Full paper for the Annual Meeting of the Population Association of America, March 31 - April 2, 2016, Washington, DC

Version 7 March 2016

Cohort Factors Impinging on Suicide Rates in the United States, 1990-2010

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Abstract

We use CDC microdata on cause of death and CPS data on populations by age to create suicide rates for five-year age groups at five-year intervals, further cross-classified by race/ethnicity, education, and marital status. We examine the suicide history 1990-2010 of U.S. birth cohorts, net of age and cohort linear trends. These de-trended cohort deviations follow familiar patterns: most pronounced in the Baby Boom, least pronounced during the Baby Bust, they illustrate the so-called Easterlin effect. Suicide rates for women show similar patterns as suicide patterns for men. We show persistence of those effects net of micro factors (especially education and marriage) implicated in suicide behavior and correlated at the macro level with relative cohort size. Analysis of suicide patterns over time for high- and low-educated men and women shows that white men with low education face a sharp increase, significantly above the linear time trends, in suicide rates among cohorts born between 1955 and 1970. This bump is mostly unrelated to secular trends of increasing average educational attainment rates, at least if no interaction between age and cohort is involved in the explanation. No obvious pattern related to cohort size is found for African-American high- and low-educated men, which makes sense given the very different historical dynamics for this minority sub-population.

Background

Suicide is generally associated with a combination of privations; life reversals (including loss of social status, income, and employment; plus divorce); despair and loss of hope, aspirations and positive expectations about the future; and a lack of social integration that might otherwise temper the various combinations of the other factors. Durkheim argued that both *integration* (the context of deep social embeddedness) and *regulation* (the capacity of individuals to maintain a correct, positive balance between realities and expectations) prevent suicides. In both sociology and epidemiology, an enduring question has been, How is it that the most extraordinary of individual behaviors can be best evidenced by factors that are potentially supra-individual? On a population level, suicide rates may increase if economic conditions are bad or rapidly decreasing (Phillips & Nugent 2014). Suicide rates can be partly explained by individual-level factors such as family size (Denney et al. 2009; Denney 2010), but have in studies with aggregate data also been shown to respond to period and cohort effects and other contextual factors (Pampel 1996; Pampel & Peters 1995; Stockard and O'Brien 2002a; 2002b; 2006). One contextual-level factor influencing suicide rates is the size of the birth cohort into which one is born, providing or depriving opportunities to succeed in realizing one's expectations. This effect of size of birth cohort has been coined the Easterlin effect (Easterlin 1980), referring to the hypothesis that individuals born into larger birth cohorts fare worse than individuals born into smaller birth cohorts. A larger birth cohort may render jobs in short supply relative to the cohort's needs, diminish or cloud economic prospects, exacerbate competition, and make upward social mobility less likely and downward social mobility more of a threat.

We describe suicide trends between 1990 and 2010 for men and women, adjusting for individual sociodemographic characteristics. We test the extent to which the Easterlin effect holds for suicide mortality for the United States, especially after adjustment for micro level correlates of suicide, and consider how and why, among birth cohorts, the propensity toward suicide operates differentially. We take advantage of information on individual-level variables to examine suicide patterns separately for gender, race, and educational level (university diploma and above versus completed high school and below), with a focus on middle-aged white men with low education, giving some further structure to the observations of Case and Deaton (2015), that rising morbidity and mortality levels specifically reflect a cohort effect. More precisely, this frustrated social generation hypothesis[1] is rapidly presented and not demonstrated by the authors; additionally, they mention an increasing educational gradient they do not consider further[2]. Our aim here is to fill these gaps: we deepen the cohort-based diagnosis and analyze how the educational gradient of suicide is related to birth cohort dynamics.

[1] "After the productivity slowdown in the early 1970s, and with widening income inequality, many of the baby-boom generation are the first to find, in midlife, that they will not be better off than were their parents." (p.5)
[2] "The final rows of <table 1> show increasing educational gradients from 1999 and 2013" (p.3)

Data

Data on suicide come from the Center for Disease Control and Prevention (CDC) Mortality Multiple Cause Files.[3] This annual micro dataset contains information annually on approximately two million decedents, including age, cause of death, and important socioeconomic descriptors. We examine the N=130,479 suicides for five-year age groups for the years 1990, 1995, 2000, 2005, and 2010. We distinguish male and female suicides, not only because of well-known differences in suicidal behavior (suicides are approximatively fourfold more common among men, despite more failed suicide attempts in women) but also because of gender-specific differences in labor force engagement and attachment (a crucial factor in the theory of the Easterlin effect) that narrowed across the cohorts represented in this study. For these male and female decedents we observe their contemporaneous level of education, a measure of relative social status and perhaps permanent income; race/ethnicity (white/African-American/other/Hispanic); and marital status (single/married/widowed/divorced or separated).

To construct a reference population at risk of suicide, we use the Current Population Survey (CPS) for the respective year.[4] For the five years corresponding to the CDC death data we obtain counts of the population for five-year age groups cross-classified by education, race/ethnicity, and marital status. Because the CPS reports level of education in terms of diplomas whereas Vital Statistics uses years of schooling, we employ a four-category schema—less than high school graduate; high school graduate (including GED); some college; and college graduate and beyond—that is relatively unambiguous with respect to both data sources. The CPS reference population consists of 1,600,481 cases, before application of extrapolation weights. We match the two micro datasets (deaths and population at risk) to compute suicide rates by age over time, cross-classified by the individual-specific sociodemographic factors. Counts of the observed population can be blended with death information and common demographic life-table assumptions to create estimates of person-years at risk. Alternatively, individuals dying due to suicide can be subtracted from similar estimates of beginning-year populations at risk to create an aggregated cross-classification of micro data: suicide (or not) × year × age × education × race/ethnicity × marital statistics. This allows computation of probabilities or odds. In practice, with the incidence of suicide being low, and intervals of observation short (a year), logs of rates, probabilities, and odds are all but indistinguishable from one another.

