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# A changing epidemiology of suicide? The influence of birth cohorts on suicide rates in the United States



Julie A. Phillips

Department of Sociology/Institute for Health, Health Care Policy and Aging Research (IHHPAR), Rutgers, The State University of New Jersey,  
26 Nichol Avenue, New Brunswick, NJ 08901-2882, USA

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## ABSTRACT

The increases in suicide among middle-aged baby boomers (born between 1946 and 1964) in the United States since 1999 suggest a changing epidemiology of suicide. Using data from 1935 to 2010, this paper conducts age-period-cohort analyses to determine the impact of cohorts in shaping temporal patterns of suicide in the United States. The analysis demonstrates that age, period and cohort effects are all important in determining suicide trends. Net of age and period effects, the cohort pattern of suicide rates is U-shaped, with cohorts born between 1915 and 1945 possessing among the very lowest suicide rates. Suicide rates begin to rise with boomers and subsequent cohorts exhibit increasingly higher rates of suicide. The general pattern exists for both men and women but is especially pronounced among males. The average suicide rate over the entire period for males is about 28 per 100,000, 95% CI [27.4, 28.7]. For males born in 1930–34, the suicide rate is estimated to be 17.4 per 100,000, 95% CI [15.9, 18.8]; for males born between 1955 and 1959, the rate is essentially the same as the average for the period while for males born between 1985 and 1989, the suicide rate is estimated to be 37.8 per 100,000, 95% CI [33.1, 43.4]. The results dispute popular claims that boomers exhibit an elevated suicide rate relative to other generations, but boomers do appear to have ushered in new cohort patterns of suicide rates over the life course. These patterns are interpreted within a Durkheimian framework that suggests weakened forms of social integration and regulation among postwar cohorts may be producing increased suicide rates.

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## 1. Introduction

Recent reports document a sharp rise in suicide rates among the U.S. middle-aged population beginning in 1999 (Hu et al., 2008; Phillips et al., 2010). For those aged 45–54, the suicide rate increased from 13.9 per 100,000 in 1999 to 19.6 per 100,000 in 2010. Rates also rose substantially for those aged 55–64, but there has been no concomitant increase in rates for other age groups (Centers for Disease Control (CDC)). This pattern has taken many by surprise as historically, the overall picture of suicide rates among middle-aged persons has been one either of stability or decline. Prior to this recent increase, rates among males aged 45–64 had declined by more than half, from approximately 60 per 100,000 in 1930 to less than 30 per 100,000 by 1986. Rates for females aged 45–64 had shown more cyclical fluctuation, but they also decreased from about 13 per 100,000 in 1930 to about 8 per 100,000 by 1986 (McIntosh, 1991).

Among the explanations offered for this recent trend among the middle-aged is one that attributes the increases to cohort effects. Cohort effects consider both the contemporaneous and the historical socio-cultural context that people experience, recognizing that certain events have different implications for suicide according to the age at which that event occurs. Some have speculated that the baby boomer birth cohort, those born between 1946 and 1964 and who occupied the middle age ranges between 1999 and 2010, may have a unique suicide risk that they carry with them through the life course. However, the challenge in explaining temporal patterns in suicide rates is to distinguish the effects of cohort membership from those of age or time period (Yang, 2007). Suicide rates vary with age due to biological or behavioral differences and accumulation of social experiences associated with age. Thus, shifts in the population's age composition or in the age pattern of suicide can produce changes in overall suicide rates. Furthermore, historical or social events that occur in a given time period, such as new treatments for depression, can affect suicide rates across all age groups, altering the overall suicide rate by either increasing or decreasing deaths across all age ranges.

E-mail address: [jphillips@sociology.rutgers.edu](mailto:jphillips@sociology.rutgers.edu).

Researchers have employed various techniques to disentangle these effects, often finding support for the presence of cohort effects in explaining changes over time in suicide in both the U.S. (Ahlburg and Shapiro, 1984; Manton et al., 1987; Stockard and O'Brien, 2002; Woodbury et al., 1988) and cross-national context (Ajdacic-Gross et al., 2006; Granizo et al., 1996; Gunnell et al., 2003; Snowden and Hunt, 2002). In the United States, attention has focused primarily on the influence of the baby boomer cohort. When attempting to explain the dramatic increase in adolescent suicide rates during the 1960s/1970s, scholars pointed to possible cohort effects that might account for the patterns, noting the ways that membership in a large birth cohort might confer disadvantage (Easterlin, 1980; Macunovich, 2002) and/or how the experience of growing up during the post-war period of economic prosperity and rapidly improving health and life expectancy prospects may have left this cohort with poor coping skills (Sudak et al., 1984; McIntosh, 1994). However, these U.S. analyses are generally limited in scope and dated. While they capture some of the decline in suicide rates that began in the 1980s, they do not consider the divergent trends by age group observed since 1999 and the marked increase in suicide rates among boomers while in middle age. Furthermore, prior work has not taken advantage of recent methodological developments in APC analysis. The *Annual Review of Sociology* (Wray et al., 2011) calls explicitly for research that capitalizes on these newly-developed techniques and links micro and macro features of life to capture the ways in which individuals are embedded within social structures such as birth cohorts and time periods.

Hence, this paper revisits the question of age, period and cohort in shaping U.S. suicide trends. As shown in Fig. 1, crude suicide rates have waxed and waned over time, with declines and relative stability between 1933 and the early 1950s, increases through to the early 1980s, and then subsequent drops through to about 2000 before rates started to creep upward again. Accompanying these overall trends is an important shift in the epidemiology of suicide – beginning in 2004, midlife replaced old age as the life course stage exhibiting the highest rates of suicide. This analysis seeks to determine the extent to which cohort effects can explain these patterns by compiling seventy-five years of data on U.S. suicide for twenty-seven birth cohorts and applying new developments in APC analysis, namely the Intrinsic Estimator (IE).

## 2. Data and methods

Data on suicide deaths by five year age group and sex are obtained from the U.S. Vital Registration System for the period 1935–2010; those who did not reside in the United States at the time of death are excluded. The corresponding information on population counts by five year age group and sex come from the Census Bureau. Using these two pieces of information, age- and sex-specific suicide rates are computed for every five year period between 1935 and 2010 for the population aged 15–74. That is, for A age groups (15–19, 20–24 ..., 70–74) and P time periods (1935–39, 1940–44, ..., 2005–2010) of equal length (five years), age–sex-specific suicide rates are computed. Over this time period, a total of 27 consecutive birth cohorts are represented, beginning with those born in 1860–64 and aged 70–74 in 1935 and ending with those born in 1990–95 and aged 15–19 in 2010. Those older than 75 are not included since information by five-year age intervals is not available for the earlier years of the study period.

