show that, in this pre-reform period, most of the states, classified into five different convergence clubs, were converging in divorce rates within each club. Figure 3 displays the evolution of the divorce rates by club. However, when we focus on the post-reform period (1973–1998), we see a different picture. The algorithm classifies the states into six homogeneous groups (four of them contain 10 states). Two remaining states are not classified into any club (Nevada and Wyoming), and for these the convergence

[3] González-Val and Marcén (2012) find that many of the structural breaks detected in the state series are located in that brief period (1970–1973).
[4] We have tried different intermediate years to split the sample and the qualitative results are maintained.
[5] Although González-Val and Marcén (2012) suggest that the long-run effect of divorce law reforms on the divorce rate observed by Wolfers (2006) may be the result of both unilateral reforms and changes in the aftermath of divorce.
[6] The estimations were performed with the Gauss code kindly provided by Donggyu Sul on his webpage.

hypothesis is rejected; California, Indiana, and Louisiana are excluded because of missing data. Figure 4 shows the path over time of the divorce rate of the states by club.

Figure 5 illustrates the transitional movements across state groups over time. This figure lists the states in each club, based on the clustering results obtained for the pre-reform period, 1956 to 1972. The figure also displays the changes in club membership that had taken place by 1998, considering the club classification obtained when considering the whole period, 1956 to 1998. The results show that there are transitions to clubs 2 and 3 when the post-reform period is included in the sample. There are 29 new members in Club 3, coming from all the other clubs. About 50% of the states in Club 1 move to Club 3 over the later time period; another 5 states move to Club 2. Most of the states initially classified in Club 2 move to Club 3, and all the states in Clubs 4 and 5 (which eventually disappear), move to Club 3. Finally, all the states from Club 3 remain in the same group. Summarizing, there is strong empirical support for a transition in divorce rates across states, and evolving membership of convergence clubs. Our findings suggest that unilateral divorce reforms could lead to convergence across states as most of them end up in the same convergence club (Club 3) after the divorce law reforms. This may indicate that divorce law reforms helped to homogenize divorce rates across states.

3.2. Determinants of the club classification

To analyze how state characteristics affect the likelihood that any pair of states become members of the same convergence club, we estimate a logit model. The dependent variable is a dummy that takes the value 1 if two states belong to the same club for each possible pair, and 0 otherwise.

We use several explanatory variables. The first two are related to the geographical distance between each pair of states, sharing a border, and a dummy variable that takes the value 1 if both states are in the same region.[7] Second, we consider political and legal differences; several dummy variables control for whether the unilateral divorce and the joint custody reform laws were approved the same year in the two paired states, and whether the governors in both states are from the same party (in the initial year). Third, we measure the demographic similarity between states, considering the absolute difference in several variables: state population (ln scale), the

[7] We consider eight regions: New England, Mideast, Great Lakes, Plains, Southeast, Southwest, Rocky Mountain and Far West.

percentage of women who finished high school, the percentage of women with one or more years of college, the female labor-force participation rate, and the marriage rate. Finally, to account for economic differences between states, we include the absolute difference in the per capita personal income (ln scale). All the differences are calculated from the average values of the period considered. The data on governor by state is taken from Sobel and Raimo (1978). Data on educational variables, and on the female labor-force participation rate are obtained from the Current Population Survey and from the Integrated Public Use Microdata (Ruggles et al. 2010).[8] The marriage rate is obtained from the Vital Statistics of the US. Lastly, population and per capita personal income data are taken from the US Department of Commerce (Bureau of Economic Analysis).

Table 4 shows the results of the logit estimations. We analyze the club classification in two sub-periods: 1956–1972 (pre-reform period), 1973–1998 (post-reform period). The whole period (1956–1998) is not considered, since most of the states are classified within the same convergence club. We report the estimated coefficients, their corresponding robust standard deviation (in parentheses), and the marginal effects. The results of the models, including all explanatory variables, are shown; nested models adding different sets of the variables do not change the results.[9]

We obtain some results common to both sub-periods. First, the location of states, measured by the variables sharing a border and same region, has no effect, so the club classification is not driven by geographical patterns. Second, the dummies controlling for whether the unilateral divorce and the joint custody reform laws were approved in the same year in the two-paired states, are also not significant. Previously, we argued that the main explanatory factor of the change in the composition of groups over time is unilateral divorce law reform, so we split the sample in 1973, the year in which most of the states approved the reform. Nevertheless, the dummy variable is not significant, because in the pre-reform period the states had not yet approved the reform, and in the post-reform period most had already changed the law. Third, the difference in the average marriage rate is the only significant variable in both sub-periods. The estimated coefficient is negative, pointing to the similarity in marital patterns as a key determinant of the club classification. Finally, differences in state population, and the percentage of women who finished high school, are not significant.