Method

Consider a generalized equation of the form

[3] Vital Statistics Data Available Online at http://www.cdc.gov/nchs/data_access/Vitalstatsonline.htm
[4] Available at cps.ipums.org.

$$\ln D_{ijk} = \ln P_{ijk} + \mu + \sum_{i=1}^{I} \alpha_i + \sum_{j=1}^{J} \beta_j + \sum_{k=1}^{K} \gamma_k + \delta^A A_i + \delta^P P_j + \delta^C C_k,$$

a representation of a Poisson model for rates (the denominator P_{ijk} being fixed by design), where—to simplify the exposition—notation for sociodemographic characteristics (and corresponding parameters) have been suppressed. The α_i, β_j, and γ_k are the familiar fixed effects for the *I* ages, *J* periods (years), and *K*=*I*+*J*-1 cohorts represented in the rectangular *A*×*P* design. The A_i, P_j, and C_k are vector transformations of the indexes *i*, *j*, and *k* (see Holford 1983 or Chauvel & Schroeder 2014 for alternative linear transformations). If we ignore the terms δ^A, δ^P, and δ^C, it is well known that even with the effect normalization

$$\sum_{i=1}^{I} \alpha_i = \sum_{j=1}^{J} \beta_j = \sum_{k=1}^{K} \gamma_k = 0$$

the model is under-identified, due to the fundamental linear combination *A*=*P*-*C* (e.g., Fienberg 2013). The alternative way of thinking about this is that only two of the three linear terms δ^A, δ^P, and δ^C are identified; but that once a decision is made on that front, the further constraints

$$G^A \sum_{i=1}^{I} \alpha_i A_i = \delta^A, \quad G^P \sum_{j=1}^{J} \beta_j P_j = \delta^P, \quad \text{and} \quad G^C \sum_{k=1}^{k} \gamma_k C_k = \delta^C,$$

do identify the α_i, β_j, and γ_k parameters, not as effects of age, period, and cohort *per se*, but as deviations from a two-dimensional plane in time. (The G^* terms are normalizing scalars depending on the linear transformation of the indexes). A restriction on one of the linear components of age, period, and cohort is in no sense a general solution to the fundamental problem of identification of so-called cohort effects in data cross-classified by age and time. Rather, it is an identification that accords with our particular conceptualization of cohort effects, as deviations from a general time trend. We define that baseline trend as the linear trend in age and cohort. We hence set $\delta^P = 0$, as per Holford's (1983) original identification, which has been taken up by O'Brien (e.g., 2014) as the (period) zero linear trend (ZLT) model and by Chauvel & Schroeder (2014) as the APC-Detrended approach (APCD). Our analysis under this model emphasizes cohort contrasts where more advantaged birth cohort are systematically above the linear age-period-cohort trend compared with the comparably disadvantaged birth cohorts below the trend. Chauvel & Schroeder (2014, 2015) provide a more detailed substantive motivation for this focus.

The general model (so conceptualized and identified) can be used with either aggregated data on rates (a Poisson error specification and a log link) or with a micro motivated (binary error, logit or probit link) specification—not that the distinction matters much computationally in these data (see comments in prior section). It can be extended to including individuals' characteristics, specifically education, marital status, and race/ethnicity. We thus test whether fluctuations in suicide by cohort are due to changes in cohort composition with respect to these measured demographic characteristics (known to be correlated at the micro level with suicide) or whether they represent more “intrinsic” cohort characteristics, such as the (mis)fortune represented by (relatively larger) cohort size.

Results

APCD logit models were run without control variates, then with education, marital status and race as covariates for men (Table 1) and women (Table 2). Being married and better educated were protective with respect to the risk of suicide, whereas being white was positively associated with risk of suicide.