The basic APC model is of a log linear regression form as follows (Mason et al., 1973):

$$\log(r_{APC}) = \log(s_{APC}/n_{APC}) = \alpha + \beta_A + \gamma_P + \delta_C,$$

where  $r_{APC}$  represents the expected suicide rate in age–period–cohort group (A, P, C);  $s_{APC}$  denotes the expected number of suicide deaths;  $n_{APC}$  is the population at risk;  $\alpha$  is the intercept or adjusted mean suicide rate;  $\beta_A$  is the effect for age groups  $A = 1, \dots, a$ ,  $\gamma_P$  is the effect for time periods  $P = 1, \dots, p$ ; and  $\delta_C$  is the effect for cohorts  $C = 1, \dots, c$ . This model cannot be estimated with conventional regression techniques because of an identification problem. Any two factors (e.g. age and birth cohort) enable us to predict exactly the third factor (here, time period) (Mason et al., 1973; Binstock and George, 2011). In other words, one can determine the time period with knowledge of a person's age and year of birth, and thus it is not possible to obtain a unique solution to the model. A substantial body of literature discusses this phenomenon and approaches to circumvent the problem, although each of these “solutions” has its own set of limitations (Mason et al., 1973; Smith, 2008).

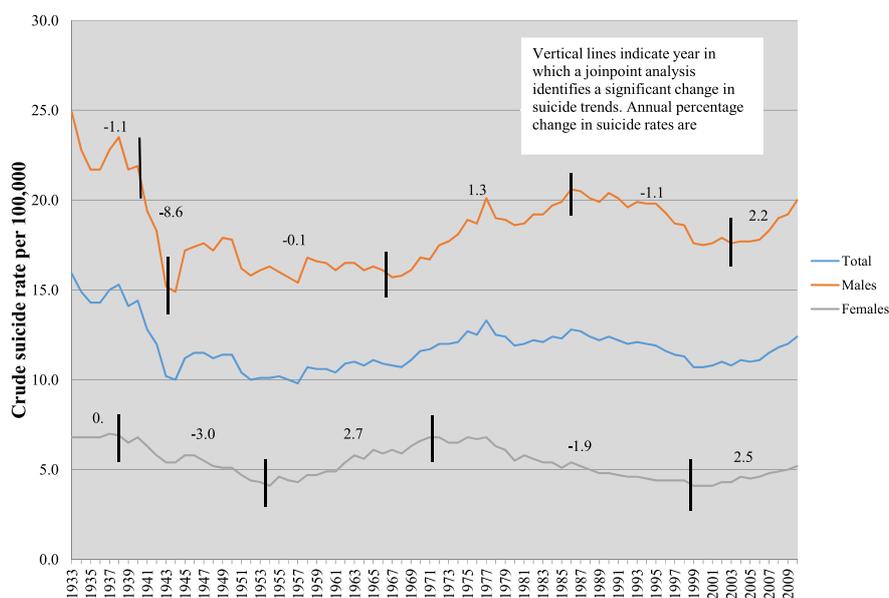


Fig. 1. U.S. crude suicide rates, overall and by sex, 1933–2010.

Source: U.S. National Center for Health Statistics (<http://mypage.iu.edu/~jmcintos/SuicideStats.html>)

Among the most common is an approach that constrains one or more coefficients in the conventional linear regression model to be equal (e.g. the coefficient for age group 15–19 equals that for age group 20–24), so as to obtain a solution. The limitation to this approach is that one must rely on potentially faulty assumptions when selecting the constraint, and results may vary wildly depending on the constraint adopted. Another method, commonly adopted by sociologists, is the “proxy variables approach”, which substitutes one of the set of age, period or cohort dummy variables with one or more alternate variables. Thus, if one believes that unemployment levels are the primary period effect of interest, the time period dummy variables would be replaced with unemployment rates for each period. The problem here is that the selected substitute variable(s) may not capture all relevant variation in that particular component.

Yang et al. (2004, 2008) recently resurrected the Intrinsic Estimator (IE) solution to APC analysis and argue that this method has several desirable statistical properties compared to prior approaches. Although the IE model applies a constraint to obtain a solution, it does so by assuming no a priori knowledge of the phenomenon being studied and uses information that is completely independent of the event rate. This model attribute may be desirable since seemingly reasonable assumptions about constraints to place on age, period or cohort effects have been shown previously to lead to counter-intuitive estimates. Furthermore, the variances or error terms of the estimated coefficients obtained from the IE approach, which are assumed to follow a Poisson distribution, are smaller than those obtained from other constrained solutions. In this sense, the IE coefficients are an average of sorts of the constrained estimates and can serve as the representative solution (O'Brien, 2011; Smith, 2004). Yang applied this new approach to the study of adult chronic disease mortality from heart disease and certain forms of cancer, finding substantial reductions in mortality between the late 1960s and late 1990s that are largely accounted for by cohort effects (Yang, 2008).

In this analysis, I use three approaches to estimate the APC model. I focus attention on the results obtained with the Intrinsic Estimator (IE) model given the advantages listed above. However, in light of concerns raised by O'Brien (2011) about the qualities of the IE approach – namely, that the IE constraint may yield biased results – I estimate two additional constrained models to assess how robust findings are to different assumptions. First, I estimate a conventional APC model using Generalized Least Squares, in which the coefficients for two adjacent time periods (1935 and 1940) are constrained to be equal (hereafter referred to as the CGLIM model). This particular constraint is chosen for the following substantive reasons: (1) the results reveal that period effects appear to be the least consequential of the three factors (age, cohort, period) in affecting suicide rates over time; (2) economic conditions – a period characteristic believed to be strongly associated with suicide rates (Luo et al., 2011) – were comparably poor in the two periods (the unemployment rate was 17% in 1935 and 14.6% in 1940); and (3) suicide rates were similar in both periods (although some caution against using the outcome to select a constraint). Second, I estimate a CGLIM with a Zero Linear Trend (ZLT) constraint following a recommendation by O'Brien (2011). The underlying assumption behind the ZLT model is that the long-run period effects on suicide have a zero linear trend – that is, there are periods of increase and decrease in suicide rates, with no particular trend over the long run. This appears a reasonable assumption in this instance (see Fig. 1). Similar results obtained across these three models should enhance confidence in the accuracy of the estimated age, period, and cohort effects. The above-described models are estimated in Stata using the `apc_cglim.ado` and `apc_ie.ado` files (see <http://www.unc.edu/~yangy819/research.html> for more detail).