[8] Information on Alaska, District of Columbia, and Hawaii is not available, so all the two-state pairs including any of these three states are excluded from the sample.
[9] Results of the nested models are available from the authors upon request.

We also obtain evidence of the particular influence of certain variables in one of the two sub-periods. Thus, if two states have a governor from the same party in 1973, they are more likely to be members of the same convergence club in the post-reform period. Furthermore, there is a significant negative correlation between female labor-force participation rates and the probability of belonging to the same club in the 1973–1998 period. Thus, two states with a similar share of female workers have a greater probability of belonging to the same club. In the pre-reform period, larger differences in the percentage of women with one or more years of college, and variations in personal income, reduce the probability of being in the same club.

4. Conclusion

In this paper, we examine the evolution of US divorce rates across states, using data for the period 1956-1998. We utilize a cluster algorithm that allows us to identify dissimilar convergent growth of divorce rates by state, finding strong support for the club classification. The empirical evidence shows that, in the whole period considered, there was a clear convergence in divorce rates across states, since the majority (35) are classified within the same convergence club. Nevertheless, there are significant variations in the number of groups and their composition when we split the sample in two sub-periods: pre-reform (1956–1972) and post-reform (1973–1998).

Our findings suggest that the liberalization of divorce laws implied a divorce convergence across the US. When we examine transitions between groups over time, we observe that most of the states move from one club to another. The analysis of the transitions reveals that unilateral divorce reform may lead to convergence across states, as most end up in the same convergence club. Therefore, divorce law reforms helped to homogenize divorce rates across states. Finally, we study what factors can help to explain the club classification in the two sub-periods. No geographical pattern can be deduced and, although some of the determinants change over time, the similarity in marital patterns, measured by differences in the marriage rate, is a key determinant of the club classification; two states with a similar marriage rate have a greater probability of belonging to the same club in both sub-periods.

