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Cohort Variations in Suicide Rates among Families of Nations

An Analysis of Cohorts Born from 1875 through 1985

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Abstract

Using data on age-specific suicide death rates from 19 modern nations and cohorts born as early as 1875–9, we find that two indicators of cohort-related social capital, relative cohort size and percentage of nonmarital births, are positively and significantly related to suicide rates. These effects are significantly stronger in the English-speaking family of nations, which have historically provided fewer political and social supports to families and children. The analytic model, an extension of the Age-Period Cohort Characteristic model, which utilizes hierarchical linear modeling, provides strong controls for age and period effects as well as for autoregressive effects within cohorts. Our data allow us to include older age groups and data from a wider range of countries than previous studies.

Key words: families of nations • nonmarital births • relative cohort size • social capital • suicide

Distinguishing age, period, and cohort effects remains a central concern of demographers. One approach to this problem is to use cohort characteristics to represent cohort effects (Kahn and Mason, 1987; Mason et al., 1973; O'Brien et al., 1999). Such models have demonstrated that variations in age-specific rates of homicide, in the United States, and of suicide, both in the United States and cross-culturally, are strongly related to cohort variations in relative size and family structure (O'Brien and Stockard, 2002, 2003; O'Brien et al., 1999; Savolainen, 2000; Stockard and O'Brien, 2002a, 2002b). These effects persist even after introducing strong controls for age and period, and can account for recent upturns in youth suicide and homicide rates in the United States and a number of other countries.

While the results from these studies of the United States have been replicated using different methods and periods of analysis, they have limitations due to data availability. Data are available on key independent variables for only a restricted range of cohorts. This limits the number of cases at older age ranges and the time span over which researchers can examine cohort effects.

Cross-cultural analyses indicate that the effects of cohort-related social capital can be modified by social context (Stockard and O'Brien, 2002a). An analysis with a sample of cohorts born in 1915 and later in 14 western industrialized countries indicated that cohort effects tend to be weaker in societies with a political tradition of collectivism and support for families and children and stronger in countries with a relatively large amount of change in traditional family roles. The study, however, was limited, not just by the range of cohorts involved, but, also by the contextual measures, which were derived only from the post-Second World War period. To the extent that these measures tap political and cultural conditions from time periods after a cohort's formative years, they may provide less than optimal measures of the context in which cohort members were raised.

This article addresses these concerns by utilizing data on cohorts born as early as 1875. We examine the influence of relative cohort size and cohort childhood family structure on age-specific suicide deaths from 1950–2000 for cohorts from 19 industrialized nations and explore the extent to which these relationships are modified by national political/cultural contexts. We introduce a new analytic strategy to help model the effects of cohorts on suicides (see O'Brien et al., 2003), treating age-period-specific suicide rates as nested within cohorts and both of these variables as nested within countries in a hierarchical model (Raudenbush and Bryk, 2002).

THEORETICAL BACKGROUND AND RELATED LITERATURE

The growing body of work focusing on cohort effects and lethal violence builds upon a broadly defined Durkheimian perspective. This literature suggests that variations in lethal violence rates across birth cohorts are related to the social integration and regulation that they experience, or *cohort effects*. At the same time, recent evidence suggests that characteristics of the broader society, or *contextual effects*, can modify the influence of these demographic characteristics.

The Cohort-Related Social Capital Hypothesis and Limitations of Earlier Work

A number of studies have documented the relationship of two demographic characteristics – family structure and relative cohort size – to age-period-specific homicide and suicide rates independent of age and period (O'Brien and Stockard, 2002, 2003; O'Brien et al., 1999; Savolainen, 2000; Stockard and O'Brien, 2002a, 2002b).[1] It is suggested that these characteristics reflect 'cohort-related social capital' and influence the social integration and regulation that birth cohorts experience. This influence occurs in at least three different ways: 1) fewer financial resources, which may result from more children within a cohort and/or fewer adults within a household; 2) less attention and supervision for children, with adult resources spread more thinly among children; and 3) a stronger influence of peers, resulting from the lower involvement of adults in children's lives.

In part, cohort effects reflect the aggregation of individual effects – the result of more children growing up in larger families or single parent families. In addition, however, building on Coleman's (1990) notions of social capital and social closure, all members of a birth cohort are affected by these characteristics, no matter what the size or composition of their own family. When cohorts are relatively large, all children, not just those from larger families, experience the lower levels of supervision and adult resources as well as the heightened influence of peers. When cohorts have more single parent families, networks of parents tend to have less closure, with a lower likelihood of parents being acquainted with parents of their children's friends and peers becoming a more important influence on all children, regardless of their own family background.[2]

Previous research has noted the possibility of different results for males and females. Males might be more likely than females to be negatively affected by larger cohort sizes because of their historically greater involvement with the labor market (Pampel, 2001). They may also be more likely than girls to interact in larger peer groups, to seek out same sex peers as companions, and to avoid adult companionship (Fagot, 1994; Maccoby, 1990, 1995). To the extent that the influence of relative cohort size and nonmarital births occurs because of their relationship to young peoples' greater peer, rather than adult, interactions, males might be more likely than females to be negatively affected by larger cohort sizes. To date, there has been some, limited, support for this hypothesis. The effect of relative cohort size has been stronger for males than for females in analyses of suicide in the United States and internationally and for analyses of homicide deaths in the United States for nonwhites, but not for whites. In the United States the effect of family structure has been stronger for males than females for analyses of homicide deaths for both whites and nonwhites, but in the analysis of suicide deaths for only nonwhites (O'Brien and Stockard, 2002; Stockard and O'Brien, 2002b). In the analysis of international data, the effect of family structure was stronger for females than for males (Stockard and O'Brien, 2002a). Only a few of these differences, however, were statistically significant.

Some researchers also have hypothesized that there might be interaction effects between age and cohort effects, resulting in a stronger influence of cohort effects at younger ages than at older ages. For instance, some (Kahn and Mason, 1987; Steffensmeier et al., 1992) argue that the effects of relative cohort size should be especially pronounced for those who are young. Clinicians who study suicide have suggested that adolescents may be more vulnerable to adversities than adults (Diekstra, 1995, citing Platt, 1984 and Rutter, 1980). Others suggest that larger cohorts may have more social and peer support in old age, as well as more political influence, thus reducing suicide rates when they become older (McCall and Land, 1994; Pampel, 2001; Pampel and Peters, 1995). Recent analyses of suicide and homicide rates using the APCC model have found only limited support for such interaction effects.[3]

While the accumulating body of work regarding lethal violence and cohort-related social capital provides strong and consistent evidence, it has several limitations related to data availability. All analyses to date have been restricted to birth cohorts born between 1915–9 and later, conforming to the time period for which reliable data on family structure are available for cohorts from the United States. As a result, analyses of suicide and homicide deaths have included limited data on older age groups: only the cohort born between 1915 and 1919 had reached the age of 75–9 by 1995 (the last year of data in most of the earlier analyses) and only two cohorts had reached the age of 70–4.

This small number of cases at older age ranges presents at least three potential limitations to the previous work. First, one of the central tenets of cohort theory is the 'lasting effects principle,' which suggests that cohort effects persist throughout the life span. According to this principle, the effects of diminished cohort-related social capital on lethal violence should be observed at all ages throughout the life span. While this tenet appears to be supported with the analyses done to date, these tests have been limited because of relatively few cases at older age groups. A more valid examination of the lasting effects principle requires additional cases at older ages. Second, the lack of cases at older age ranges results in less than optimal tests of possible interaction effects between cohort characteristics and age. This problem is especially acute regarding tests of the interactions of cohort characteristics with older age groups. Third, including a broader range of cohorts may provide additional insights into the comparative influence of family structure and relative cohort size. Analyses that have included cohorts born between 1915 and 1984 have consistently found that the influence of family structure has been stronger than that of relative cohort size. While these analyses have included cohorts with a fair amount of variation in measures of both cohort characteristics, an analysis with a more extensive range of cohorts, and greater variation in these independent variables, should provide an important test of these findings.

We address these limitations by utilizing data on cohorts born as early as 1875 and as late as 1984. This provides many more cases in the older age ranges. Based on earlier tests of this hypothesis, which have included a more limited set of periods and cohorts, we hypothesize that, throughout the set of nations in our analysis, relatively large cohorts and those that have a higher incidence of nontraditional family structures experience less integration and regulation and thus will be more likely to have higher suicide rates. As noted above, some literature suggests that these cohort effects will be stronger for males than for females and that they may be stronger at younger ages and weaker at older ages. Including a broader range of cohorts in our analysis provides a stronger test of these hypotheses.