Following the sequence of analyses recommended by Yang (2004, 2008), I first present a descriptive analysis of age-specific suicide rates by sex for selected time periods and birth cohorts. This analysis provides a contextual account of suicide patterns over time. To obtain a quantitative evaluation of the sources of suicide change, I fit a sequence of one and two factor models to the data. A one-factor or gross-effect model is estimated for age effects (A), period effects (P) and cohort effects (C) only. Three two-factor models are estimated to determine the effect of age and period (AP), age and cohort effects (AC), and period and cohort effects (PC) only. Finally, a full APC model, which controls for age, period and cohort simultaneously, is estimated. Two model selection criteria, Akaike's information criterion (AIC) and the Bayesian information criterion (BIC), are used to assess the fit of each model.

### 3. Results

#### 3.1. Descriptive analysis

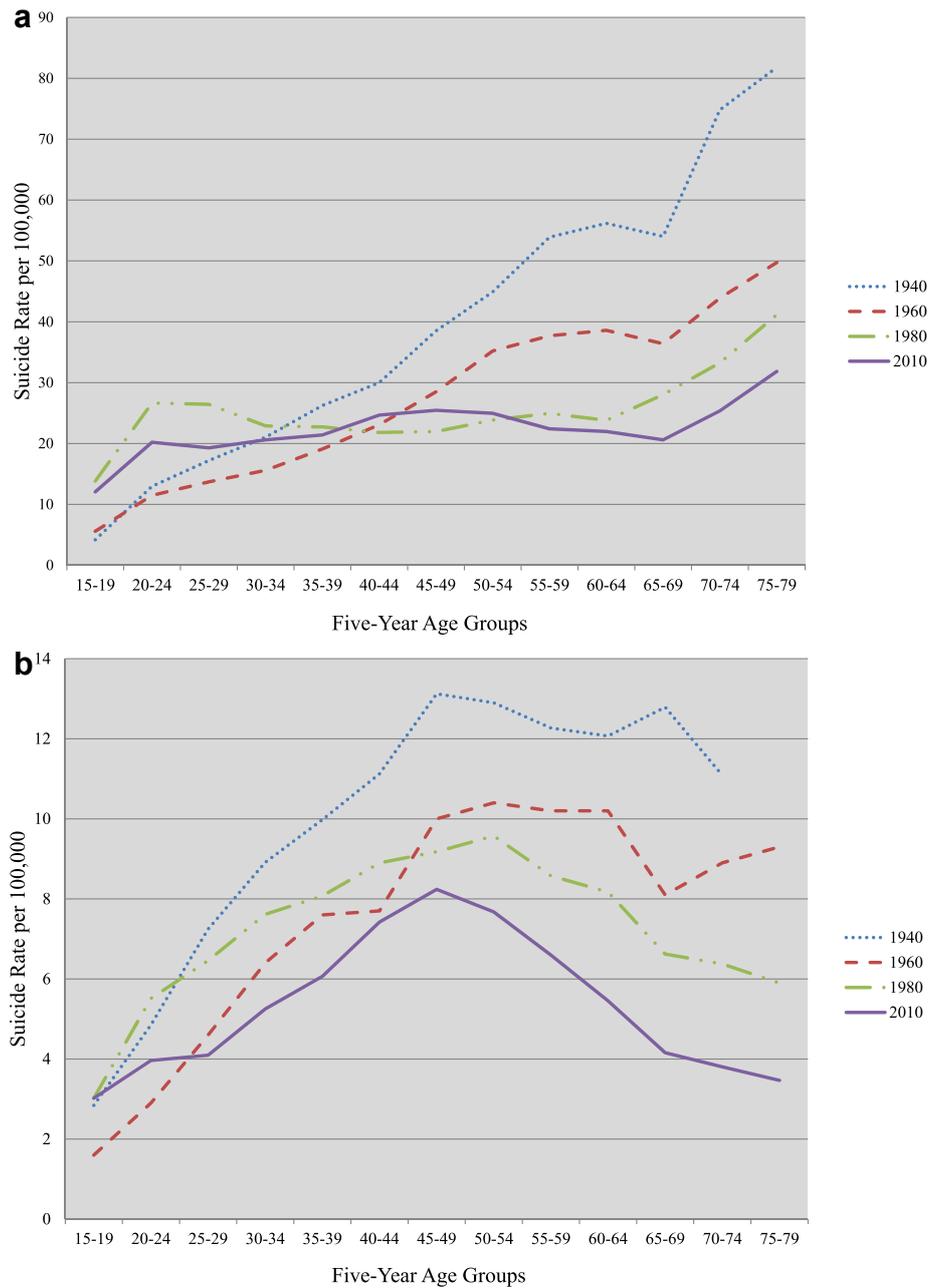
*Age-specific suicide rates, by time period.* Fig. 2a and b displays the age pattern of deaths from suicide for selected years between 1940 and 2010 for males and females respectively. Among males, during the earlier portion of this period (1940–1970), there appears to be a steady and fairly linear increase in the rate of suicide with age. For example, in 1940, the rate of suicide among 15–19 year olds was about 5 per 100,000; the suicide rate for those aged 75–79 was over fifteen times that rate, at 80 per 100,000. In the period since 1980, the age pattern has been such that rates increase dramatically in adolescence, achieve stability in the middle age ranges, and then rise steeply again in the older age ranges (65+).

As a result of these patterns, we find that since 1970, male suicide rates have been higher among adolescents and young adults relative to earlier time periods. In contrast, rates for the elderly have dropped substantially in more recent time periods. We also observe an increase in suicide rates among the middle-aged in 2010, although these contemporary rates for the middle-aged do not approach the historical highs of the pre-1970 period. The age curves of suicide are not parallel across time periods, suggesting that cohort effects are present.

The age pattern of suicide is distinct for females, who exhibit far lower suicide rates than do males at all age ranges. Historically, female suicide rates have risen sharply through young and middle adulthood and then stabilized or slightly declined during the older age ranges. More recently (post-1970), the overall shape of the curve has a more pronounced inverted U-shape, with rates for females peaking in the middle age ranges and then declining in the older age ranges (with the exception of 1990). While not completely uniform, female suicide rates within each age range exhibit a general pattern of decline over the 70-year period.