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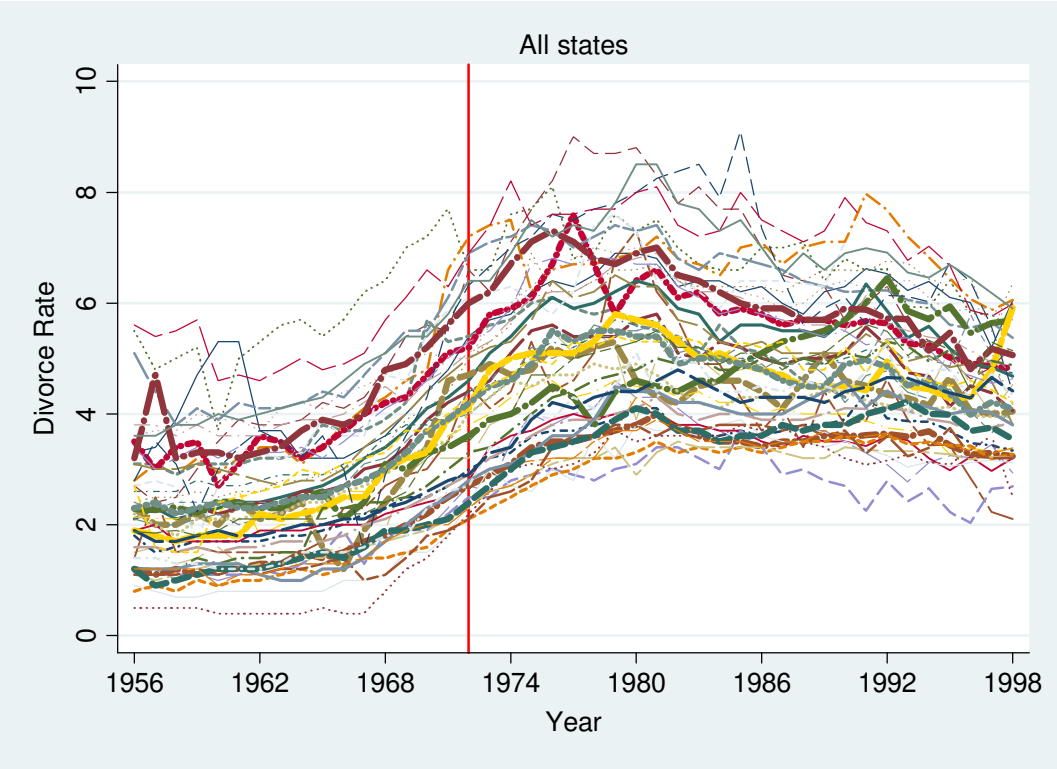
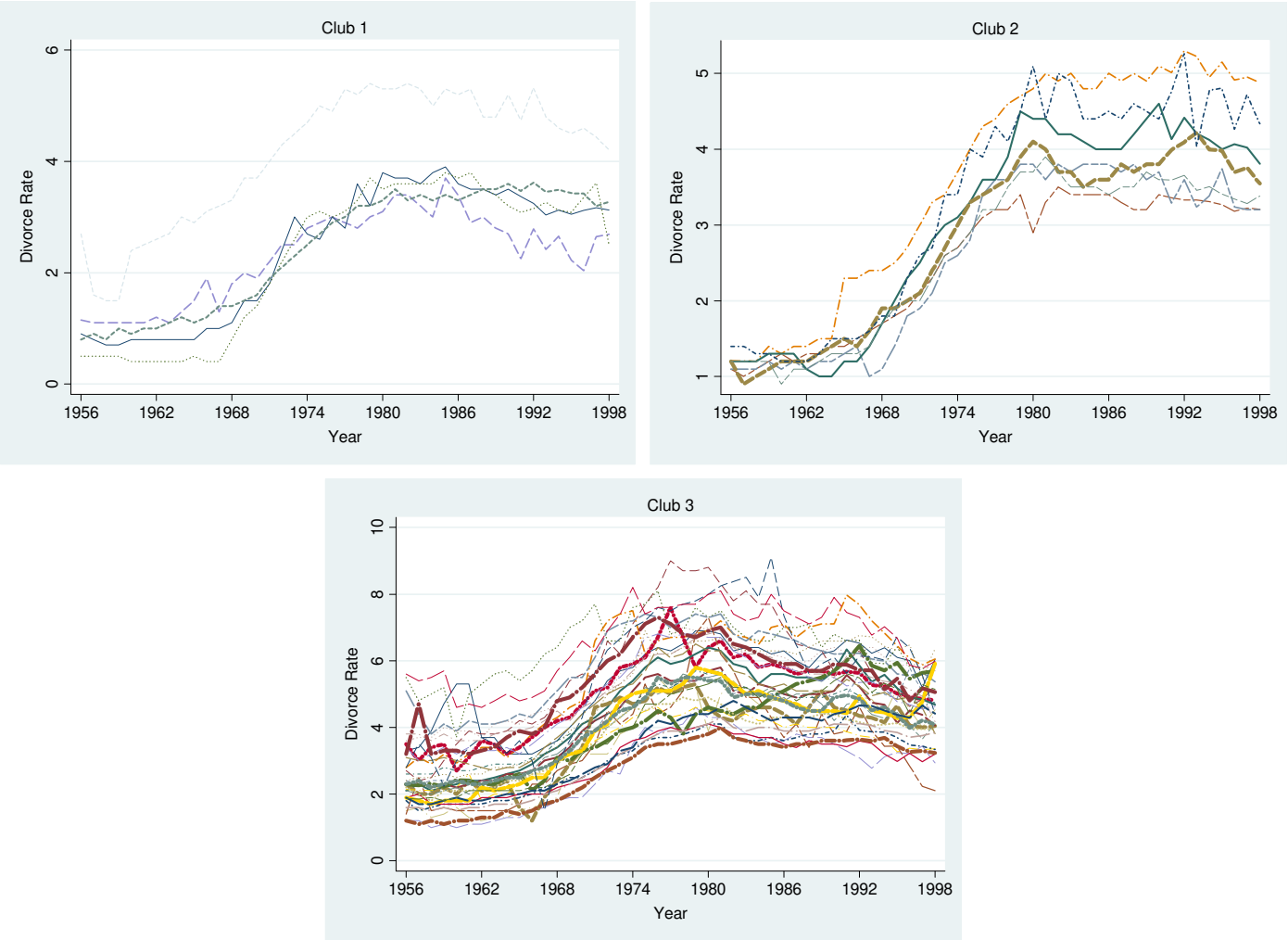
Figure 1. States' divorce rate paths, 1956–1998