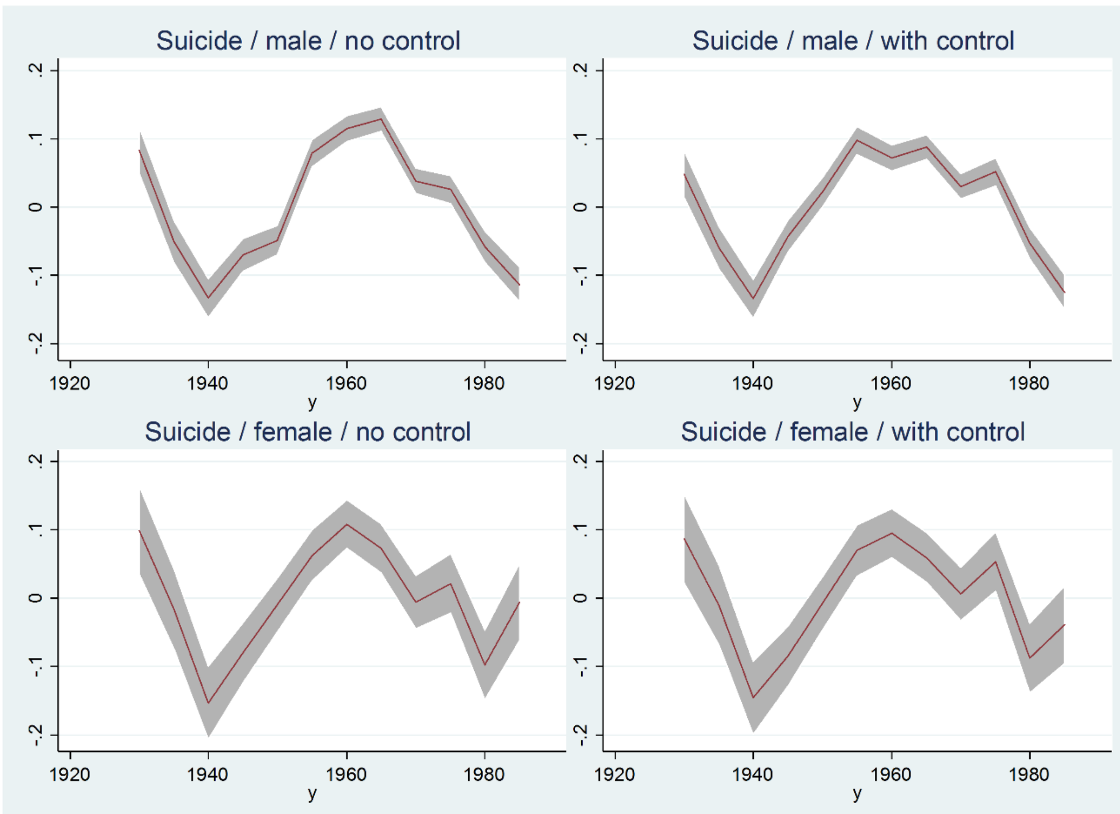
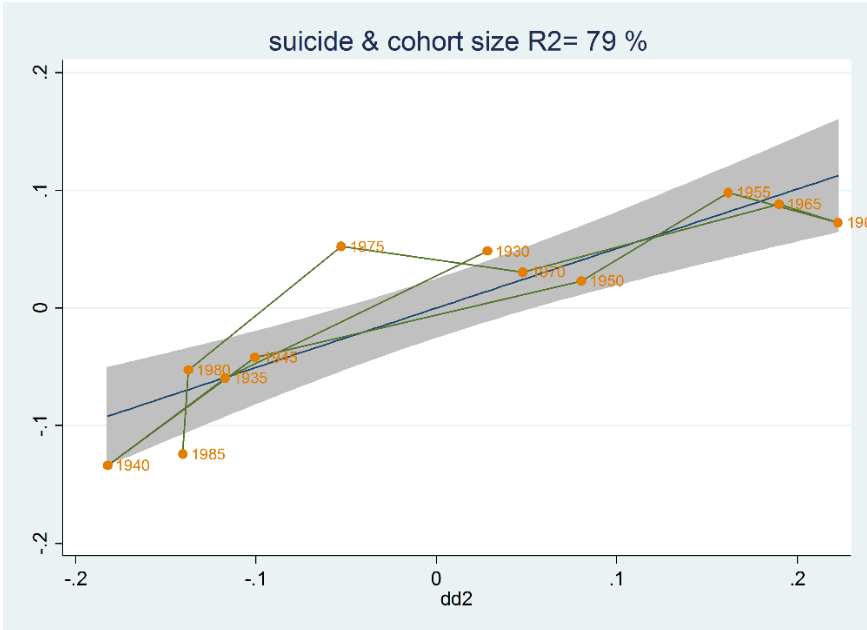


Figure 1. Cohort fluctuation for suicide by birth cohort with and without controls, males and females.

Figure 1 shows the cohort fluctuations in suicide rates for both men and women, without control variates (left hand side) and with control variates (right hand side). For men, the 1940 cohort has a -10 % reduced risk, whereas the 1960 cohort faces a +10 % increased risk of dying of suicide, relative to the linear trend in age and cohort. The control variables (see Tables 1 and 2) confirm what we know from Durkheim (1897) and his followers about suicide: marriage is protective. Although there is no doubt that, on average, so-called white Americans have more of the good things in life, they are also at higher risk of suicide, due (in the traditional Durkheimian interpretation) to lower levels of regulation. Education, however, has a protective role, and white Americans have on average more education. But the relative density of whiteness, and/or education, is changing only slowly across cohorts, and largely in a linear fashion, which is not what we are looking at with these cohort coefficients. When controls are included in the APCD model, the profile of cohort fluctuation remains almost unchanged, except for a slight reconfiguration of the relative rates of the peak suicide cohorts among males: without micro level controls, the apparent peak is for the cohorts circa 1965; whereas with controls the peak occurs for the circa 1955 birth cohorts, with a gradual fall-off (not rise) during the next decade of births. These are minor variations on the overall theme.

Plotting the APCD coefficients for cohort deviations from trend in rates of suicide against the APCD coefficient of size of birth cohort reveals a strong positive correlation of size of birth cohort and suicide rates per cohort for both genders (Figure 2). Specifically, in our reference population sample from CPS, we generate logN, the logged (extrapolation weighted) sample in each cell of the A×P table, and extract the cohort APCD vector which represents the relative detrended size of the cohort. We correlate these cohort coefficients with those previously obtained for the APCD model for suicide rates (Tables 1 and 2). The R^2 of this correlation suggests that, for men, a total of 79 % of the variance in suicide rates can be explained by the relative size of the birth cohort (upper panel of Figure 2). For women, associations of birth cohort size and suicide fluctuation are smaller—as should be expected given their lower implication in the labor force, particularly in the birth cohorts represented here—but show the same pattern as size of birth cohort contributes almost largely to the variance in suicide fluctuations ($R^2 = 72$ %; lower panel of Figure 2). Models were also run including both individual-level variables and contextual-level size of birth cohort. As expected given the strong correlations observed in Figure 2, the explanatory contribution of size of birth cohort absorbed almost all the effect of cohort contrasts.