Contextual Effects and Families of Nations

Societies may develop ways to supplement or replace diminished cohort-related social capital. In addition, other conditions within a society could heighten the impact of low levels of cohort-related social capital. In other words, national contextual factors may alter the effects of cohort-related social capital. Using a sample of 14 modern industrialized countries, including the United States and Canada, Stockard and O'Brien (2002a) found that the effects of relative cohort size and family structure on age-period-specific suicide rates were weaker in societies that provided alternative sources of social integration and regulation (greater levels of support to mothers and families). They were stronger in societies that had experienced more rapid change in traditional family roles. While these results were strong and held across a variety of models, the quantitative measures of social context were based on data from 1955 and later, even though cohorts in the analysis were often born in much earlier periods.

This article addresses this concern by using a more qualitative measure of contextual effects, based on the notion of 'families of nations,' groups of countries that share geographic, linguistic, cultural, and/or historical characteristics. These nations often share similar legal traditions and public policies (Castles, 1993), as well as similar social problems and issues. We assume that these 'familial' characteristics have been relatively constant over the last century and a half, the periods in which the cohorts that we study were born. Among the countries that we examine, legal scholars often delineate two major groups: those that follow the tradition of English common law, sometimes referred to as 'Anglo-American common law' and those that adhere to 'Continental European Civil Law' (Therborn, 1993: 246). The latter group includes three general subgroups: the Romanist family, the Germanic family, and the Nordic family.[4]

The English-speaking countries have historically provided the least support for families and children. For instance, the Anglo-Saxon legal tradition defined illegitimate children as *filius nullius*, or 'nobody's child' and deprived an illegitimate child of any legal rights, including a claim on its mother (Teichman, 1982; Wimperis, 1960). It was not until the 1970s that most of these countries legally affirmed an illegitimate child's right to inherit on an equal basis with legitimate children (Therborn, 1993; see also Baldock, 2003, Heffernan, 2003; Rubellin-Devichi, 1997). In addition, the English-speaking countries tend to have relatively low levels of state support for families and a general tendency for governments to deliberately avoid interfering in individuals' private lives. The teenage pregnancy rates in many of the English-speaking countries are relatively high (Coney, 1993). Many of these teenage mothers do not marry. As a result, the problems associated with high numbers of nonmarital births are increased by having larger numbers of very young parents with relatively few economic and educational resources. In recent years, some of the English-speaking countries – particularly New Zealand, Canada, and Australia – have experienced substantial change in traditional family roles, making it more difficult to replace declining levels of cohort-related social capital. At the same time values and attitudes have, at least in some countries, remained relatively traditional, which could promote greater strain (Bradshaw, 1997; Cornia and Danziger, 1997; Drummond, 1997; Kennedy and McCormack, 1997). Not surprisingly, when compared to the other highly developed nations, all of the English-speaking countries tend to have higher proportions of their children living in poverty (Bradshaw, 1997; Cornia, 1997; Hantrais, 1997; Kamerman and Kahn, 1997; Kennedy and McCormack, 1997; Maclean, 1994).

The legal traditions of the Romanist nations have also been very restrictive regarding nonmarital births. They have subscribed to the 'Code of Napoléon' of 1804, which forbade an illegitimate child to search for its father and establish a legal relationship. Most of the countries retained these strict prohibitions until late in the 20th century, with Belgium not giving illegitimate children an equal right to inheritance or a right to paternity proceedings until the 1980s (Rubellin-Devichi, 1997; Saraceno, 1997; Therborn, 1993). With the exception of the Netherlands, the Romanist countries tend not to have 'collectivist' political institutions, political systems that provide universal programs for social assistance and strong programs of social protection. They do, however, often provide strong societal wide systems of support for families and children through means such as child allowances, government sponsored preschools, and tax schemes that favor families (e.g. Bergmann, 1996; Dumon, 1997; Hantrais, 1997; Kamerman and Kahn, 1981; Muller-Éscoda and Vogt, 1997).

While not as restrictive as the Romanist countries, those in the Germanic tradition severely restricted the rights of illegitimate children for many years. Until the 1970s illegitimate children in Austria were totally excluded from inheritance possibilities if there were 'legitimate' heirs; in Switzerland, illegitimate children could only inherit half of what legitimate heirs were entitled to (Therborn, 1993). Neither of the Germanic countries in our sample (Austria and Switzerland) has strong collectivist political institutions and, while Austria tends to provide relatively strong support for women and children, Switzerland is far less likely to do so and has a long history of traditional family structures and a political structure that supports this traditionalism (Popenoe, 1988).

The Nordic nations are most likely of all the nations in our analysis to have legal and social structures that might supplement cohort-related social capital. In 1915 Norway was the first European country to pass legislation that specifically protected the rights of a child born out of marriage, guaranteeing the child the right to a father, his name, his support and his inheritance and giving the state the responsibility for establishing paternity. The other Nordic countries soon followed this pattern (Rubellin-Devichi, 1997; Teichman, 1982; Therborn, 1993; Wimperis, 1960). The Nordic countries are most likely of those in our sample to have collectivist political institutions and also have a long tradition of providing extensive and relatively generous supports to families and working women, both through guarantees of legal equality and supportive policies and programs such as provisions for part-time work and high quality, state-supported and guaranteed childcare. In contrast to the English-speaking countries, the welfare of children is typically viewed as an important concern of the state, resulting in very low rates of child poverty and excellent child health (Hantrais, 1997; Hort, 1997; Kamerman and Kahn, 1981, 1997; Kröger et al., 2003). Age of marriage in the Nordic nations has typically been relatively high and the number of teen parents relatively low. Children born outside of marriage are more likely in the Nordic countries, than in the English-speaking countries, to be raised within a stable consensual union (Granström, 1997; Meisaari-Polsa, 1997).

Based on results obtained in earlier cross-national tests of the cohort-related social capital hypothesis (Stockard and O'Brien, 2002b), we expect that the influence of cohort-related social capital will be modified by national political and cultural characteristics. Specifically, we hypothesize that the effects of cohort-related social capital will be reduced in nations that provide more supportive atmospheres for families and children. The English-speaking countries have, historically, provided relatively low levels of support for families and children, and we expect the strongest effects of cohort-related social capital in those nations. In contrast, the Nordic countries have provided the greatest levels of support, and we expect the weakest influences in those nations.

METHODOLOGY

To examine our hypotheses we use data from 19 industrialized countries: six English-speaking countries: Australia, Canada, Ireland, New Zealand, the United Kingdom, and the United States; seven countries within the Romanist family: France, Belgium, the Netherlands, Spain, Portugal, Greece, and Italy; two within the Germanic family: Austria and Switzerland; and four in the Nordic family: Denmark, Sweden, Norway, and Finland. We have data on each of our two key measures of cohort-related social capital (relative cohort size and family structure) for cohorts born from the late 19th century forward in 15 of these countries. In the remaining countries we have data for cohorts born from the early 20th century on. All of the nations included in our sample have long had well-established systems for gathering vital statistics and are populous enough to ensure that suicide rates for five-year age groups are statistically reliable from one year to the next. While they are all currently industrialized democracies, they vary in their demographic, economic and political characteristics, their social and political history, their cultural attitudes, and their rates of suicide.[5]

Data and Measures

Our data on suicide rates come from the World Health Organization (WHO). To parallel the growing body of work on cohort effects on lethal violence and to obtain reliable estimates we use age-specific rates for five-year-age groups (O'Brien and Stockard, 2002, 2003; O'Brien et al., 1999; Savolainen, 2000; Stockard and O'Brien, 2002a, 2002b). Data for all except one of the countries used in the analysis were available from the early to mid-1950s through 2000. Except for the earliest and last years, the data on suicides broken down by age are for 1950, 1955, 1960, . . . , 2000.[6] We omit from our analysis death rates for people younger than the age of 15, when suicide is rare, and for people over the age of 75, when population sizes become much smaller and rates are less reliable. Thus, the oldest cohort in our analysis (70–4 years of age in 1950) was born in 1875–9, and the youngest cohort in our analysis (15–19 years of age in 2000) was born in 1980–4.