Note: Nevada is not included because its extremely high divorce rate distorts the graph.

Figure 2. States' convergence clubs, 1956–1998



Note: Nevada is not shown within Club 3 because its extremely high divorce rate distorts the graph.

Figure 3. States' convergence clubs, 1956–1972

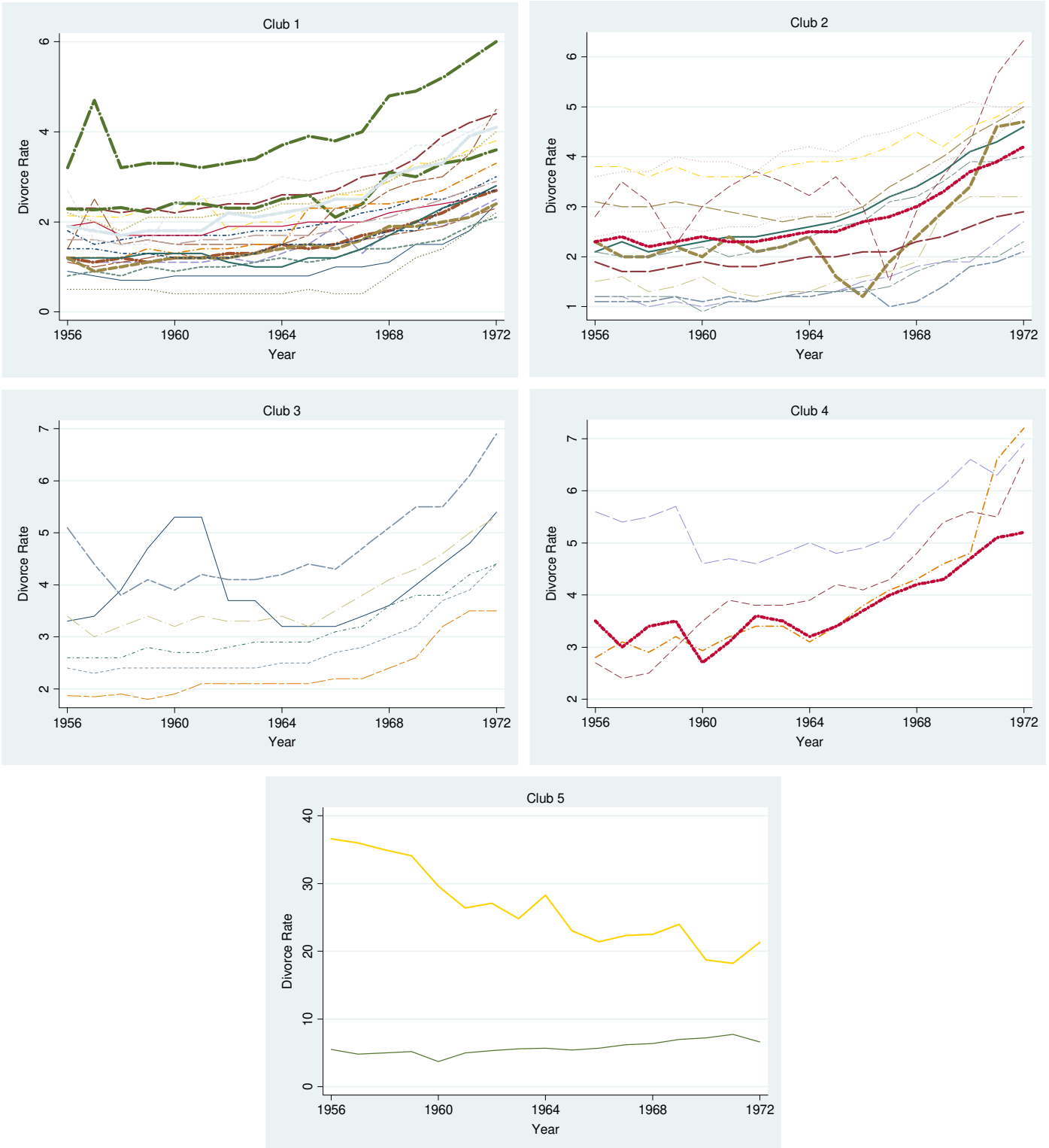


Figure 4. States' convergence clubs, 1973–1998

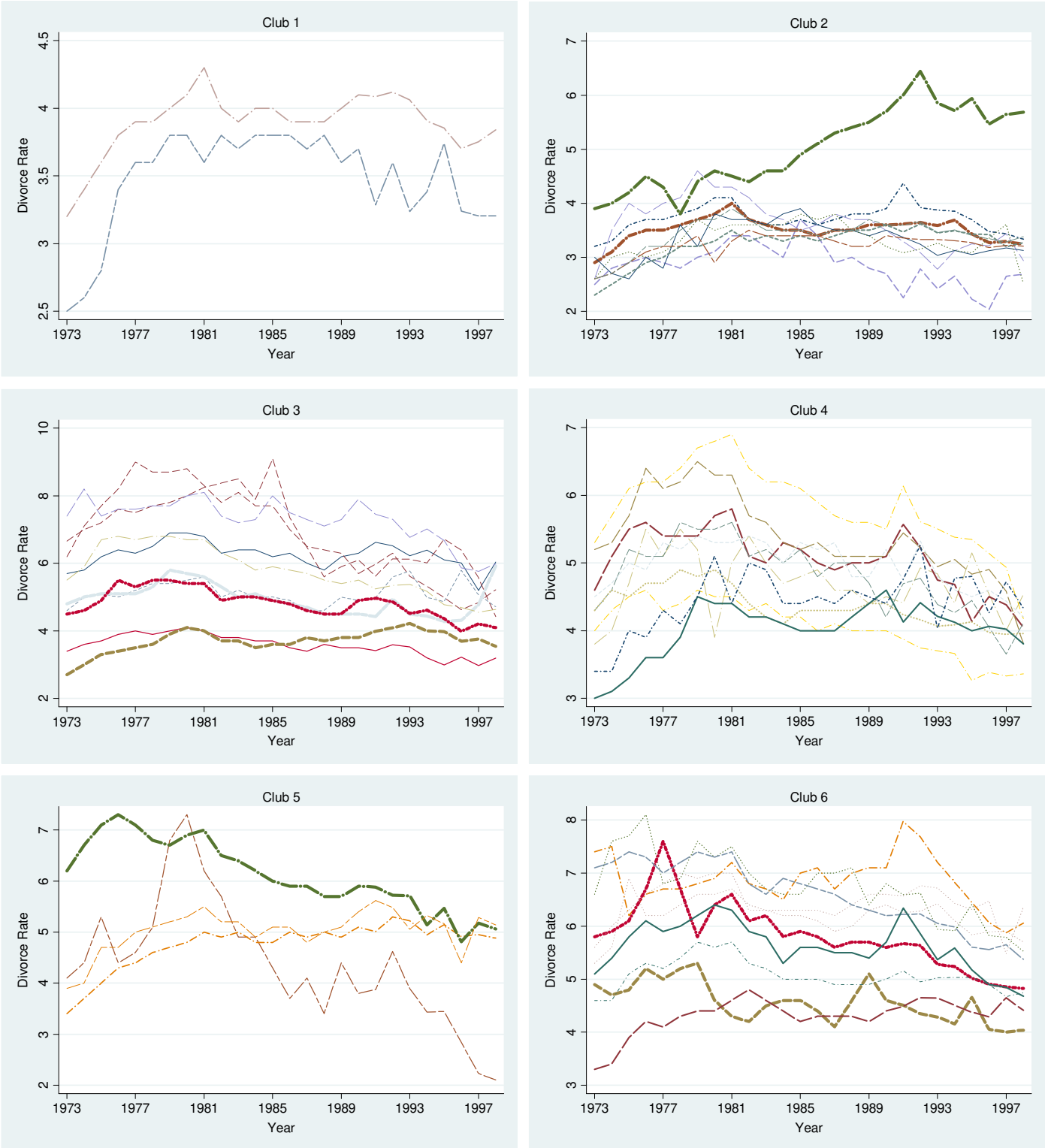