*Age-specific suicide rates, by birth cohort.* Fig. 3a and b shows the age-specific suicide rates for selected male and female birth cohorts, respectively. The five year birth cohorts are mapped to popular conceptions of social generations so as to facilitate interpretation of the patterns with a less complicated figure. The graphs of each five year birth cohort plotted separately are available upon request.

The age pattern for males is such that among the earliest cohorts (the GI generation), suicide rates tend to rise linearly with age. For the silent generation (cohorts born between 1925 and 1945) and early boomers (cohorts born between 1946 and 1954), suicide rates increase sharply during adolescence and then plateau in the middle age ranges before rising again in the older age ranges (in the case of the silent generation). However, the late boomers (born between 1955 and 1964) show a sharp increase in their suicide rate at ages 35–54 years. As the boomer cohorts have yet to reach the oldest



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**Fig. 2.** a). Age-specific male suicide rates, 1940–2010. b) Age-specific female suicide rates, 1940–2010.

age ranges, we do not know how their rates will change as they become elderly. Given this pattern of age-specific suicide rates by cohort, the largest between-cohort differences are in the age ranges of 20–29, with the GI and silent generations (1900–45 birth cohorts) registering the lowest rates, the baby boomers (1946–1964 birth cohorts) among the highest rates, and the more recent birth cohorts (post 1970) in the middle. In the middle age ranges (ages 40–50), suicide rates are remarkably similar across all birth cohorts. Note that the late boomers have among the highest rates in this age range as well as in adolescence. In contrast, early boomers, for whom data are available through age 64, do not exhibit rates that are unusually high relative to other cohorts in midlife.

Among women, there are larger between-cohort differences in age-specific suicide rates in the middle age ranges than in adolescence, where rates appear to be quite similar across birth cohorts. Suicide rates for those aged 20–29 are highest for the early boomer generation, a similar pattern to what we observed in males. The GI generation experienced their peak suicide rate at ages 50–54, and the silent generation at ages 40–44, with steady declines in rates thereafter. The earliest boomer women appear to reach their high suicide rate at ages 55–59, but this peak is quite a bit lower than prior cohorts. However, by 2010, only those boomers born before 1950 had reached ages 60–64; when 2015 data become available, we will know for certain whether the early boomers' rates peaked

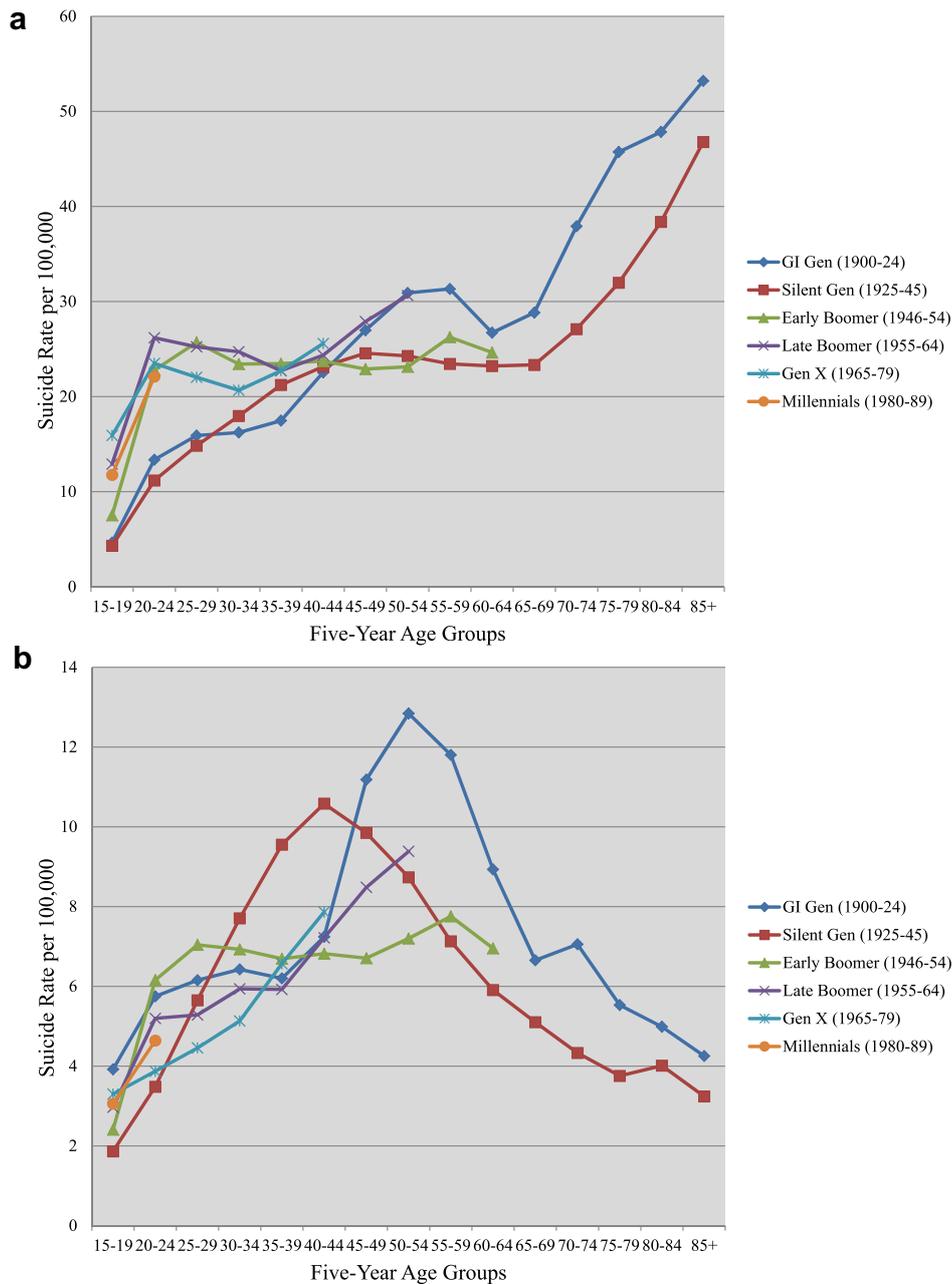


Fig. 3. a) Age-specific suicide rates for males, by cohort. b) Age-specific suicide rates for females, by cohort.

at ages 55–59 or later. For late boomers, who are already in their 40s and 50s, suicide rates are still climbing and thus they will reach their peak later in life relative to the cohorts that preceded them. For both men and women, the age curves by cohort exhibit non-parallelism, indicating that period effects are present.