We examine two general aspects of cohort-related social capital: family structure and relative cohort size. Our measure of family structure is the percentage of nonmarital births (NMB) within a cohort. This measure is used because it can be directly linked to birth cohorts and, unlike measures such as the percentage of children in single parent homes at a given age, is available for birth cohorts born in the 19th and early 20th century. The *Demographic Yearbook* published by the United Nations supplied data on nonmarital births for cohorts born after about 1945 and Wimperis (1960) provided data for earlier years. To obtain the percentage of nonmarital births for each cohort born after 1945, we summed the appropriate percents and divided by the number of years. For earlier cohorts the data were provided for five-year groupings of birth cohorts.

We operationalize relative cohort size (RCS) as the percentage of the adult population that was in a five-year cohort when the cohort was young (O'Brien, 1989; O'Brien and Stockard, 2002; Savolainen, 2000; Stockard and O'Brien, 2002a, 2002b). This operationalization is preferred for several reasons. It compares the size of the birth cohort with the size of the cohorts that preceded it, the measurement is made at a crucial age when the birth cohort is entering the job market (leaving home, potentially ready to marry, bear children, etc.), and the age range covered by this measure includes the parental generation of adults who were mainly responsible for rearing members of the birth cohort. Yearly population figures were obtained from the WHO for cohorts born after 1950; information for earlier cohorts was obtained from national censuses, usually from 10-year intervals, except when wars or national crises resulted in longer gaps (Mitchell, 1992).[7] We were usually able to measure relative cohort size for two different five-year birth cohorts during a single decennial year: the percentage of those 15–64 who were 15–19 when the cohort was 15–19 and the percent of those 10–59 who were 10–14 when the cohort was 10–14. When both measures were available for a cohort, we used the average of the two figures in our analyses. The correlation between these two measures when both were available was .99. In a very few cases we had to base the measure on other age groups.[8] Most commonly this measure involved the percentage of those 5–54 who were 5–9 when the cohort was 5–9, but again, when data for this and other measures of the cohort's relative size were available the measures were very highly correlated (r = .98 for the measure based on ages 15–19; and r = .99 for the measure based on ages 10–14).[9]

Table 1 presents information for the United Kingdom to illustrate the data that are used in our analysis. The rows and columns indicate period and age and each cell contains the age-period-specific suicide rate. Cohort 1 was born between 1875 and 1879, cohort 2 between 1880 and 1884, and so on. The last cohort in our analysis (cohort 22) was born between 1980 and 1984. By following a particular cohort diagonally through the table, one can see the way in which cohorts move through the space of time and age. (Cohort 1 has one observation and is the upper right hand cell of the table, cohort 2 is on the second diagonal, and so on, with cohort 22 represented by the bottom left-hand cell.) The marginal at the bottom of the table contains two values that remain the same for each cohort over time. The top and bottom entries represent, respectively, the relative cohort size and the percentage of the cohort members who were born to unwed mothers.

Data like those in Table 1 are examined for all 19 countries with each age-period-specific suicide rate representing a case. The full matrix of data (for all 22 cohorts) is available for 10 countries (Australia, Denmark, United Kingdom, Finland, France, Italy, the Netherlands, New Zealand, Norway, and Switzerland). Data for cohorts 2–22 (born in 1880 and later) are available for Austria, Belgium, and Sweden; data for cohorts 3–22 are available for Portugal, for 5–22 for Spain, for 9–22 for the United States, and for 10–22 (born in 1920 and later) for Canada, Greece, and Ireland.[10] Earlier studies of the effects of cohort-related social capital have only included cohorts born in 1915 and later (cohorts 9 and later) to parallel data available for the United States. Examination of Table 1 illustrates how this restriction severely limits the number of cases at older age ranges – only three cases within the 70–4 age group, four within 65–9, etc. The broader range of cohorts included in the present analysis provides, for the majority of countries in the study, an equal number of observations of age groups at each period.

We explore the extent to which the results obtained in other analyses of the relationship of cohort-related social capital and lethal violence are replicated when this much broader range of cohorts and a greater number of countries are included in the analysis. This analysis provides an extensive and broad ranging test of the theory of cohort effects on lethal violence. Based on earlier studies we hypothesize that lower levels of cohort-related social capital, as measured by larger relative cohort sizes and more nontraditional family structures, will be related to higher suicide rates. Our broad range of cohorts allows improved tests

Table 1 Age-period-specific suicide rates and measures of cohort characteristics, United Kingdom, 1950–99

	Age												
Period	**15–19**	**20–4**	**25–9**	**30–4**	**35–9**	**40–4**	**45–9**	**50–4**	**55–9**	**60–4**	**65–9**	**70–4**	
1950	**12**	**11**	**10**	**9**	**8**	**7**	**6**	**5**	**4**	**3**	**2**	**1**	
	1.87	3.93	4.64	5.38	8.06	10.67	14.84	17.82	21.42	21.86	25.37	24.22	
1955	**13**	**12**	**11**	**10**	**9**	**8**	**7**	**6**	**5**	**4**	**3**	**2**	
	1.64	3.49	6.84	6.73	8.33	11.07	15.26	19.89	22.88	25.68	27.95	26.79	
1960	**14**	**13**	**12**	**11**	**10**	**9**	**8**	**7**	**6**	**5**	**4**	**3**	
	2.3	5.92	7.48	8.98	10.58	11.43	15.5	18.16	22.41	23.39	23.49	23.29	
1965	**15**	**14**	**13**	**12**	**11**	**10**	**9**	**8**	**7**	**6**	**5**	**4**	
	3.05	5.91	8.42	8.59	11.04	14.54	14.77	16.98	20.48	21.41	20.78	22.23	
1970	**16**	**15**	**14**	**13**	**12**	**11**	**10**	**9**	**8**	**7**	**6**	**5**	
	2.25	5.97	5.99	7.49	8.55	10.6	11.87	12.64	15.08	15.46	16.3	16.42	
1975	**17**	**16**	**15**	**14**	**13**	**12**	**11**	**10**	**9**	**8**	**7**	**6**	
	2.69	6.96	7.64	7.76	8.41	10.05	11.88	12.00	11.53	14.19	12.12	14.55	
1980	**18**	**17**	**16**	**15**	**14**	**13**	**12**	**11**	**10**	**9**	**8**	**7**	
	3.03	6.95	8.39	9.35	11.55	12.56	13.06	13.91	15.11	14.25	14.86	15.34	
1985	**19**	**18**	**17**	**16**	**15**	**14**	**13**	**12**	**11**	**10**	**9**	**8**	
	2.97	7.63	9.36	10.93	10.79	11.98	13.84	13.71	14.23	14.66	14.29	13.79	
1990	**20**	**19**	**18**	**17**	**16**	**15**	**14**	**13**	**12**	**11**	**10**	**9**	
	3.73	10.25	11.02	10.35	11.23	10.87	12.28	10.25	10.48	9.05	10	9.55	
1995	**21**	**20**	**19**	**18**	**17**	**16**	**15**	**14**	**13**	**12**	**11**	**10**	
	3.96	9.1	11.01	10.27	10.87	11.95	8.82	9.79	8.64	7.04	7.83	7.09	
1999	**22**	**21**	**20**	**19**	**18**	**17**	**16**	**15**	**14**	**13**	**12**	**11**	
	4.21	9.27	10.91	11.35	11.36	10.70	10.49	9.13	9.16	7.41	6.66	6.94	
RCS	9.43	9.06	10.35	12.1	12.79	11.54	10.71	12.05	10.41	9.67	9.56	10.21	10.98
Cohort	**22**	**21**	**20**	**19**	**18**	**17**	**16**	**15**	**14**	**13**	**12**	**11**	**10**
NMB	14.42	9.82	8.54	8.18	6.42	4.86	4.82	5.5	6.8	4.2	4.3	4.5	4.3
RCS	11.05	12.51	13.32	13.96	13.66	14.42	14.58	15.87	16.19				
Cohort	**9**	**8**	**7**	**6**	**5**	**4**	**3**	**2**	**1**				
NMB	5.5	4.3	4	3.9	4.1	4.2	4.6	4.8	4.7				

Note: Within each row in the body of the table the cohort number is on top and bolded, and the age-period-specific suicide rate is below. In the bottom row the top figure is the measure of RCS, the middle figure is the cohort number, and the bottom figure is the measure of NMB. Cohort 1 was born 1875–79.

of hypotheses regarding interaction effects between cohort characteristics and age, and we also examine results separately for males and females. We explore contextual effects and specifically hypothesize that the influence of cohort-related social capital will be stronger in English-speaking nations and weaker in Nordic countries given their historical, political, and cultural traditions regarding support to families and their greater tolerance for nonmarital births.