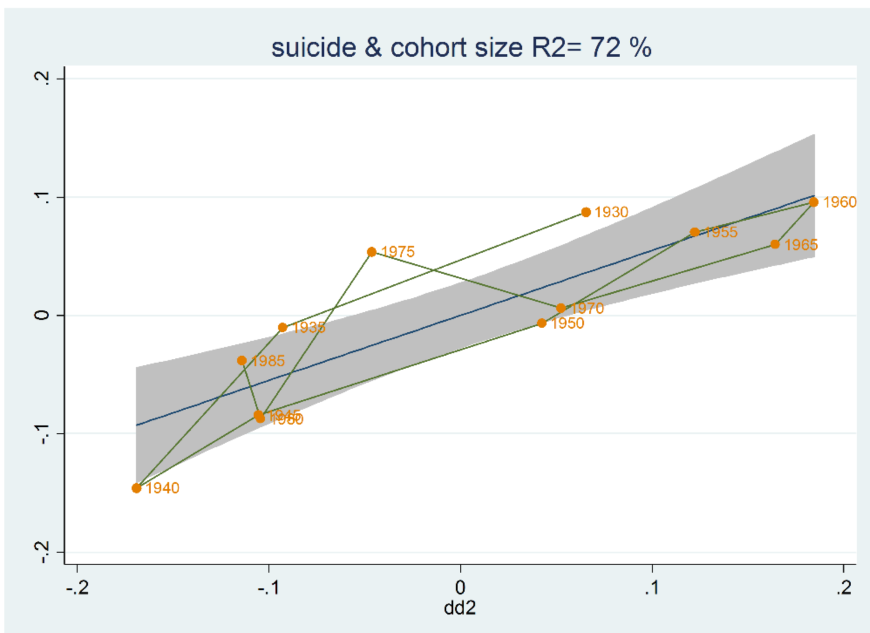


Figure 2. Associations of birth cohort size and suicide rates for men (upper graph) and women (lower graph) after controls.

The covariates added to the general APCD model assume that the “effects” of each demographic or socioeconomic variable are additive, just as the basic model does itself with the canonical terms age, period, and cohort since, for example, the only interactions of age and period are assumed to be captured in one specific linear combination, cohort. We now consider various interactions, in the sense of age patterns that vary by combinations of demographic and socioeconomic characteristics, as well as interactions between the canonical time measures that are not captured in the definition of cohort. Suicide rates (per 100,000) for different age groups in different survey years (1990-2010) are plotted in Figure 3 by gender and education (high [some college or above] versus low [no college]) for whites, and by education for African-American men. Any two dimensional rendering of rates (in this case age within periods) risks conflating period and cohort effects, mis-estimating age patterns—it has ever been thus! We take the risk nonetheless, since identification of these "layered" interactions has potential to add to our understanding of this historical social process.

For white men (certainly the better-educated white men), white women, and African-American men with high education, no clear patterns of suicide by age are observed in the cross-section (i.e., within periods, specific years of observation). For low-educated African-American men (lower right-hand panel of Figure 3), suicide rates are highest in the 20s then diminish continuously with age. There is one group with a dramatic bump in suicide rates: white low-educated men born around 1960 face a large increase in suicide rates. This increase is restricted to the time when this cohort enters middle age, as it is only visible above age group 40, no clear pattern being observable below (lower-left panel of Figure 3; lower left hand side). This represents a cohort-age or period-age interaction not specified in the additive APCD model. Net of longer term trends in expectations for well-being, the recent economic crisis—which we pick up in these data for the year 2010—was experienced especially badly (insofar as suicide is an indicator) by lower-educated men of middle age. Presumably they were at an age with responsibilities if not unmet expectations, little time, and fewer resources to cope. For women, suicide rates are lower, so deviations are less pronounced, but there is also evidence of a parallel bump for those with lower education among the cohorts reaching middle age in 2010 (middle panel of lower tier of Figure 3; translated into logs, the bumps for white men and women are similar in intensity). In contrast, the African-American population, which has a long history of exclusion and poverty, and quite a bit of experience with coping with challenges, prejudices, and thwarted ambitions does not seem to have been affected in the same way as the white less-educated population. The latter group has recently experienced an acceleration in a sense of loss, of a prior relatively secure (if not dominant) social and economic status, as per the 1960s model of middle-class America. We infer that that diminished opportunities for training and upward social mobility at that age contribute to increases in despair and gaps between aspirations and realities, resulting in increased population-level suicide rates.