In sum, the descriptive analysis suggests that, compared to previous cohorts, the baby boomer generation had among the highest rates of suicide in adolescence, for both males and females. However, current rates of suicide in middle age among early boomers are in line with, and somewhat lower than, those shown by prior cohorts. On the other hand, the patterns among late boomers in middle age do appear somewhat anomalous in terms of level and trend. Analytic models will determine the degree to which these patterns hold once age, period and cohort effects are assessed simultaneously.

### 3.2. Analytic models

Table 1 displays the results of six reduced log linear models. Among the three one-factor models, age variation best accounts for temporal patterns of suicide over the period for both men and women. This result is similar to those of past analyses of temporal change in mortality and is not surprising given the strong association between age and suicide risk for both sexes. Variations across time periods are least effective in explaining suicide trends for both sexes. For men, the age-cohort model exhibits the superior fit among the two-factor models; for women, the age-period model best explains the temporal variation. However, both for men and women, the model accounting for all three components of change, age, time period and cohort, best explains the changes in suicide rates over time, confirming the patterns found in the descriptive analysis.

**Table 1**  
Goodness-of-fit statistics for age–period–cohort log linear models of U.S. suicide rates between 1935 and 2010, by sex.

	Age	Period	Cohort	Age–period	Age–cohort	Period–cohort	Age–period–cohort
	A	P	C	AP	AC	PC	APC
<b>Males</b>							
Deviance	17273.21	37361.65	20908.28	14858.83	6736.10	8045.71	2679.30
AIC	19025.96	39122.40	22689.04	16641.58	8538.85	9858.47	4512.05
BIC	19065.05	39174.52	22773.73	16729.53	8659.38	9955.28	4681.44
DOF	180	176	166	165	155	150	140
<b>Females</b>							
Deviance	5321.62	10686.96	8439.84	1536.49	3202.99	5650.00	875.36
AIC	6848.39	12221.73	9994.61	3093.26	4779.76	7236.77	2482.13
BIC	6887.48	12273.85	10079.30	3181.21	4900.29	7373.59	2651.52
DOF	180	176	166	165	155	150	140

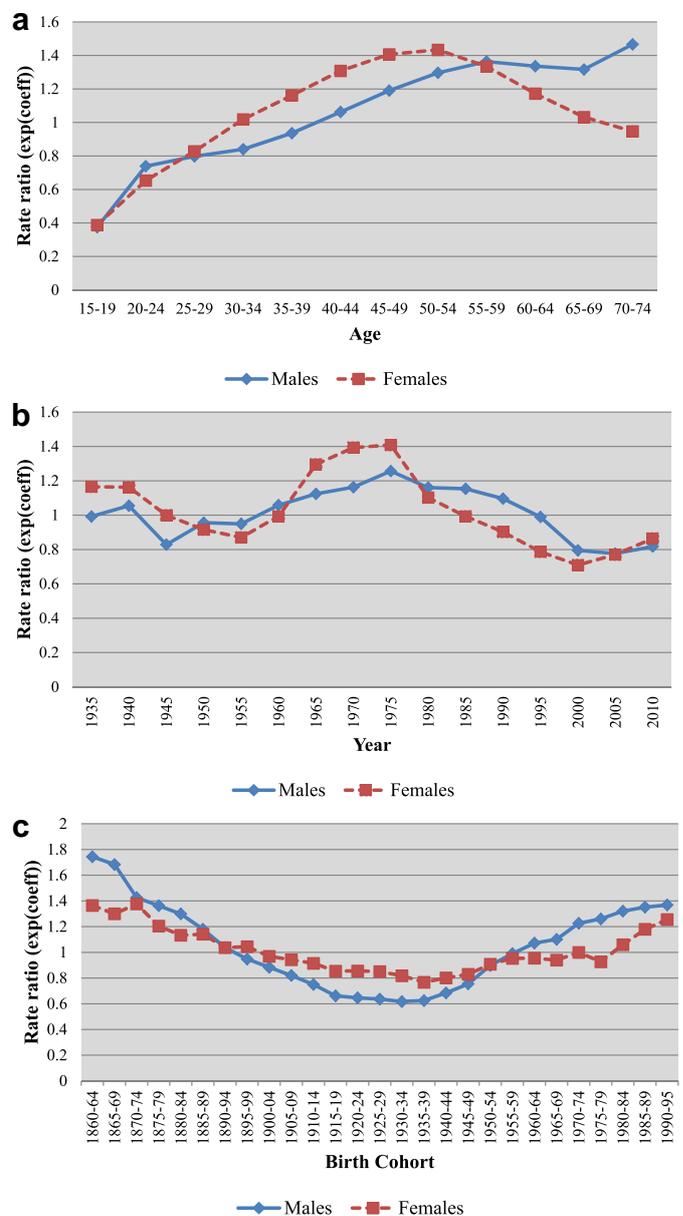
Notes. AIC = Akaike's Information Criterion. BIC = Bayesian Information Criterion. DOF = Degrees of Freedom. The smaller the AIC and BIC, the better the model fit.

Table A1 shows the results from the estimation of the full APC models using the IE method for males and females. Following Yang (2008), I created graphs of the exponentiated coefficients to show the net effects of age, time period, and cohort on overall suicide trends. The results largely confirm those of the descriptive analyses. Looking first at Fig. 4a, we see the different age pattern of suicide for males and females. For both sex groups, suicide rates rise sharply in adolescence. Female rates continue to rise steadily through middle age before declining in the older age ranges. The rate of suicide for males appears to grow steadily over the life course. For both groups, the age effects are the largest of the three.

Fig. 4b shows the variation in suicide rates by time period and reveals a pattern similar to that observed in Fig. 1, which shows the crude suicide rate and applies a joinpoint analysis to identify changing trends over the period. The pattern of period effects is similar for both males and females, although the effects appear somewhat more pronounced for females. Net of age and cohort effects, we observe the increase in suicide rates beginning in the late 1950s and continuing through until 1980, with sharp declines until 2000. Since 2000, suicide rates have been rising.