Analysis

We analyze these hypotheses by applying hierarchical linear modeling (HLM) to an Age-Period-Cohort Characteristic model (APCC), using PROC MIXED in SAS. Our model may be conceived as a three-level hierarchical model, with age and period effects nested within cohorts, which are in turn nested within countries. We use dummy variables to represent groupings within the 'families of nations.' Grouping the data for all nations in a family together, we first regress the dependent variable (the logged age-period-specific suicide rates for males and for females) on dummy variables for age-groups and periods and include randomly varying intercepts for cohorts and countries as well as a first-order autoregressive term for cohorts. The dummy variables for age and period provide strong controls for these effects (Pampel and Peters, 1995). Factors that change over time and whose effects are constant across age groups within each country, such as variations in the media, income inequality, levels of divorce, or political strife, are controlled through the period dummy variables. Factors related to age that are constant across periods such as the tendency for older people to commit suicide more than the young, are controlled through the use of dummy variables for age categories. These dummy variables control for the main effects for variables that are not included in the model to the extent that they are related to either period or age. Including cohorts and countries as random effects allows us to assess the variance between cohorts and countries in suicide rates that remains after controlling for age and period. The autoregressive term for cohorts allows us to model dependencies between the residuals within cohorts.

Following other studies in the area, we use the natural log of the age-period-specific suicide rates as the dependent variable in all of our analyses. The use of the natural log can compensate for the skew of the dependent variable and is important for theoretical reasons since our interest focuses on proportional changes in suicide death rates.[11] Because, however, the natural log of zero is undefined, the use of the natural logs produces unique problems when suicide rates equal zero. This condition occurred with some of our data points.[12] To include as many cases as possible within our analysis, yet preserve the substantive essence of very low rates in these cases, we substituted the lowest rate recorded for any age group of that gender in that country for the zero, a procedure that probably resulted in estimates for these age-groups that were slightly elevated, and, of course, slightly higher estimates of the suicide rates. To examine any bias that this procedure might have introduced we also analyzed the data in two other ways: 1) adding one to the rate for all cases and then taking the natural log of this value; and 2) counting the zero rates as missing values. In each case the substantive results are quite similar to those reported in this article. These results are available on request.

Following O'Brien and Stockard (2003) we also log the measures of the two cohort characteristics. This transformation ensures that proportionate shifts in the percentage of nonmarital births or relative cohort size are associated with proportionate changes in suicide rates. It also facilitates the interpretation of coefficients ('b') associated with the cohort characteristics: a one percent change in the independent variable will be associated with a 'b' percent change in the dependent variable.

We test hypotheses regarding the interaction of cohort characteristics and age by including interaction terms of both cohort effects (LNRCS and LNNMB) and a young age group (ages 15 to 29) and an older age group (ages 60–74), which results in a set of four interaction terms. Previous literature hypothesizes that the effects of cohort characteristics might be somewhat stronger for younger age groups, but weaker for older ages. To examine the possibility that cohort characteristics have a stronger effect for males than for females we conduct separate analyses for the two sex groups.

Below we report descriptive statistics and then explore analytic models that are progressively more complex. In our analytic models, we first examine the baseline model with the dummy variables for age and period, an autoregressive term within cohorts, and randomly varying intercepts for cohorts and countries. Second, we add the measures of cohort characteristics and, third, the interactions of age with these cohort characteristics. We then add the dummy variables for families of nations and interactions of the cohort characteristics with the family of nations measures. We examine the relative fit of the models using both the Bayesian Information Criterion (BIC) and the likelihood ratio chi-square statistics. BIC values assess the fit of a model correcting for the number of degrees of freedom used by the model and the number of cases in the analysis. Lower values indicate a better fit. The chi-square statistic is used to test for the significance of the improvement of fit from one model to another. Preliminary analysis indicated that the influence of the Germanic and Nordic countries on the effects of RCS and NMB was quite similar, and they were grouped together as the reference category when dummy variables for English and Romanist nations were used in the analysis.[13]

RESULTS

Descriptive Results

Table 2 presents the average suicide rates for 15–19-year-olds and 50–4-year-olds for each country and the dataset as a whole. Females almost always have substantially lower rates of suicide than males.[14] There is wide variation, however, between the countries and the families of nations in their suicide rates. Among both males and females, the Germanic and Nordic countries have higher average rates than either the English or Romanist countries. Within the older age group the rates in the English-speaking countries and the Romanist countries are similar. In contrast, within the younger age group, the rates are somewhat higher within the English-speaking countries than in the Romanist nations. This is especially true among young men, for whom the average rate over the years studied is 9.4 per 100,000 in the English-speaking countries, contrasted to a value of 4.5 in the Romanist nations. The maximum value for teenage males in the English-speaking countries (32.99 for 15–19 year olds in New Zealand) is substantially higher than that found in the Romanist countries (14.47 for males in Belgium).

Table 2 Suicide rates by sex for 15–19-year-olds and 50–4-year-olds for each country in the analysis, 19 countries, 1950–2000

		15–19-year-olds		50–4-year-olds	
		Males	**Females**	**Males**	**Females**
English-speaking:					
Australia	Mean	9.76	3.12	26.44	10.89
	Range	4.32–17.15	1.00–6.62	19.21–40.44	5.30–20.44
Canada	Mean	12.63	3.27	26.87	9.74
	Range	2.61–21.38	1.08–5.15	21.85–30.81	6.85–14.46
Ireland	Mean	5.96	1.12	14.17	4.83
	Range	0.00–19.83	0.00–4.31	4.97–29.28	0.00–9.24
New Zealand	Mean	13.16	4.74	22.61	10.88
	Range	1.62–32.99	0.00–10.60	16.52–35.92	5.02–23.72
UK	Mean	4.17	1.54	17.71	10.49
	Range	2.50–6.45	0.79–2.13	13.62–24.40	4.19–16.37
USA	Mean	10.76	2.61	28.10	9.78
	Range	3.50–18.05	1.32–3.73	21.36–35.17	6.54–13.09
Germanic/Nordic:					
Austria	Mean	17.08	5.48	54.70	21.51
	Range	9.58–21.94	1.35–12.09	37.95–76.74	12.48–29.42
Denmark	Mean	8.50	3.24	53.69	29.38
	Range	3.66–11.94	0.72–6.24	22.77–78.75	12.00–44.70
Finland	Mean	20.03	4.47	57.70	18.90
	Range	7.31–40.10	2.56–6.96	46.99–78.61	14.70–25.27
Norway	Mean	9.89	2.81	23.60	8.52
	Range	0.97–20.57	0.00–8.43	16.84–33.23	4.61–19.57
Sweden	Mean	8.51	4.48	41.51	16.78
	Range	4.84–13.11	0.91–7.67	22.31–53.47	10.28–31.71
Switzerland	Mean	15.01	5.25	46.33	19.25
	Range	9.37–22.87	3.65–7.34	33.99–62.54	13.45–31.56
Romanist:					
Belgium	Mean	8.46	3.26	38.08	17.87
	Range	3.83–14.47	1.06–5.64	31.24–45.07	13.04–24.85
France	Mean	6.44	2.91	42.27	14.47
	Range	3.89–9.09	1.94–4.31	32.48–50.23	10.93–18.66
Greece	Mean	2.34	2.38	8.08	3.03
	Range	0.59–3.24	0.54–7.79	2.49–17.51	1.33–4.65
Italy	Mean	3.05	1.92	15.56	6.18
	Range	2.06–4.40	0.78–3.44	10.41–22.89	4.70–7.47
Netherlands	Mean	4.12	1.31	17.00	12.30
	Range	1.20–7.36	0.52–2.44	11.90–23.07	8.39–15.49
Portugal	Mean	3.91	3.44	23.02	5.45
	Range	2.61–5.19	0.91–5.55	6.96–35.31	1.53–7.78
Spain	Mean	3.54	1.24	13.54	4.69
	Range	1.55–6.23	0.39–2.41	10.39–16.80	3.77–5.76
Total	Mean	8.80	3.06	29.95	12.34
	Range	0.00–40.10	0.00–12.09	2.49–78.75	0.00–44.70