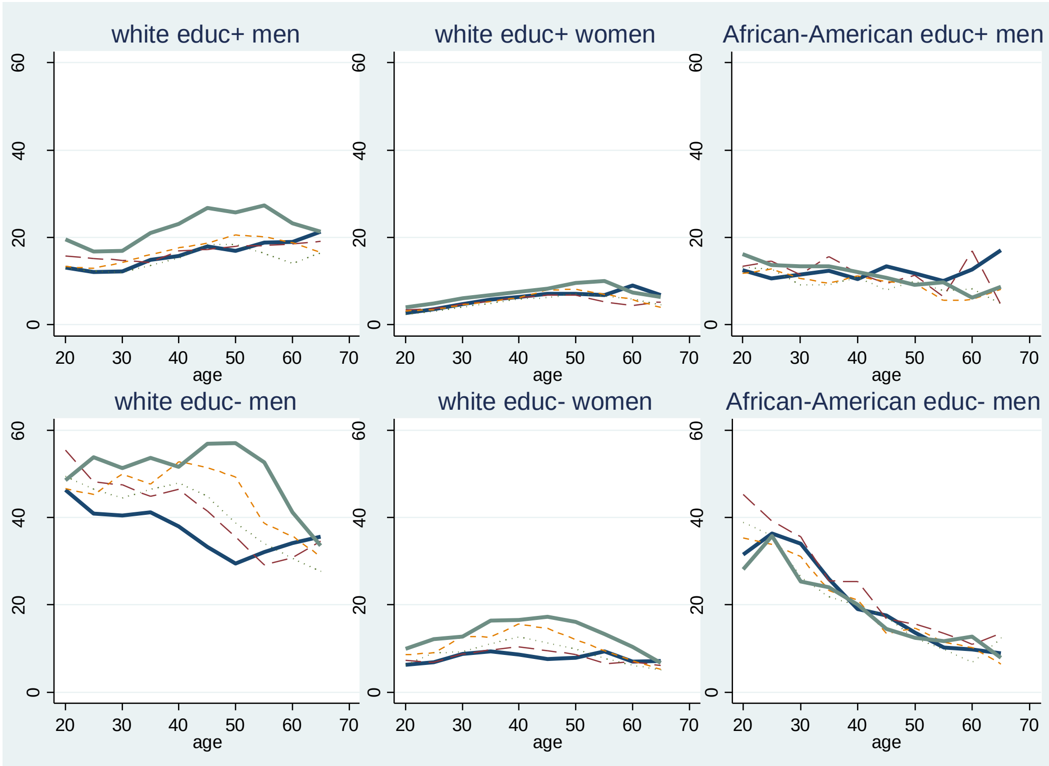


Figure 3. Suicide rates (per 100,000), age and education 1990-2010 in groups of the American population: men, women (both not separated by race), and African American men separate for high education and low education. Black thick lines = 1990, dark grey thick lines = 2010; dashed, dotted and short dashed lines represent 1995, 2000 and 2005 respectively.

What of Durkheim's theory, that the lack of social integration is main contributor to the risk of suicide? The group most vulnerable to social stress identified in our analyses is white low-educated men. Comparing the suicide fluctuations for married versus non-married high-educated versus low-educated men (Figure 4), we clearly see that the vulnerable group of white low-educated men is at dramatically higher risks of dying of suicide if not socially protected by marriage (rates above 100 suicides per 100,000 residents for the 2010 survey). Although higher educated non-married white men do have higher rates of suicide than similar men who are married, the relative risk of dying of suicide associated with marital status is substantially higher for low-educated white males. Education, or social status, is in part a substitute for social integration or, to turn it around slightly, in the absence of high(er) educational status, the absence of a spouse makes men especially prey to suicide.

Figure 4. Suicide rates (for 100,000), age, education and marital status 1990-2010 in groups of the American white male population, separate for married (right hand side) versus non-married (left hand side) and high education (associate university diploma and above, upper graphs) versus low education (completed high school and below, lower graphs). Black thick lines = 1990, dark grey thick lines = 2010; dashed, dotted and short dashed lines represent 1995, 2000 and 2005 respectively.

Finally, consider the relative contribution of size of birth cohort (*It's the demography, stupid!* Easterlin, 1987) and educational expansion (*It's the education, stupid!* Dowd & Hamoudi, 2014) to explaining fluctuations in suicide rates. Regressing suicide rates on indicators of cohorts population size (from CPS calculated as APC-D cohort coefficients of the logarithm of the population size per period and age cell) and educational average attainment (calculated as APC-D cohort coefficients of the logit model of probability to be associate university diploma or above per period and age cell), we find that population size explains the preponderance of cohort deviations in suicide rates (R^2= 55% with cohort size, 89 % if cohort size is lagged by five years; see Figure 5), whereas educational attainment only marginally contributes further. We generated several alternatives for the cohort nonlinearities of education. The one we retain here is based on the CPS based cohort APCD vector in a logit specification for some college or above versus high school graduate or below. When suicide APCD is correlated with education (after adjusting for cohort size), it simply adds 2 % to the overall R^2 with no lag, and 22% with a lag of five years of education. In a multiple linear regression of suicide APCD coefficients on those of education and of cohort size, the role of education is not significant. Education might matter, but it is largely by way of conditioning the effects of social integration (marriage), in particular among the white population. Cohort trends are more likely Easterlin effects, as Figure 6 makes clear[5].