Finally, Fig. 4c displays the net cohort effect on suicide rates for males and females. For males, rates are lowest for birth cohorts born between 1915 and 1939. Rates begin to increase for the boomer cohorts and have continued to do so with subsequent birth cohorts. The average suicide rate over the entire period for males is about 28 per 100,000 (the exponential of the intercept,  $-8.1792$ ). Males born in 1930–34 have an average rate that is reduced by a factor of 0.62–17.4 per 100,000; for males born between 1955 and 1959, the average rate is essentially the same as the average for the period while for males born between 1985 and 1989, the average rate is estimated to be 37.8 per 100,000, 1.35 times greater than the overall average. Relative to the 1930–34 birth cohort with the lowest rate over the period, male late boomers have an average suicide rate that is roughly 1.6 times as large. Males born in the 1985–89 period have an average rate about 2.2 times that of the 1930–34 cohort. Patterns are similar, but less pronounced for females, with the average rate beginning to rise with the boomer cohorts and those that followed.

To assess the robustness of the IE findings, I ran two additional models that apply different constraints, namely the traditional CGLIM model and the new ZLT constraint model. Since time period proves to be the least consequential determinant of suicide in the reduced log linear models, I chose to constrain the effect for time periods 1935 and 1940 to be equal in the CGLIM approach. The ZLT model applies a zero linear trend to the period coefficients. Both the CGLIM and the ZLT models yield similar substantive results to those of the IE solution. The age effects are virtually identical for males and females across the three models. The period effect estimated by



**Fig. 4.** a) Intrinsic estimates of age effects of suicide, by sex. b) Intrinsic estimates of period effects of suicide, by sex. c) Intrinsic estimates of cohort effects of suicide, by sex.

the ZLT model for males and females is also very similar to that estimated by the IE approach; the period effect for males only in the CGLIM models shows more of a steady decline than that estimated by the IE models. Of particular interest, both models show a similar pattern to the cohort effects. Suicide rates are lowest among those born in the early part of the 20th century but start to rise with baby boomers and continue to increase for subsequent cohorts for both males and females. Results are available in the [Online appendix](#).

#### 4. Discussion

To explain the age, period and cohort patterns of suicide rates, Durkheim's classic statement about the relationship between social structure, namely levels of social integration and regulation, and suicide remains germane in many respects, despite its introduction over a century ago (Durkheim, 1951). For example, suicide risk is closely tied to age for both men and women, but the age pattern is specific to sex and may reflect the differing levels of stress and social isolation experienced by men and women at various stages of the life course. Historically, suicide for women peaks in the middle age ranges when they are active parents, experiencing the empty nest, and undergoing physical (menopause) changes. In contrast, rates for men peak in old age when their primary forms of social connection and support disappear through death of a spouse and retirement (Grzywacz et al., 2002).

The period effects appear similar by gender, although somewhat stronger for women than men, a finding consistent with those from Switzerland (Ajdacic-Gross et al., 2006). A period effect that likely produced declining suicide rates during the 1940s is World War II (Thomas and Gunnell, 2010) while the declines during the late 1980s and 1990s have been linked by some to the development of new antidepressants (Gibbons et al., 2005). Fluctuations in rates of unemployment (Luo et al., 2011), alcohol consumption (Gruenewald et al., 1995), religious composition and immigration (Phillips, 2013) are also associated with suicide trends, although most longitudinal studies of suicide examine only bivariate associations and do not consider the confluence of factors that may produce change. Nonetheless, the period effects commonly linked to fluctuating suicide rates are manifestations of the underlying degree of social integration and regulation within a society.

Finally, I find evidence of cohort effects, with sharp differences in suicide rates across the 27 cohorts included in this analysis. Taking a historical view, the baby boomer generation is *not* anomalous in its suicide rate. In fact, early boomers have a lower than average rate of suicide while late boomers hover around the average suicide rate for all cohorts combined. However, boomers do appear unusual in marking the beginning of an important shift in patterns across cohorts. I find that suicide rates for male cohorts begin to rise sharply with the baby boomer generation and continue to do so for all subsequent birth cohorts. A similar increased rate in suicide for the postwar generations is observed for U.S. women although less pronounced than that for men. This general pattern of rising suicide rates among recent birth cohorts concurs with findings from other countries (Ajdacic-Gross et al., 2006; Gunnell et al., 2003).

How can we make sense of these cohort patterns? Certainly, the 1960s ushered in broad social and cultural changes which weakened traditional forms of social integration and regulation. Divorce rates rose while marriage rates declined, leading to a dramatic rise in the percentage of the population living alone; 28% of all households were comprised of a single adult in 2010, compared to just 9% of households in 1950 (Klinenberg, 2012). Similarly, evidence suggests that religious involvement is in decline in the United States; the postwar cohorts have not shown the increase in religiosity that earlier cohorts have displayed as they aged (Hout and Fischer,

2009; Miller and Nakamura, 1996) and increasing numbers of Americans self-report as Atheists or without a religious identity (Rosch, 2012). Rising obesity rates are compromising the health status of cohorts, particularly as they move into the older age ranges. During the first decade of the twenty-first century, a variety of unstable economic conditions signaled insufficient social regulation. For instance, the nature of work has changed over the past two decades or so (for men especially), with declines in long-term employment and increases in short-term employment for males as they enter their thirties and beyond (Farber, 2007).

These societal transformations are undoubtedly reflected in the period effects, but at the same time, they likely also operate as cohort effects. The impact of such changes is determined in part by the stage of the life course at which people experience the event. Economic shocks, for example, probably have a disproportionate adverse effect on those in midlife, as such individuals are more likely to be family breadwinners and supporting dependents. In contrast, the young and the elderly are less likely to be in the labor force and often have the support of family and in the case of the elderly, social security. Thus, economic conditions, a period effect, interact with age to create a quintessential cohort effect. Certainly, evidence suggests this was the case with regard to the economic recession of 2007–2009 (Chang et al., 2013). In contrast, the experience of growing up during the Great Depression did not appear to elevate the suicide rate for the Silent Generation, perhaps because this cohort learned to persevere through difficult times. Another period effect, rising divorce rates, also probably intersects with age – boomers largely experienced this trend as young adults and now again in midlife as the gray divorce rate increased between 1990 and 2009 (Brown and Li, 2012). The effect is likely different, and perhaps more detrimental, for subsequent cohorts, who experienced these shifts as children and teenagers.