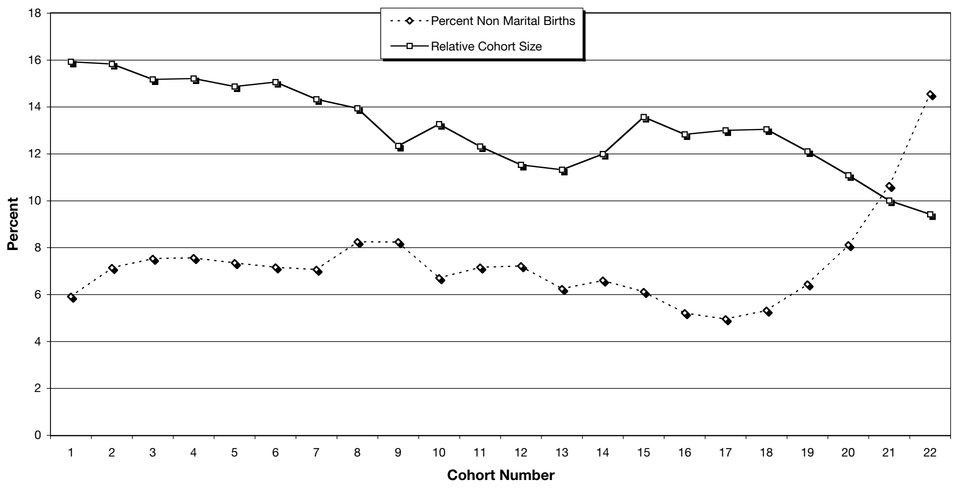
Table 3 gives average values of NMB and RCS for each country. On average, the Germanic and Nordic countries have the highest percentage of nonmarital births, although there is substantial variation within this group, with Austria, Denmark and Sweden having relatively high values, and Switzerland one of the lowest. The Romanist and English-speaking countries have similar average levels of nonmarital births, with Portugal and France having the highest values and Greece, the Netherlands and Ireland the lowest. Many of the nations have had substantial variation in levels of nonmarital births over the years covered in our research, as shown by the range in values, with the highest levels of variation in Sweden and Austria, in the Germanic/Nordic group, as well as the English-speaking New Zealand. Variations in relative cohort size across the families of nations are less striking than those with nonmarital births. The nations with, on average, the largest youthful cohorts are Portugal, Ireland, and Spain; those with the smallest are Belgium, France, and United Kingdom.

Figure 1 displays the average values of both nonmarital births and relative cohort size from the earliest cohort (cohort 1, born in 1875–9) to the most recent group (cohort 22, born in 1980–4). The data regarding nonmarital births indicate relatively stable rates from cohort 1 through cohort 7 (born 1905–09),[15] but a

Table 3 Percentage of nonmarital births and relative cohort size for each country in the analysis

	Percentage of nonmarital births			Relative cohort size		
Country	**Mean**	**Range**	***N***	**Mean**	**Range**	***N***
English-speaking:						
Australia	5.15	3.90–13.76	130	12.87	10.27–16.54	130
Canada	4.36	2.15–14.78	85	13.64	9.61–15.69	94
UK	5.13	3.90–14.42	130	11.57	9.06–16.19	130
Ireland	2.95	1.76–6.24	85	14.36	12.11–16.76	130
New Zealand	5.37	2.30–22.96	130	13.54	10.50–18.76	130
USA	4.55	2.10–19.61	94	12.88	10.42–15.33	94
Germanic/Nordic:						
Austria	18.70	11.66–26.50	117	11.75	8.39–16.14	117
Denmark	10.22	6.96–37.94	130	12.70	7.64–16.33	130
Finland	6.96	4.14–13.82	130	13.51	8.96–16.14	130
Norway	6.54	3.52–17.76	130	12.94	9.01–17.32	130
Sweden	13.76	9.46–44.22	129	11.78	8.75–15.98	130
Switzerland	4.06	3.40–5.30	130	11.77	8.22–15.27	130
Romanistic:						
Belgium	4.43	2.06–8.80	117	11.25	8.66–15.43	117
France	8.25	5.94–14.40	130	11.53	8.31–13.55	130
Greece	1.12	0.00–1.56	60	14.22	9.58–18.12	101
Italy	4.64	2.00–7.60	130	13.28	7.77–15.98	130
Netherlands	2.01	1.24–5.90	130	13.79	8.38–16.68	130
Portugal	12.26	7.20–15.60	116	14.64	9.64–17.04	117
Spain	4.76	1.46–6.10	100	14.01	9.34–15.69	130
Total	6.84	0–44.22	2203	12.93	7.64–18.76	2330

Figure 1 Average values of cohort characteristics by cohort



sharp increase for cohorts 8 and 9, which were born in the First World War era. Rates then dropped to levels at or slightly below the pre-war levels, reaching their lowest average levels for cohort 17, born in 1955–9. The average rates for later cohorts were substantially higher, reaching unprecedented levels after the birth of cohort 20 (1970–4). With respect to relative cohort size, cohorts born in the late 19th century were relatively large, compared to those born later, and the average size of cohorts across the 19 countries in our sample fell relatively steadily until the end of the Great Depression and the start of the Second World War. The impact of the post-Second World War baby boom is reflected in the larger size of cohorts 15–18 (born between 1945 and 1964), while cohorts born after that point have all been much smaller. As shown in Figure 1, there is a tendency, more pronounced for recent cohorts, for the two measures to move in opposite directions.[16]

The literature on the cohort-related social capital hypothesis suggests that cohorts that have higher levels of RCS and/or NMB will have higher rates of suicide; that the effects, especially of RCS may be stronger for males than for females; and that the cohort effects may be stronger at younger than at older ages. The contextual effects hypothesis suggests that these effects should vary across the families of nations, with stronger effects in the English-speaking nations and weaker effects in the Nordic countries. The data shown above indicate substantial variation in cohort-related social capital and in age-period-specific suicide rates in our sample. As a result, our dataset is well suited to test these two hypotheses.

Testing the Model

Table 4 compares the fit indices separately for men and women for six different models of the age-period-specific suicide rates in 19 countries. The baseline

Table 4 Comparison of models for suicide rates for males and females across 19 countries from 1950 to 2000 (cohorts born between 1875–9 and 1980–4)

	Models					
	1	**2**	**3**	**4**	**5**	**6**
1. Age and period dummies with randomly varying intercepts for cohorts and countries[a]	x	x	x	x	x	x
2. LNRCS and LNNMB[b]		x	x	x	x	x
3. LNRCS and LNNMB by age interactions[b]			x	x	x	x
4. English, English × LNRCS, English × LNNMB[b]				x	x	x
5. Romanist[b]					x	x
6. Romanist × LNRCS, Romanist × LNNMB[b]						x
Men –2 × LogLikelihood	1227.2	738.0	686.7	668.5	656.9	653.0
BIC	1309.7	832.2	804.5	795.2	786.5	788.4
Women –2 × LogLikelihood	2482.5	1915.0	1849.1	1822.8	1813.5	1812.5
BIC	2565.0	2009.2	1966.9	1949.4	1943.0	1948.0

[a] Autocorrelation within cohorts is modeled using a first order autoregressive term.
[b] The effects Age and Period are fixed within countries and allowed to randomly vary across countries as are LNRCS and LNNMB and their interactions with age. All variables involving country level indicators are fixed.

model includes the age and period dummy variables as fixed effects (within countries), as well as a random variable for cohorts and countries and an autoregressive term for age-period-specific suicide rates within cohorts. When compared to this baseline model, Model 2, which adds the two measures of cohort characteristics, provides a significantly better fit for both men and women.[17] Similarly, for both sex groups, Model 3, which adds the interactions of age and cohort characteristics, is a better fit than Model 2;[18] Model 4, which adds the dummy variable for English-speaking countries and the interactions of this dummy variable with cohort characteristics, is a better fit than Model 3; and Model 5, which adds a dummy variable for Romanist countries, is a better fit than Model 4.[19] The addition of the interactions of cohort characteristics and the dummy variable for Romanist countries in Model 6 does not provide a significantly better fit (for males, chi-square change = 3.9, d.f. = 2, $p > .20$; for females, chi-square change = 1.0, d.f. = 2, $p > .50$). Further, the BIC statistic indicates that Model 5 fits better than Model 6 for both males and females.

Table 5 includes the coefficients associated with Models 1, 2, 3, and 5 for both men and women. Model 1 serves as a baseline model, while the additional models sequentially add the cohort characteristics, the interactions of the cohort characteristics with age, and the country-level coefficients and interactions. All of the models include dummy variables for the effects of age and period and allow these coefficients to vary randomly across countries. Coefficients associated with these variables are not included to simplify the presentation of the data, but are available upon request from the authors.