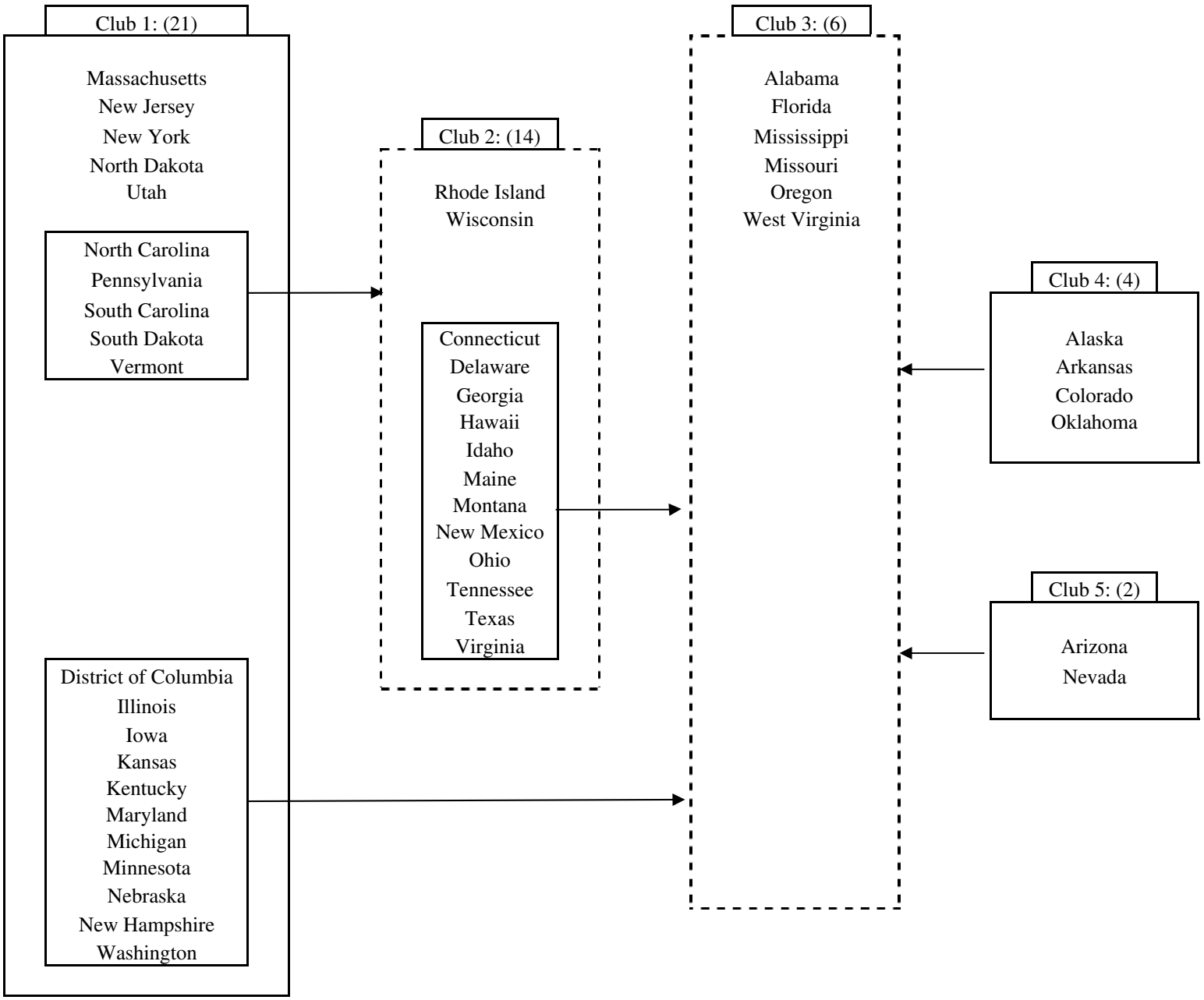


Figure 5. Transitioning to Clubs 2 and 3



Notes: Transitions between groups. Club classification in 1956–1972 versus club classification in 1956–1998.