The APC analysis finds larger cohort effects for men, which is not surprising in some respects. The societal changes that occurred beginning in the late 1950s and early 1960s, including the rise in female labor force participation, the increasing prevalence of women in institutions of higher education, and changes in family formation patterns, have had profound implications for gender relations (Rosin, 2012). In a variety of ways, the status of men is threatened as institutions that historically have been good for them – education and marriage, for instance – are being transformed in fundamental ways. A college education is becoming increasingly more critical in this service economy, but women have been outpacing men for several decades in their acquisition of degrees and providing increasing competition for good jobs (National Center for Education Statistics). As economic opportunities for women grow, marriage becomes more dispensable. Declining marriage rates mean that male partners are less likely to reap the benefits to marriage, such as those to health which appear greater for men than for women (House et al., 1988; Umberson, 1992). Furthermore, relative to the past, women are increasingly likely to be the primary breadwinner within a marriage but research indicates that this dynamic can have *detrimental* effects on men's health outcomes (Springer, 2009). All these forces may combine to affect psychological well-being and stress levels, and to enhance feelings of social isolation, among birth cohorts of American men in the young and middle-age ranges in recent decades. In essence, the playing field has changed dramatically for men, and many may not yet have adjusted to the new rules, producing increases in anomic suicide.

The evidence for increasing suicide rates among recent cohorts of American women is more tenuous. Many of the above-described changes have been beneficial for women (Schnittker, 2007) – steady full-time work has positive effects on physical and mental health among mothers (Frech and Damaske, 2012) and in 2000, 61% of women with children under the age of three worked outside the

home, compared to just 34% in 1975 (Committee on Ways and Means, U.S. House of Representatives, 2000). However, there are potential costs associated with these changes (Bianchi et al., 2006; Hochschild, 1989). Working mothers in dual-career marriages report that, compared to their husbands, they take on far more of the housework and child care responsibilities, coming home to what Arlie Hochschild coined a “second shift”. Even in families in which the father is involved in child care, the number of which has undoubtedly increased in recent decades, men do more of the “fun” child activities while women conduct more of the basic child maintenance. These factors may increase stress levels and negatively affect quality of life as both women and men navigate a rapidly-changing work and family environment (in Durkheimian terms, weak social regulation); if so, they may play a role in explaining the pattern of rising rates of suicide among recent birth cohorts of women, particularly for those who are more disadvantaged (Frech and Damaske, 2012).

Other forms of social integration and regulation within our society are evolving and may affect the postwar cohorts (younger than age 60) more profoundly than preceding generations, producing rising levels of egoistic suicide. Robert Putnam in his book *Bowling Alone* documents declining rates of civic participation (e.g. in the political process, community groups and voluntary associations) that have reduced the numbers of group social interactions (Putnam, 2000). Rising rates of mobility and suburbanization may have accelerated this trend, although some contend that traditional forms of civic participation have simply been replaced with alternative ones (Lemann, 1996). More recently, attention has focused on how new forms of social media affect social integration and alter the ways in which we communicate and connect in dramatic ways. These tools have greatly expanded our networks of friends and acquaintances and enhanced our levels of connectedness yet these new types of connections have emerged alongside growing rates of loneliness. Studies indicate that loneliness declines with a greater relative frequency of face-to-face contact but increases with a greater proportion of online interactions (Cacioppo and Patrick, 2008), suggesting that to the extent that online interaction reduces personal connections, it may contribute to higher levels of isolation. Turkle (2008) argues that the link is direct, with technology promoting superficial and tenuous connections and shallow degrees of intimacy.

Perhaps as a reflection of these changes, we have witnessed rising rates of diagnosed mental and psychiatric disorders over the past three decades. In 1987, the rate of outpatient treatment for depression was 0.73 per 100 persons, but rose to 2.33 by 1997 (Olfson et al., 2002). However, changing diagnostic criteria and increased use of managed care make it difficult to ascertain whether these rising rates represent a true increase in the prevalence of psychiatric disorders. Since the 1980s, rates of suicide by firearms have been declining at the same time that rates of suicide by drug poisoning, particularly for women, have been increasing in the early 2000s, perhaps related to the sharp increase in the use of antidepressants and prescription drugs more generally (Zuvekas, 2005). If these kinds of shifts over time in methods used to commit suicide are related to birth cohorts, they also offer a potential explanation for the cohort effects observed.

This analysis provides a first important step – using seventy-five years of data to identify, describe and distinguish the patterns of suicide rates across age groups, periods and cohorts – and notes a number of important societal-level changes that have affected forms of social integration and regulation and are consistent with the observed patterns. Future research that can directly assess the extent to which the above-described factors explain rising rates of suicide among recent cohorts is now needed. The development of hierarchical APC models may be one fruitful way in which to test

some of the conjectures put forward as individual-level data mitigates the possible problem of the ecological fallacy inherent in Durkheimian explanations of suicide (see also Makinen, 1997). Moreover, efforts to explore variation in suicide rates among individuals within particular cohorts and by race/ethnicity may shed light on explanations for the observed patterns.

As with all work of this nature, there are limitations to the measurement of suicide deaths. Between 1935 and 2010, the National Center for Health Statistics categorized deaths according to six different International Classification of Death (ICD) schemes and these revisions may have affected the recording of suicides over time. However, consistent with data from England and Wales (Thomas and Gunnell, 2010), there do not appear to be obvious discontinuities in the overall suicide rate or suicide rates by gender in the years in which successive revisions of the ICD were introduced. In addition, the under-reporting of suicide deaths in official statistics may have declined over time as suicide becomes less stigmatized, although medical examiners are not greatly affected by pressure exerted by relatives to avoid the stigma of suicide (Timmermans, 2005). Distorting the statistics in the other direction is the declining rate of autopsies in the U.S. since the 1970s, which would tend to increase under-reporting. Shifts in method over time also affect reporting – drug poisoning accounts for a smaller proportion of suicides today than in the past and so suicides may be less likely to be under-reported in the more recent period since such incidents are more likely to be classified as “undetermined” or “accidental”. Related to this point, there are important gender differences in the method of suicide used (females are more likely to commit suicide using drug poisoning relative to men) and hence female suicides may be underreported relative to men and affect the reported results.

## 5. Conclusion

Suicide rates declined during the last two decades of the 20th century due to a variety of period effects and a favorable age composition, but patterns have started to reverse as the large boomer cohort (particularly males) moves into older age ranges with traditionally higher suicide rates and with the development of increasing suicide rates among the middle-aged. Meanwhile, the oldest age ranges, those with the highest suicide rate (for males), are currently occupied by the Silent Generation (born 1925–45), the cohort with the lowest suicide rate of all. Thus, we have a new epidemiology of suicide whereby middle-aged suicide rates exceed those for the elderly.