Table 5 Unstandardized coefficients showing the influence of country groupings on the effects of RCS and NMB on the logged age-period-specific suicide rates for men and women, 1950–2000

	Men				Women			
	Model 1	**Model 2**	**Model 3**	**Model 5**	**Model 1**	**Model 2**	**Model 3**	**Model 5**
FIXED EFFECTS								
COHORT-LEVEL COEFFICIENTS								
LNRCS		.8590***	.9349***	.6401**		.5718**	.5152***	.1546
		(.2295)	(.2113)	(.2388)		(.1858)	(.1662)	(.1461)
LNNMB		.4514***	.2618**	.0489		.3086***	.0177	–.1655**
		(.0898)	(.0909)	(.0843)		(.0906)	(.0881)	(.0795)
LNRCS × YOUNG			–.3475	–.3758*			–.0418	–.0683
			(.1943)	(.1820)			(.1794)	(.1773)
LNRCS × OLD			.3989	.3566			.2553	.1413
			(.2115)	(.2021)			(.2253)	(.1951)
LNNMB × YOUNG			.2050**	.1964**			.4000***	.3915***
			(.0655)	(.0619)			(.0894)	(.0801)
LNNMB × OLD			.0076	.0130			.2204	.2085
			(.1670)	(.1684)			(.1917)	(.1933)
COUNTRY-LEVEL COEFFICIENTS								
English				–.6597*				–.6717***
				(.2335)				(.2484)
English × LNRCS				.8856*				1.1017***
				(.4238)				(.2471)
English × LNNMB				.6276***				.5825***
				(.1314)				(.1235)
Romanist				–.9011***				–.8299**
				(.2250)				(.2393)
RANDOM COEFFICIENTS								
	Men				Women			
VARIANCE	**Model 1**	**Model 2**	**Model 3**	**Model 5**	**Model 1**	**Model 2**	**Model 3**	**Model 5**
Cohort intercepts	.0778***	.0372***	.0244***	.0246***	.0191***	.0117***	.0075***	.0071*
	(.0070)	(.0044)	(.0034)	(.0034)	(.0035)	(.0025)	(.0020)	(.0020)
Country intercept	.2860**	.2935**	.3044**	.1566**	.3133**	.2989**	.3080**	.1749**
	(.0957)	(.0977)	(.1010)	(.0531)	(.1050)	(.1003)	(.1032)	(.0601)
AR(1)	.1962***	.2876***	. 2714***	.2728***	.1088***	.0545	.0111	.0140
	(.0311)	(.0323)	(.0328)	(.0327)	(.0300)	(.0308)	(.0309)	(.0308)
LNRCS		.7347**	.5590**	.4734**		.4896**	.3037**	.0948
		(.2980)	(.2451)	(.2134)		(.2026)	(.1439)	(.0708)
LNNMB		.1057*	.0855**	.0285		.1180**	.0758**	.0229
		(.0476)	(.0410)	(.0205)		(.0490)	(.0385)	(.0189)
LNRCS × YOUNG			.3292	.2497			.1153	.1156
			(.2074)	(.1771)			(.1592)	(.1440)
LNRCS × OLD			.4255*	.3606			.3946	.2038
			(.2454)	(.2202)			(.2612)	(.1788)
LNNMB × YOUNG			.0101	.0039			.0489	.0250
			(.0195)	(.0149)			(.0380)	(.0256)

Table 5 Continued

	RANDOM COEFFICIENTS							
	Men				Women			
VARIANCE	**Model 1**	**Model 2**	**Model 3**	**Model 5**	**Model 1**	**Model 2**	**Model 3**	**Model 5**
LNNMB × OLD			.2811*	. 2911*			.3617*	.3835*
			(.1557)	(.1577)			(.2024)	(.2040)
RESIDUAL	.0545***	.0506***	.0502***	.0502***	.1105***	.0885***	.0852***	.0854***
	(.0022)	(.0023)	(.0023)	(.0023)	(.0041)	(.0034)	(.0032)	(.0032)
MODEL FIT								
–2 × LogLikelihood	1227.2	738.0	686.7	656.9	2482.5	1915.0	1849.1	1813.5
BIC	1309.7	832.2	804.5	786.5	2565.0	2009.2	1966.9	1943.0

Note: Numbers in parentheses are standard errors. The significance tests are one-tailed for each of the predicted hypotheses: LNRCS, LNNMB, English, English × LNRCS, English × LNNMB, and Romanist. The tests are two-tailed for the other fixed effects and for all of the variance component tests: * $p < .05$, ** $p < .01$, *** $p < .001$.

The first part of Table 5 gives the fixed effects coefficients associated with the measures of cohort-related social capital, the interaction effects, and the country-level coefficients associated with the four models for each sex group. Results regarding the cohort-related social capital hypothesis may be seen by examining the coefficients for LNRCS and LNNMB and their interactions with age across the four displayed models (the cohort level coefficients). Results regarding the contextual hypothesis may be seen by examining the country-level coefficients. The negative coefficients associated with the dummy variables for English-speaking and Romanist countries reflect the fact that these two sets of countries have much lower suicide rates than the Germanic and Nordic nations. The positive interactions with LNNMB and LNRCS show the increased salience of the measures of cohort-related social capital in the English family of nations.

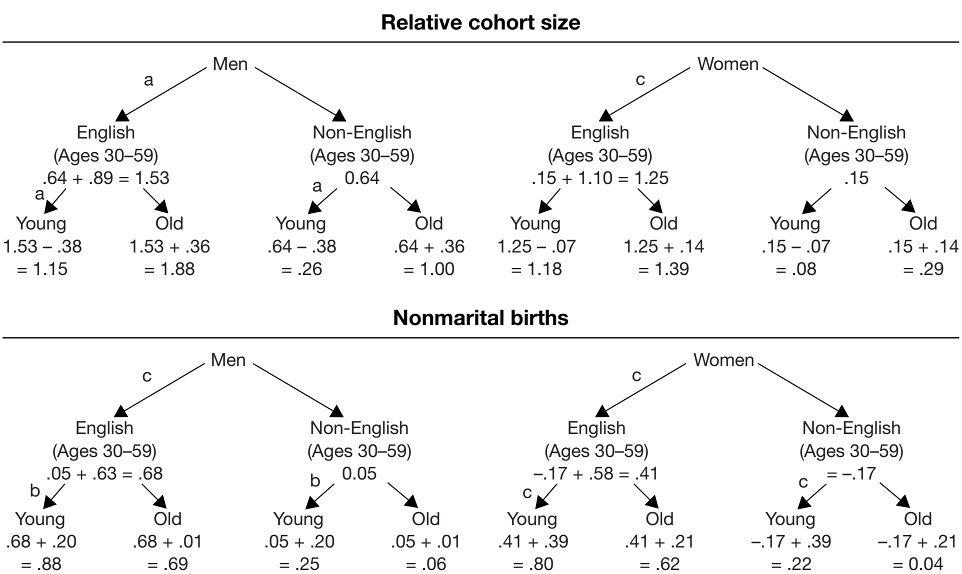
For both sex groups, as hypothesized, the coefficients associated with the two measures of cohort-related social capital are positive and significant in Model 2 (e.g. the coefficient associated with LNRCS is .8590, $p < .001$ for men; and .5718, $p < .01$ for women). With the addition of the interaction effects in Model 3, we see, by comparing values across the various interactions, that the effects of RCS and NMB vary between age groups. With the addition of the country level variables (Model 5), the picture becomes slightly more complex, indicating that, as the contextual hypothesis suggests, the influence of cohort-related capital can be muted or enhanced within cultural contexts.

For men, the influence of relative cohort size on suicide rates is positive and highly significant across all the countries within the analysis (b = .64, $p < .01$), but significantly higher within the English-speaking countries (interaction b = .89, $p < .05$). In direct contrast to the hypothesis that the effect of relative cohort size would be smaller for older age groups than younger age groups, the effect of RCS is substantially higher for older groups (b = .36, Model 5) and significantly smaller for the younger age group (b = –.38, $p < .05$, Model 5). For women, the influence of relative cohort size is nonsignificant for the non-English-speaking countries, once country-level variables are added to the model (b = 0.15, Model 5). However, the influence of relative cohort size is strong and significant for women within the English-speaking countries (interaction b = 1.10, $p < .001$), with a magnitude close to that for men. Neither of the interactions of age with relative cohort size is significant for women, although the pattern of a stronger effect for older age groups also appears.