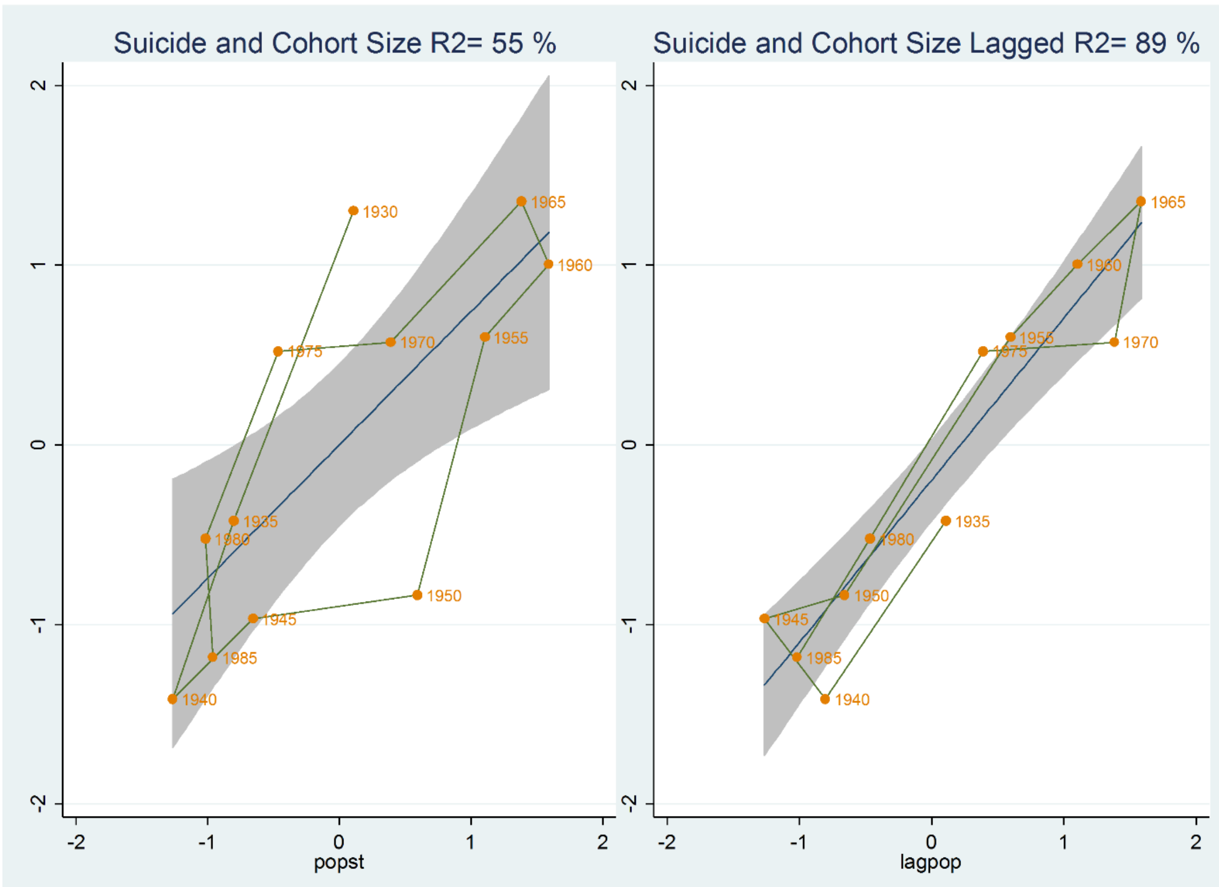


Figure 5. Illustration of the association of suicide rates and cohort size, adjusting for educational expansion, for the whole population, using concurrent cohort size (left panel) and five-year lagged cohort size (right panel).

[5] We tried a third type of explanation: "*It's the economy, stupid!*", with a third APCD detrended vector based on poverty variable. To generate this, we compute in the CPS the median family income equivalized by the square root of the household size and define relative poverty threshold as half of the median. This poverty dimension is very well (negatively) correlated with education (R=-.86) but poorly with suicide (R=-0.07). When poverty is introduced in the regression of suicide on population size and education, the adjusted R^2 declines; the variance inflation factor VIF diagnoses a strong collinearity of poverty with the other regressors (at first: education). We consequently avoid this economic perspective on suicides.

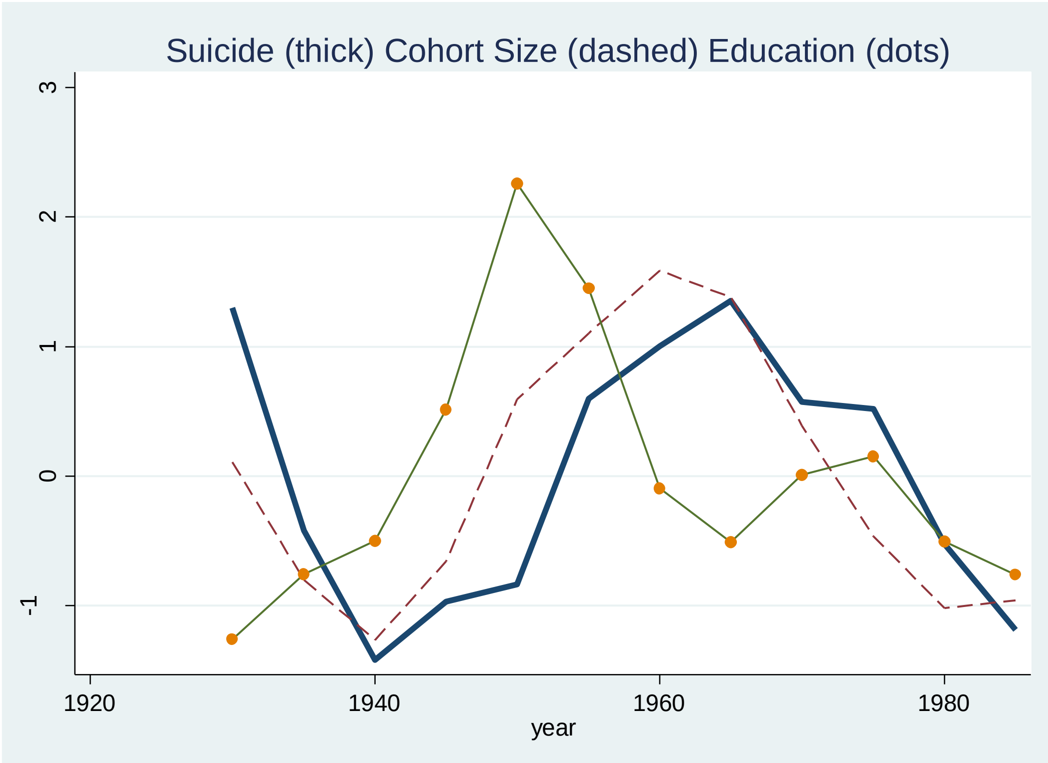


Figure 6. Suicide rates, educational expansion and cohort size APC-D standardized cohort coefficients (male population).