Applying a Durkheimian lens, I speculate that the broad social and economic changes introduced in the 1960s may have weakened traditional forms of social integration and regulation for the postwar cohorts, leading to a pattern of rising suicide rates. As younger cohorts move through the life course and more data become available, we should closely monitor the extent to which these emerging patterns hold. As new family and work structures become more normative and as people recalibrate their expectations, patterns may shift. The size and composition of cohorts can be an engine for social change, altering both the structures and functions of institutions and our perception of them.

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Appendix A

**Table A1**  
Intrinsic estimates for U.S. suicide rates, 1935–2010.

	Males			Females		
	Exp(Coeff.)	95% conf. interval		Exp(Coeff.)	95% conf. interval	
		LCL	UCL		LCL	UCL
Intercept	0.0003	0.0003	0.0003	0.0001	0.0001	7.7E-05
<b>Age</b>						
15–19	<b>0.374</b>	0.348	0.402	<b>0.388</b>	0.357	0.421
20–24	<b>0.739</b>	0.700	0.780	<b>0.653</b>	0.613	0.696
25–29	<b>0.798</b>	0.756	0.841	<b>0.828</b>	0.781	0.878
30–34	<b>0.840</b>	0.797	0.886	1.018	0.965	1.075
35–39	<b>0.936</b>	0.888	0.986	<b>1.162</b>	1.103	1.224
40–44	<b>1.062</b>	1.009	1.119	<b>1.308</b>	1.244	1.376
45–49	<b>1.190</b>	1.130	1.253	<b>1.407</b>	1.338	1.479
50–54	<b>1.296</b>	1.230	1.366	<b>1.434</b>	1.361	1.510
55–59	<b>1.364</b>	1.292	1.439	<b>1.334</b>	1.262	1.411
60–64	<b>1.336</b>	1.261	1.416	<b>1.172</b>	1.101	1.248
65–69	<b>1.317</b>	1.237	1.401	1.032	0.963	1.106
70–74	<b>1.466</b>	1.373	1.566	0.947	0.876	1.023
<b>Period</b>						
1935	0.992	0.923	1.067	<b>1.165</b>	1.081	1.256
1940	1.055	0.982	1.134	<b>1.163</b>	1.078	1.254
1945	<b>0.828</b>	0.761	0.901	0.999	0.923	1.082
1950	0.956	0.886	1.032	<b>0.916</b>	0.845	0.993
1955	0.949	0.878	1.025	<b>0.870</b>	0.802	0.943
1960	1.058	0.983	1.139	0.994	0.922	1.071
1965	<b>1.124</b>	1.046	1.207	<b>1.296</b>	1.214	1.383
1970	<b>1.163</b>	1.085	1.246	<b>1.393</b>	1.310	1.482
1975	<b>1.256</b>	1.178	1.339	<b>1.408</b>	1.327	1.495
1980	<b>1.160</b>	1.090	1.236	<b>1.102</b>	1.034	1.175
1985	<b>1.153</b>	1.086	1.225	0.993	0.931	1.060
1990	<b>1.096</b>	1.034	1.162	<b>0.902</b>	0.845	0.964
1995	0.989	0.934	1.048	<b>0.787</b>	0.735	0.842
2000	<b>0.795</b>	0.749	0.844	<b>0.710</b>	0.662	0.761
2005	<b>0.776</b>	0.732	0.823	<b>0.770</b>	0.721	0.823
2010	<b>0.817</b>	0.771	0.866	<b>0.864</b>	0.809	0.923
<b>Cohort</b>						
1860–64	<b>1.743</b>	1.320	2.302	1.364	0.917	2.029
1865–69	<b>1.684</b>	1.401	2.023	<b>1.300</b>	1.008	1.676
1870–74	<b>1.426</b>	1.224	1.662	<b>1.376</b>	1.144	1.656
1875–79	<b>1.363</b>	1.201	1.547	<b>1.204</b>	1.029	1.409
1880–84	<b>1.299</b>	1.165	1.448	<b>1.133</b>	0.991	1.295
1885–89	<b>1.180</b>	1.069	1.303	<b>1.141</b>	1.017	1.280
1890–94	1.038	0.945	1.140	1.037	0.933	1.152
1895–99	0.948	0.866	1.038	1.044	0.949	1.148
<b>G1 Generation</b>						
1900–04	<b>0.883</b>	0.810	0.964	0.969	0.886	1.060
1905–09	<b>0.820</b>	0.755	0.891	0.942	0.866	1.025
1910–14	<b>0.750</b>	0.692	0.812	<b>0.914</b>	0.844	0.990
1915–19	<b>0.662</b>	0.611	0.717	<b>0.854</b>	0.790	0.923
1920–24	<b>0.646</b>	0.597	0.700	<b>0.855</b>	0.792	0.923
<b>Silent Generation</b>						
1925–29	<b>0.636</b>	0.587	0.689	<b>0.851</b>	0.787	0.919
1930–34	<b>0.617</b>	0.568	0.670	<b>0.818</b>	0.755	0.886
1935–39	<b>0.625</b>	0.577	0.677	<b>0.767</b>	0.708	0.831
1940–44	<b>0.684</b>	0.635	0.736	<b>0.801</b>	0.743	0.863
<b>Early Boomers</b>						
1945–49	<b>0.755</b>	0.706	0.807	<b>0.829</b>	0.773	0.889
1950–54	<b>0.898</b>	0.844	0.956	<b>0.908</b>	0.849	0.971
<b>Late Boomers</b>						
1955–59	0.990	0.931	1.053	0.953	0.890	1.020
1960–64	<b>1.072</b>	1.005	1.144	0.954	0.885	1.029
<b>Generation X</b>						
1965–69	<b>1.101</b>	1.023	1.186	0.939	0.860	1.027
1970–74	<b>1.226</b>	1.129	1.331	0.999	0.903	1.106
1975–79	<b>1.260</b>	1.146	1.386	0.927	0.820	1.047
<b>Millennials</b>						
1980–84	<b>1.320</b>	1.184	1.472	1.059	0.924	1.215
1985–89	<b>1.351</b>	1.179	1.549	1.179	0.998	1.393
1990–95	<b>1.368</b>	1.072	1.745	1.255	0.951	1.657
<b>Deviance</b>	2679.3			875.36		
<b>DOF</b>	140			140		

Notes. Bold = statistically significant at  $p < 0.05$ . Exponentiated coefficients interpreted as rate ratios.

Appendix B. Supplementary data

Supplementary data related to this article can be found at <http://dx.doi.org/10.1016/j.socscimed.2014.05.038>.

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