For men, the results in the final model indicate that, as with relative cohort size (and as hypothesized), the effect of nonmarital births is significantly larger in the English-speaking than in the non-English-speaking countries; and, as hypothesized, the effect is significantly stronger among the younger age groups. Similarly, for women, the influence of nonmarital births is significant only for the English-speaking countries, but the effect is significantly stronger for the younger age groups.

Given the significant interactions with age and families of nations, we turn our attention to Table 6. This table summarizes these patterns by showing the effects of logged NMB and RCS on logged suicide rates estimated in Model 5 for both sex groups across the three age groups (young, middle, and old) and the English-speaking and non-English speaking countries.[20] Recall that, because both the independent and dependent variables are logged, the coefficients indicate percent change in the dependent variable (suicide rates) that would be associated with a one percent change in the independent variable when the

Table 6 Summary of model effects of logged suicide rates on LNRCS and LNNMB by age, country, and sex for Model 5



Significantly different from all other ages: [a] $p < .05$; [b] $p < .01$; [c] $p < .001$.

other cohort characteristic, age, and period are controlled. For instance, the value of 1.88 associated with RCS for older men in the English-speaking countries (the highest in Table 6) indicates that a one percent shift in relative cohort size is associated with a 1.88 percent shift in the age-period-specific rate of suicide for men in this age group and family of nations, controlling for age, period and nonmarital births.

The data in Table 6 provide mixed support for the hypothesis that the effects of cohort-related social capital are stronger for men than for women. The effects of RCS are stronger for men than for women for all age groups in the non-English-speaking countries and for all but the youngest age group in the English-speaking countries. With nonmarital births, the effects are approximately equal in each age group and country group, except possibly (although not significantly so) for middle-aged men in the English-speaking countries, where the effect is stronger for them than for women (.68 versus .41).

Comparing the relative influence of our two measures of cohort-related social capital, the influence of relative cohort size is stronger than that of nonmarital births for men in all age and country groups, with the exception of young men in non-English-speaking countries, where the influence is approximately equal. For women, the influence of relative cohort size is stronger than that of nonmarital births for all ages in the English-speaking countries, but there is no consistent pattern across the age groups in the non-English-speaking countries. The variation in nonmarital births, however, has exceeded that for relative cohort size in all of the groups of countries over the time span covered in our research.

The second part of Table 5 includes the random coefficient estimates of variance. Comparing values across models allows us to estimate the extent to which our models account for differences in suicide rates between cohorts (see line 1 for cohort intercepts) and countries (line 2 for country intercepts). The values for Model 1 indicate that there are significant differences between cohorts (.0778, $p < .001$ for men; and .0191, $p < .001$ for women) and countries (.2860, $p < .01$ for men; and .3133, $p < .01$ for women) in their suicide rates when only age and period dummy variables are included. The drop in these values across the models indicates the extent to which variables included in the models account for the variance between cohorts and countries. For men, the cohort and country-level variables included in Model 5 account for 68 percent of the variance between cohorts and 45 percent of the variance between countries. For women, the variables in Model 5 account for 63 percent of the variance between cohorts and 44 percent of the variance between countries.[21] Not surprisingly, the cohort characteristics and their interactions with age (Models 2 and 3) play the major role in explaining the variance of cohorts, while the country-level variables introduced in Model 5 explain variance between countries.

The autoregressive coefficient indicates the extent to which the residuals within cohorts are correlated. This coefficient remains significant across all models for men, but is significant for women only in Model 1. The random coefficient variance estimates associated with the LNRCS and LNNMB estimate the differences in the effect (slope) of these independent variables across the countries. These estimates decline markedly from Model 2 (the one that first includes cohort characteristics) to the final model (36% for LNRCS for males, 73% for LNNMB for males, and 81% for both LNRCS and LNNMB for females) and are statistically insignificant for males for LNNMB and for females for both cohort characteristics. Of the estimates associated with the interactions of the cohort characteristics and age groups, only that for the interaction of LNNMB and older age groups is significant in the final model (.2911, $p < .05$ for men; .3835, $p < .05$ for women). Finally, the residual variance declines slightly from Model 1 to Model 5 (8% for men and 23% for women), but remains significant for all models.

DISCUSSION

The analysis in this article provides an extensive test of the cohort-related social capital hypothesis by including a very broad range of cohorts (born from 1875 to 1984) from 19 separate countries. The results provide support for both the cohort-related social capital hypothesis and the contextual hypothesis. We find strong effects of both relative cohort size and nonmarital births on age-specific rates of suicide, but only in the English-speaking countries are both of these effects statistically significant. Because, as have other studies, our model includes strong controls for both age and period, these results provide strong evidence that these are indeed cohort effects and not those related to uncontrolled influences of age and period.

The broad range of cohorts in our sample provides a large number of cases for both young and old age groups. This inclusion of more cases at older age groups provides improved tests of hypotheses regarding the interaction of cohort characteristics with age. Our results directly contradict the expectation that the effect of relative cohort size would be stronger for those in younger age groups (Kahn and Mason, 1987; Steffensmeier et al., 1992) and perhaps even negative for older ages (McCall and Land, 1994; Pampel, 2001; Pampel and Peters, 1995).[22] In fact, the positive relationship of relative cohort size with suicide rates is weaker for younger age groups, for both men and women, although this difference is statistically significant only for men. Our results with nonmarital births provide support for the views of clinicians that adolescents may be more vulnerable to adversities than adults (Diekstra, 1995). For both males and females, the influence of nonmarital births is stronger for the younger age groups.

The inclusion of more cases at older age groups also provides a better test of the 'lasting effects' principle of cohort theory. Our results appear to provide substantial support for this principle. Our measures of cohort characteristics are directly related to the youthful experiences of the cohorts in our analysis. Our results suggest that these characteristics have a significant influence on suicide rates throughout all ages of the life cycle and periods and that in some cases the influence is even more pronounced at older ages.[23]

Our results support the hypothesis that cohort effects are stronger for males than for females with respect to relative cohort size, but not nonmarital births. This finding supports Pampel's (2001) suggestion that males would be more negatively affected by relative cohort size because of their greater involvement with the labor market.

In contrast to results with a more limited range of cohorts, the present results also indicate that larger relative cohort sizes produce more of a risk than nonmarital births. In fact, some analyses with a more limited range of cohorts have found positive, but statistically insignificant, effects of relative cohort size.[24] Our analysis, however, has a broader range of values of relative cohort size (especially larger size cohorts) and a larger sample size.[25] As shown in Figure 1, the largest cohorts tend to be those born in the late 19th and earlier 20th centuries and are, consequently, overrepresented in the older age groups in our analysis.

Concern over the relative importance of relative cohort size and nonmarital births should not overshadow important common results. Both relative cohort size and the proportion of nonmarital births are indicators of cohort-related social capital – the resources, support, opportunities, constraints, and social networks – that are available to birth cohorts. Reductions in cohort-related social capital can arise from different sources in different historical eras. For instance, birth cohorts of the late 19th century as well as the post-Second World War era have been disadvantaged by their relative size. Cohorts born in the 1970s and 1980s, as well as to some extent those born during the First World War and the Depression era, at least in some countries, have been disadvantaged by their nontraditional family structures. Whatever the source of the loss of cohort-related social capital, it seems to have lifelong effects in terms of a cohort's propensity to suicide. From a policy perspective, knowing which cohorts are at risk and providing ways to minimize that risk is important, no matter which source may have prompted that risk.

The results of our examination of the 'contextual effects' hypothesis help point toward effective policy interventions. Earlier examinations of contextual effects were potentially limited by using quantitative measures obtained from the last half of the 20th century. To counter this limitation we used a more qualitative indicator of national culture. As hypothesized, we found that cohorts within the English-speaking nations, with their historical tradition of providing relatively limited public financial support for families and children, were particularly susceptible to the negative effects of low levels of cohort-related social capital. Cohorts in the Nordic countries, with a long tradition of cultural and political tolerance toward nontraditional family forms, were not significantly less likely than those in Romanist or Germanic countries to experience the effects of low levels of cohort-related social capital. Instead, the effects of cohort-related social capital were remarkably similar across all of the non-English-speaking countries, even though they vary rather strongly in their historical support for children born to nonmarital unions. These results may suggest that the most important contextual effect on the impact of cohort-related social capital is not the legal context or views of a society regarding nonmarital births (the factor that distinguishes the Nordic nations), but instead the traditions of political and financial support for families and children, the factor that distinguishes the English-speaking nations from others in our sample.