Discussion

Our analysis demonstrates that *size of birth cohort* is associated with risk of dying of suicide with data of the United States at the time between 1990 and 2010. As the overall explained variance suggests, size of birth cohort is actually the main driver of suicide rates, which has been coined the Easterlin effect earlier (Pampel & Peters 1995). Suicide is but one of the many symptoms of lack of social integration and regulation that is caused by large birth cohorts leading to overcrowding and competition for scarce resources. Recent evidence suggests that homicide (O'Brien & Stockard 2002; 2006) and homelessness (Sharygin, Smith et al., in preparation) are also strongly affected by the size of birth cohort.

Suicide fluctuations reveal little *gender differences*, with persistence of cohort effects for both genders even after controlling for important micro-level indicators of social status (education) and social integration (marital status). Despite the incidence of suicide being smaller for the female population, similar effects seem to operate here, and we find pronounced associations of suicide rates with size of birth cohort for both genders.

Regarding associations of educational attainment with the availability of opportunities, aspirations and associated (lower) suicide rates, one may argue that higher education could help in the competition for scarce resources in large birth cohorts, as education is an essential prerequisite for upward social mobility and increasing life chances. Consequently, lower-educated members of those cohorts may be at even higher risk for bad outcomes in terms of employment and status. However, this suggestion of members of large birth cohorts being even more dependent on educational resources is not supported by our data. Further, changes in stratification of general mortality (e.g. Olshansky et al. 2012) have often been explained with a changing meaning of education associated with *secular trends of increasing educational attainment* across time (i.e. those low educated after educational reforms are characterized by higher disadvantage compared to earlier-born cohorts who could not benefit from educational expansions, e.g. Dowd & Hamoudi 2014). However, this possible changed meaning of education associated with secular educational expansion trends is not reflected in our findings. Rather, what we find is a dramatic increase in suicide rates specifically for the circa-1960-born white low-educated men *at the time they enter middle age*. This finding reflects a composite of (a) *particular levels of social stress related with the economic crisis* and (b) a *particular vulnerability of the life stage above middle age*, presumably reflecting objective diminished opportunities for training and development for those cohort members after reaching middle age who may at this point realize that they will be constantly deprived of opportunities for social upward mobility still possible for younger age groups.

This confirms Easterlin's (1980) original predictions, but our micro-data based research shows that only some segments of the baby boom generation have been affected: the mechanism seems to play out mainly if not only in lower-educated whites. African-Americans may, by dint of a constant exposure to discrimination and precarity that dwarfs the demographically-induced social and economic fluctuations to which the dominant white population is exposed, may have functionally high levels of social integration (or experience with diminished aspirations) that buffer effects of transitory macro-level social stress and economic insecurity. Further, a detailed analysis on subgroups of the white male U.S. population shows that in particular low-educated *non-married* men, i.e. those without a classic form of social integration, are at particular risk, confirming Durkheim's propositions. As possible explanations for this rise we have put forward social stress related to economic crisis, economic insecurity, and a realization of diminishing opportunities and life chances. Our findings show that the 1955 to 1970 cohorts are, likely due to changes in socioeconomic regimes in growth, affected at a particularly vulnerable life stage. This group, if not socially protected (integrated) via marriage, is most affected by social stress, resulting in dramatically increased suicide rates. As many social observers and political commentators have pointed out, the labor market situation in the U.S. of white, lower educated-men is not good, especially relative to what it once meant to be a white male in the United States. These men are looking for jobs in an economy that is, from their perspective, increasingly restricted, with many jobs having been transferred to other countries and with increasing competition for the few jobs remaining due to the increased labor market participation of women of similar ages. This increasing competition has for many men led to lower-paid, lower-valued jobs, fewer opportunities for upward mobility, and increased subjective threats to social status. Further explanations may be found in the different socialization of the 1960-born cohort compared to earlier and later born cohorts, and in the specific impact of the economic crisis on *mid-age*, when possibilities to start over and to gain in income and wealth—or at least to *imagine* doing same—are lower compared to younger ages due to familial and other constraints. What is hopelessness, but the absence of hope?

Conclusion

Our study extends earlier findings on mortality and morbidity increases for white non-Hispanic mid-aged men with low education (Case & Deaton 2015) by providing a *robust test of cohort fluctuations* between 1990 and 2010 with micro-data on suicide rates, resulting in the demonstrations that, first, larger birth cohorts face increased risks of dying of suicide. Second, whites entering middle age in 2010 faced especially increased risk of dying of suicide if disadvantaged in terms of educational resources. One limitation of our study may be the use of five-year intervals when use of yearly intervals would be possible; however, the general patterns at issue here are well observable in the time intervals we chose. Further, despite rich micro-level indicators of age, gender, education, and race, research on suicide would largely benefit from information on religious affiliation (and religiosity) at the individual level, especially insofar as considerations of the relative social integration of whites and African-Americans is concerned. Another limitation, obviously not resolvable at this point in time, is that we cannot test or forecast how the 2010 increase for middle aged white men with low education will evolve over time. As this current "epidemic" of rising mortality and morbidity rates (Case & Deaton 2015) is a serious threat to individual well-being, and has some ugly implications for societal well-being as well, it will be important moving forward to keep an eye on the extent that the phenomenon is transitory (as cohort effects are) or an indicator of a more fundamental breakdown of the integration of specific classes of individuals within U.S. society.