Table 1. State divorce rates: Descriptive statistics

State	Unilateral	Observations	Average	Max.	Min.	State	Unilateral	Observations	Average	Max.	Min.
Alabama	1971	1956–1998	5.4	6.9	3.2	Montana	1973 (1975)	1956–1998	4.6	6.5	2.7
Alaska	1935	1956–1998	5.8	9.0	2.4	Nebraska	1972	1956–1998	3.1	4.3	1.5
Arizona	1973	1956–1998	6.4	8.1	3.7	Nevada	1967 (1973)	1956–1998	18.6	36.6	8.1
Arkansas	.	1956–1998	5.7	8.0	2.8	New Hampshire	1971	1956–1998	4.0	5.9	1.7
California	1970	1956–1990	4.6	6.1	2.9	New Jersey	.	1956–1998	2.4	3.9	0.7
Colorado	1972 (1971)	1956–1998	5.0	7.6	2.7	New Mexico	1933 (1973)	1956–1998	5.6	9.1	1.5
Connecticut	1973	1956–1998	2.8	4.6	1.0	New York	.	1956–1998	2.3	3.8	0.4
Delaware	1968 (.)	1956–1998	3.6	5.5	1.2	North Carolina	.	1956–1998	3.6	5.3	1.2
District of Columbia	.	1956–1998	3.5	7.3	1.4	North Dakota	1971	1956–1998	2.5	3.6	0.8
Florida	1971	1956–1998	5.9	7.4	3.8	Ohio	.	1956–1998	4.0	5.5	2.2
Georgia	1973	1956–1998	4.6	6.4	2.1	Oklahoma	1953(.)	1956–1998	6.6	8.2	4.6
Hawaii	1972 (1973)	1956–1998	3.7	5.3	1.2	Oregon	1971 (1973)	1956–1998	5.0	6.8	3.0
Idaho	1971	1956–1998	5.5	7.0	3.6	Pennsylvania	.	1956–1998	2.5	3.5	1.0
Illinois	.	1956–1998	3.5	4.6	1.8	Rhode Island	1975 (1976)	1956–1998	2.6	3.8	1.0
Indiana	1973	1956–1990	5.1	7.6	1.7	South Carolina	.	1956–1998	3.0	4.6	1.0
Iowa	1970	1956–1998	3.1	4.4	1.5	South Dakota	1985	1956–1998	2.8	4.2	0.9
Kansas	1969	1956–1998	4.2	5.8	2.2	Tennessee	.	1956–1998	5.1	6.9	2.4
Kentucky	1972	1956–1998	4.1	6.4	2.1	Texas	1970 (1974)	1956–1998	5.1	6.9	3.6
Louisiana	.	1971–1983	3.4	4.3	2.5	Utah	1987 (.)	1956–1998	4.1	5.4	1.5
Maine	1973	1956–1998	4.0	5.6	2.0	Vermont	.	1956–1998	3.3	5.3	1.2
Maryland	.	1956–1998	3.0	4.1	1.7	Virginia	.	1956–1998	3.4	4.8	1.7
Massachusetts	1975	1956–1998	2.3	3.7	1.1	Washington	1973	1956–1998	5.3	7.3	3.2
Michigan	1972	1956–1998	3.7	4.9	1.8	West Virginia	.	1956–1998	4.0	5.6	1.8
Minnesota	1974	1956–1998	2.7	4.0	1.1	Wisconsin	1978 (.)	1956–1998	2.6	3.9	0.9
Mississippi	.	1956–1998	4.2	5.8	2.3	Wyoming	1977	1956–1998	6.1	8.5	3.6
Missouri	.	1956–1998	4.3	5.7	2.6						

Note: Year of unilateral divorce is from Gruber (2004) and from Friedberg (1998) in parentheses.