It is also important to note that the variables included in our models cannot account for all of the variation between cohorts or countries. While the random coefficients variance associated with cohort and country intercepts are markedly reduced in our final model, significant variation remains. This indicates that neither the cultural differences between the 'families of nations' in our analysis, nor the measures of cohort-related social capital, are sufficient to account for all of the systematic variation in the post-Second World War age-period-specific suicide rates. Further research could attempt to identify additional explanations of cohort variation. An anonymous reviewer of this article noted that a potentially fruitful area of inquiry could involve the role of regional economic factors that affect contiguous nations (e.g. Switzerland and France or Austria and Italy) as well as economic connections between countries that are not geographically based (e.g. France and Canada, especially Quebec).[26]

Finally, we note the importance of the method of analysis that we use. As we mentioned at the beginning of our article, distinguishing age, period, and cohort effects remains a central concern of demographers. We used a new approach to such analyses by treating the cohort observations as a level-two variable in a hierarchical linear model. This allows us to estimate the variation between cohorts after controlling for age and period effects (O'Brien et al., 2003).[27] When we introduce two cohort characteristics to the model, we can assess how much of this cohort variation is associated with those characteristics. This approach, we believe, provides the best technique yet developed for estimating and assessing age, period, and cohort effects.

NOTES

1 O'Brien and Stockard (forthcoming) demonstrate that age-specific rates of suicide and homicide vary together and that these parallel movements are strongly influenced by common indicators of cohort-related social capital.

2 These effects are structural in nature and not based on how much single parents or those with larger families love their children or want them to do well. It is simply more difficult to provide needed attention and supervision when there is only one adult in the house or when there are more children. These effects are heightened by the higher levels of poverty found in single parent households or the smaller amount of money per child in larger families or cohorts.

3 Specifically, O'Brien and Stockard (2003), in an analysis of homicide offending, found that the effect of nonmarital births was significantly stronger for younger age groups (those 15–19 and 20–4), and the effect for those in older age groups (40–5 and 45–9) was significantly weaker. None of the interaction effects with RCS was significant.

4 Esping-Anderson's (1990) categorization of social welfare states (liberal, corporatist-statist, and social democratic) parallels the distinction we use, but is less specific in that the Germanic and Romanist countries are both included within the corporatist-statist model.

5 Japan was omitted because data were not available on nonmarital births for a large proportion of the cohorts. Germany and the Eastern European countries were omitted because of concerns regarding the reliability and the availability of data.

6 For the earliest year the data for suicides from Denmark, Italy, Norway, Spain, Sweden, and Switzerland are for 1951, from Finland for 1952, and for Austria, Belgium, and Portugal for 1955, and from Greece for 1956. For the other countries the earliest year of data is 1950. The last year of data in the analysis is 2000 for all countries except for Denmark, France, Greece, and the United Kingdom, for which it is 1999, and for Belgium, for which it is 1997.

7 Details on computations for each country are available upon request from the authors.

8 A measure based on the percentage of the population ages 5–54 that was in the cohort when the cohort was ages 5–9 was used for cohort 10 (born in 1920–4) for Australia, Norway, and the United Kingdom and for cohort 11 (born in 1925–9) for Italy. In addition, because of a lack of other data, a measure based on an average for the ages of 0–4 (the percentage of the population age 0–49 that was 0–4 when the cohort was age 0–4) and 20–4 (the percentage of the population age 20–69 that was 20–24 when the cohort was age 20–24) was used for cohort 11 for the United Kingdom.

9 The years in which census data were gathered varied slightly from one country to another. Because, however, birth rates change relatively slowly, we have no reason to suspect that these variations substantively affect our results.

10 Data were not available on nonmarital births for Spain for cohort 12 (born in 1930–4), or for Greece for cohorts 14–15 (born in 1940–9). Thus, these cohorts are missing from the analysis for these countries.

11 That is, we expect cohort effects to proportionately affect age-period-specific suicide rates.

12 Of the 2412 observations for each sex group there were four cases of no suicides recorded for an age group for males and 36 such cases for females. Most of the cases of no recorded suicides involved Ireland. It is important to emphasize that these zeros represent observed rates of zero and not rates of suicide with so few cases that the rates were considered 'too unreliable to report.' Thus, keeping these cases in our analysis is important.

13 Before combining these groups, we tested for differential effects of LNRCS and LNNMB between these two groups of nations and for differences in the intercepts of these groups of nations in our most fully specified model in the analysis conducted later in this article. This was Model 5 in Table 5 with the country variable changed to Germanic and the interactions changed to Germanic × LNRCS and Germanic × LNNMB. When this analysis was conducted on only the Germanic and Nordic group of countries the reference group was the Nordic group of countries. We found no statistically significant differences in these nation groups on these variables.

14 A major exception is Greece, where in several years, at younger ages, the rates for females exceed those for males.

15 The increase from cohort 1 to 2 primarily occurs because we had no data for Austria for cohort 1, and Austria has an extraordinarily high rate of nonmarital births.

16 For the entire sample, the correlation between RCS and NMB is –.25. For cohorts 1–7, correlation is .02, for cohorts 8–14 it is –.18, and for cohorts 15–22 it is –.42.

17 We entered the LNRCS and LNNMB terms at the same time because both are theorized to affect the level of suicide in cohorts and to omit one would misspecify the relationship. Savolainen (2000) stresses this point when he controls on a family structure variable and finds that the effect of RCS then emerges.

18 We entered the age by LNRCS and age by LNNMB interactions separately and found a significant improvement when both sets of interactions were in the equation. We report the results from Model 3 that contain both sets of interactions.

19 For males, comparing Model 2 with Model 1, chi-square change = 489.2, d.f. = 2, $p < .001$; comparing Model 3 with Model 2, chi-square change = 51.3, d.f. = 4, $p < .001$; comparing Model 4 with Model 3, chi-square change = 18.2, d.f. = 3, $p < .001$; comparing Model 5 with Model 4, chi-square change = 11.6, d.f. = 1, $p < .001$. For females, comparing Model 2 with Model 1, chi-square change = 567.5, d.f. = 2, $p < .001$; comparing Model 3 with Model 2, chi-square change = 65.9, d.f. = 4, $p < .001$; comparing Model 4 with Model 3, chi-square change = 26.3, d.f. = 3, $p < .001$; comparing Model 5 with Model 4, chi-square change = 9.3, d.f. = 1, $p < .01$.

20 These values were calculated by simply adding the coefficients associated with a particular group in Model 5, the final model. For instance, the value of 1.00 for RCS for 60–74-year-old men (the older group) in the non-English-speaking countries represents the sum of the main effect for RCS (.64), and the interaction of RCS and the older age group (.36).

21 These values are calculated, in the PRE format, by subtracting the random coefficient of variance in Model 5 from Model 1, dividing by the coefficient for Model 1, and multiplying by 100 to attain a percent. For example, for cohort intercepts for men, $(.0778 - .0246)/.0778) \times 100 = 68.4$.

22 Note, however, that these studies operationalize the size of old cohorts differently than we do.

23 In all cases the overall effects of LNRCS and LNNMB are positive for the oldest age group and youngest age group. Only in the case of LNNMB for women in the middle age group in the non-English-speaking nations is the effect of a cohort characteristic negative.

24 Specifically, using cohorts born in 1915 and later, relative cohort size has been found to have an insignificant influence on suicide rates for nonwhites in the United States (Stockard and O'Brien, 2002b); on suicide rates for women in a sample of 14 industrialized countries, all of which are in the present sample (Stockard and O'Brien, 2002a); and on homicide rates of whites in the United States (O'Brien and Stockard, 2002).

25 For instance, Stockard and O'Brien's sample of cohorts from 14 nations, including only those born in 1915 or later, had relative cohort sizes that ranged from 7.6 to 16.1. As shown in Table 3, relative cohort size ranges from 7.6 to 18.8 in our sample.

26 The control for period in our models controls for factors such as upturns and downturns in the nations' economies, but only to the extent that those upturns and downturns are constant across countries. For this control to be problematic for our hypotheses it would be necessary for there to be a relationship between such upturns

and downturns in the economies of pairs or triplets or even larger combinations of countries that are related to both our cohort characteristics and to age-period-specific rates of suicide.

27 The nesting of these effects within families of nations and examining differences in the slopes and intercepts associated with these families is a more typical use of hierarchical linear models.

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