Acknowledgments

This work was supported by the Fonds National de la Recherche Luxembourg (FNR/P11/05 and FNR/P11/05bis, PI: L. Chauvel) and by NICHD R24 HD-044964-11 (PI: H. Smith). Thanks in advance go to Julie Phillips for offering us her data to validate our results.

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Table 1. Coefficients with Individual-Level Controls for Men.

	Coeff.	Std. Err.	p	CI	
Cohort					
1930	0.048	0.017	0.005	0.015	0.081
1935	-0.059	0.016	0.000	-0.090	-0.029
1940	-0.134	0.014	0.000	-0.162	-0.106
1945	-0.042	0.012	0.000	-0.065	-0.019
1950	0.023	0.011	0.035	0.002	0.044
1955	0.098	0.010	0.000	0.078	0.118
1960	0.072	0.010	0.000	0.054	0.091
1965	0.088	0.009	0.000	0.070	0.106
1970	0.031	0.009	0.001	0.012	0.049
1975	0.052	0.010	0.000	0.032	0.072
1980	-0.053	0.012	0.000	-0.076	-0.030
1985	-0.124	0.013	0.000	-0.150	-0.099
Age					
20	-0.117	0.010	0.000	-0.135	-0.098
25	-0.045	0.009	0.000	-0.062	-0.028
30	0.013	0.009	0.135	-0.004	0.030
35	0.042	0.009	0.000	0.024	0.059
40	0.088	0.009	0.000	0.070	0.107
45	0.087	0.010	0.000	0.068	0.106
50	0.081	0.010	0.000	0.061	0.101
55	0.035	0.011	0.001	0.014	0.056
60	-0.032	0.011	0.004	-0.054	-0.010
65	-0.152	0.012	0.000	-0.176	-0.127
Period					
1990	0.021	0.005	0.000	0.011	0.030
1995	0.046	0.006	0.000	0.034	0.058
2000	-0.089	0.007	0.000	-0.102	-0.076
2005	-0.041	0.006	0.000	-0.053	-0.030
2010	0.064	0.004	0.000	0.055	0.073
Cohort (linear)	0.195	0.021	0.000	0.153	0.237
Age (linear)	0.197	0.011	0.000	0.175	0.220
High school grad	-0.018	0.009	0.048	-0.036	0.000
Some college	-0.751	0.011	0.000	-0.773	-0.730
College grad	-1.008	0.012	0.000	-1.031	-0.985
Married	-1.001	0.008	0.000	-1.018	-0.985
Widowed	0.387	0.021	0.000	0.347	0.428
Divorced	0.076	0.010	0.000	0.057	0.095
African-American	-0.801	0.013	0.000	-0.826	-0.776
Race: other	-0.633	0.019	0.000	-0.671	-0.596

Hispanic	-0.721	0.012	0.000	-0.744	-0.697
constant	-7.261	0.010	0.000	-7.280	-7.241

Note. Reference group for education is without completed high school; for marital status it is never married/single; and for race it is white.

Table 2. Coefficients with Individual-Level Controls for Women.

		Coeff.	Std. Err.	p	CI	
Cohort						
	1930	0.087	0.033	0.007	0.023	0.151
	1935	-0.010	0.029	0.723	-0.068	0.047
	1940	-0.146	0.027	0.000	-0.199	-0.093
	1945	-0.084	0.022	0.000	-0.128	-0.041
	1950	-0.007	0.020	0.733	-0.045	0.032
	1955	0.070	0.019	0.000	0.034	0.107
	1960	0.096	0.018	0.000	0.061	0.131
	1965	0.060	0.018	0.001	0.024	0.095
	1970	0.006	0.019	0.750	-0.032	0.044
	1975	0.054	0.022	0.014	0.011	0.096
	1980	-0.087	0.025	0.001	-0.137	-0.037
	1985	-0.038	0.028	0.177	-0.094	0.017
Age						
	20	-0.423	0.022	0.000	-0.466	-0.380
	25	-0.137	0.019	0.000	-0.175	-0.099
	30	0.093	0.018	0.000	0.057	0.129
	35	0.188	0.018	0.000	0.153	0.223
	40	0.270	0.017	0.000	0.235	0.304
	45	0.250	0.018	0.000	0.215	0.285
	50	0.231	0.019	0.000	0.194	0.267
	55	0.110	0.020	0.000	0.071	0.148
	60	-0.136	0.021	0.000	-0.178	-0.095
	65	-0.445	0.025	0.000	-0.493	-0.397
Period						
	1990	0.078	0.009	0.000	0.060	0.096
	1995	-0.020	0.012	0.096	-0.044	0.004
	2000	-0.106	0.013	0.000	-0.132	-0.081
	2005	-0.038	0.011	0.001	-0.060	-0.017
	2010	0.087	0.008	0.000	0.070	0.103
Cohort (linear)		0.485	0.041	0.000	0.404	0.565
Age (linear)		0.397	0.022	0.000	0.354	0.441

High school grad	0.100	0.021	0.000	0.059	0.140
Some college	-0.359	0.022	0.000	-0.403	-0.315
College grad	-0.535	0.024	0.000	-0.582	-0.489
Married	-1.004	0.018	0.000	-1.039	-0.969
Widowed	-0.150	0.030	0.000	-0.209	-0.090
Divorced	0.002	0.019	0.926	-0.036	0.039
African-American	-1.334	0.028	0.000	-1.388	-1.280
Race: other	-0.514	0.032	0.000	-0.578	-0.451
Hispanic	-1.026	0.027	0.000	-1.078	-0.973
constant	-8.834	0.023	0.000	-8.880	-8.789

Note. Reference group for education is without completed high school; for marital status it is never married/single; and for race it is white.