Table 2. State convergence clubs

1956–1998		1956–1972		1973–1998	
Club	$\hat{\beta}_1$ (t-statistic)	Club	$\hat{\beta}_1$ (t-statistic)	Club	$\hat{\beta}_1$ (t-statistic)
1 [5]	-0.633 (-1.572)	1 [21]	0.223 (0.421)	1 [2]	-1.099 (-0.601)
2 [7]	-1.406 (-1.576)	2 [14]	0.324 (0.835)	2 [10]	-1.037 (-1.099)
3 [35]	-0.132 (-0.531)	3 [6]	-0.036 (-0.100)	3 [10]	-0.812 (-0.903)
		4 [4]	0.649 (0.496)	4 [10]	-0.882 (-0.778)
		5 [2]	-0.004 (-0.012)	5 [4]	-0.157 (-0.408)
				6 [10]	-1.229 (-1.566)

Notes: The numbers in brackets are the number of states. The corresponding t-statistic in the regression is constructed in the usual way by using HAC standard errors. At the 5% level, the null hypothesis of convergence is rejected if the t-statistic < -1.65. All the t-statistics reported are higher than -1.65, indicating that we cannot reject the null hypothesis at 5% in any case.

Table 3. States within clubs

State	Club (1956–1972)	Club (1973–1998)	Club (1956–1998)	State	Club (1956–1972)	Club (1973–1998)	Club (1956–1998)
Alabama	3	3	3	Montana	2	4	3
Alaska	4	3	3	Nebraska	1	1	3
Arizona	5	6	3	Nevada	5		3
Arkansas	4	6	3	New Hampshire	1	3	3
California				New Jersey	1	2	1
Colorado	4	6	3	New Mexico	2	3	3
Connecticut	2	2	3	New York	1	2	1
Delaware	2	4	3	North Carolina	1	5	2
District of Columbia	1	5	3	North Dakota	1	2	1
Florida	3	6	3	Ohio	2	3	3
Georgia	2	6	3	Oklahoma	4	3	3
Hawaii	2	6	3	Oregon	3	3	3
Idaho	2	6	3	Pennsylvania	1	2	2
Illinois	1	4	3	Rhode Island	2	1	2
Indiana				South Carolina	1	4	2
Iowa	1	2	3	South Dakota	1	3	2
Kansas	1	4	3	Tennessee	2	6	3
Kentucky	1	2	3	Texas	2	4	3
Louisiana				Utah	1	4	1
Maine	2	4	3	Vermont	1	4	2
Maryland	1	3	3	Virginia	2	6	3
Massachusetts	1	2	1	Washington	1	5	3
Michigan	1	4	3	West Virginia	3	5	3
Minnesota	1	2	3	Wisconsin	2	2	2
Mississippi	3	3	3	Wyoming			
Missouri	3	6	3				

Table 4. Estimation results from logit model

	1956–1972		1973–1998	
Share a border	0.043	(0.264)	0.172	(0.324)
	0.008		0.009	
Same region	0.168	(0.235)	0.054	(0.302)
	0.031		0.003	
Same year unilateral divorce reform	-0.194	(0.354)	-0.367	(0.474)
	-0.033		-0.016	
Same year joint custody reform	0.161	(0.246)	0.211	(0.284)
	0.030		0.012	
Governor from the same party (initial year)	0.009	(0.147)	0.506***	(0.177)
	0.001		0.026	
Abs. Dif. Average Ln Population	0.138	(0.090)	-0.113	(0.109)
	0.025		-0.006	
Abs. Dif. Average % Female Some College	-7.311*	(3.764)	3.958	(3.694)
	-1.305		0.206	
Abs. Dif. Average % Female High School	5.422	(3.527)	0.354	(3.402)
	0.968		0.018	
Abs. Dif. Average Female labor force participation	0.028	(0.021)	-0.060**	(0.026)
	0.005		-0.003	
Abs. Dif. Average Marriage rate	-0.014***	(0.004)	-0.227***	(0.064)
	-0.002		-0.012	
Abs. Dif. Average Ln Per capita personal income	-1.143*	(0.594)	0.968	(0.830)
	-0.204		0.050	
Observations	1128		1128	
Wald (p-value)	25.57	(0.007)	30.32	(0.001)
Pseudo R^2	0.024		0.049	

Notes: Parameter estimates, (robust standard errors) and marginal effects. All regressions include a constant.