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Cohort Fluctuations of Mortality in Relation to Poverty, Education, Marriage and Cohort Size in the United States, 1989-2015

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Abstract

What are the aggregate social factors that track recent cohort-specific mortality trends? Using CDC multiple cause-of-death microdata and CPS data on populations cross-classified by age, sex, race/ethnicity, education, marital status, we create mortality rates for five-year age groups at five-year/yearly intervals. We examine mortality rates of U.S. birth cohorts between 1989-2015, net of aging and trends in historical time, via APC-D(eviations) modeling. These cohort deviations are correlated with cohort characteristics, e.g., relative poverty rates and cohort size. Cohort-specific poverty and mortality rates are particularly strongly associated, persisting net of cohort compositional factors. Education-stratified mortality patterns over time show sharp increases in mortality rates, significantly above linear time-trends, in low-educated white women of cohorts born >1960. No such patterns are found for the African-American men nor women. Findings suggest that recently occurring privations in the white U.S. population are small relative to the historically persistent mortality-associated privations of the African-American sub-population.

Background

Case and Deaton (2017) confirm their earlier (2015; cf. Gelman and Auerbach 2016) diagnosis concerning mortality in the U.S.: a strong singularity in mortality trends is observed among middle-aged less-than-university-educated white Americans of both genders. Mortality has ceased to improve for these folks, and for some it is getting worse.

Excesses in mortality are associated with social factors including: income and education (Cutler et al. 2006), poor lifestyle (smoking, diet, exercise [lack thereof]), lack of public control on communicable or uncommunicable epidemics, unsatisfactory market and public health services, etc. Some factors are more associated with individual characteristics and behaviors, and others with social environments, even if the distinctions are often blurred: one can be directly disadvantaged by his/her own poverty, or indirectly by the social stress of being in a category with high risks of poverty. These “social stressors” such as peaks of unemployment risks, strong modification in socioeconomic outlook, and more local, less predictable, transformations in circumstances and/or the environment, can contribute to the destabilization of health status.

For Case and Deaton (2017: 429), “cumulative disadvantages” explain why some birth cohorts born after the 1960s face *relative* scarcity (Chetty et al. 2017), hence frustrations with potentially lethal consequences. From a Durkheimian interpretation, this relative deprivation stressor is more consequential in the white population that for many years *benefited* from discriminations but which, post-circa-1970, now faces an unexpected gap between expectations and realities (Case and Deaton 2017: 430). Or thinks it does. One only need look to the former Soviet Union to see that the objectively better-off elements of a society can get into a negative behavioral spiral with strong implications for mortality when the world changes in a way that runs contrary to their assumptions and expectations (Guillot, Gavrilova, and Pudrovska 2011). Whereas in a long-term discriminatory environment, deprivation was (and remains) ever present: The expectations of minorities may never have been so great as to engender so consequential an increase in perceived deprivation.

If some descriptive cohort analysis are offered by Case and Deaton (2017), there is no strong attempt to isolate cohort singularities in the process (Cutler and Lleras-Muney 2017): Sometimes the young seniors born in the 1960 are at the forefront of the relative increase in mortality, sometimes it is the generations of adults born after 1980. A more systematic cohort analysis can better identify birth years with rapidly increasing risks as well as the better protected ones. These cohort fluctuations—deviations from the linear trend of mortality—can be compared to different cohort compositional characteristics including the “usual suspects,” such as education expansion, cohort size (the “Easterlin effect”) [1], poverty rates (relative poverty defined by level of living below 50% of the median), percentage of isolated singles in a cohort, etc.

We describe mortality trends between 1989 and 2015 for men and women, adjusting for individual sociodemographic characteristics. We take advantage of information on individual-level variables to

[1] This Easterlin (1980) effect is 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.

examine mortality patterns separately for gender, race, and educational level (university diploma and above versus completed high school and below), with a focus on the middle-aged white population (men and women) 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. Our aim here is to fill these gaps: we deepen the cohort-based diagnosis and analyze how the educational gradient of mortality is related to birth cohort dynamics.

Data

Data on mortality come from the Center for Disease Control and Prevention (CDC) Mortality Multiple Cause Files.[2] This annual micro dataset contains information annually on approximately two million decedents, including age, cause of death, and important socioeconomic descriptors. We consider the general mortality and a 9 category recode for one-year age groups for each year 1989-2015. We distinguish male and female mortality, not only because of well-known sex differences 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). The size of this decedents dataset is N≅56,000,000.

To construct a reference population at risk of death, we use the Current Population Survey (CPS) for the respective year.[3] For the one year corresponding to the CDC death data we obtain counts of the population for one-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 and consistent with respect to both data sources. The CPS reference population consists of somewhat less than five million observations, before application of extrapolation weights. We match the two micro datasets (deaths and population at risk) to compute mortality 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 can be subtracted from similar estimates of beginning-year populations at risk to create an aggregated cross-classification of micro data: dead (or not) × year × age × education × race/ethnicity × marital statistics. This allows computation of probabilities or odds. In practice, with the incidence of

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

mortality being relatively low, and intervals of observation short (a year), logs of rates, probabilities, and odds are all but indistinguishable from one another (Clogg and Eliason 1987, pp. 28-34).

Method

Consider a generalized equation of the form

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

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)[4]. Our analysis under this model emphasizes cohort contrasts where more advantaged birth cohort are systematically above the linear age-

[4] The APCD can be downloaded as a Stata ado file by typing "ssc install apcd" in Stata.

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[5].

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 mortality 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 mortality) or whether they represent more "intrinsic" cohort characteristics, such as the (mis)fortune represented by (relatively larger) cohort size, pervasiveness of poverty, extent of social isolation, etc.

Results

From a descriptive point of view, Figure 0, it is well known mortality is higher with age with a close to linear relation of age and logged mortality rates (the standard Gompertz–Makeham law). In the same age groups, men's' mortality is higher than women's, African Americans' than Whites', and those of the less educated than those of holders of diplomas. The change over time is supposed to go in the direction of a reduction of mortality. In Figure 0, we should expect the red line (2015) to be lower than the black one (1990). There are exceptions: At young adult ages educated men and women show a small increase in mortality over time; things seem in general to have gotten worse for less educated white males; and except for the oldest less educated white women, mortality has increased dramatically.

[5] In terms of methods, we have developed our general linear model (glm) based apcd on the base of a critique of hierarchic hapc (see Chauvel & Schroeder 2015) because the hapc cannot deliver robust detrended results. The Stata constrained glm model is much more reliable in this purpose. But now we need improvements in some multilevel APCD models able to detect in a proper single run estimation the coefficients for individual level variables (like education of individuals) and for contextual cohort variables like poverty, general level of education, or cohort size. Newer Stata constrained multilevel glm models could provide the base for this methodological improvement but the existing algorithm have to gain in rapidity and reliability since hours or days of constrained multilevel glm deliver more problematic results than few minutes with the old cglm.

Figure 0. Log of Mortality by Age for Eight Specific Subgroups of the Population (Thick Black Line=1990; Red=2015).

Figure 0 shows some fluctuations of small intensity compared to the Gompertz law—for example, cohorts born at the end of the 1950s did worse than the others—but these are plots of mortality by age across different years, not within cohorts. A more systematic analysis of cohort specific mortality is offered by the APCD model. APCD Poisson models were run without control variates, then with education, marital status and race as covariates for men and women (Annex Table 1).

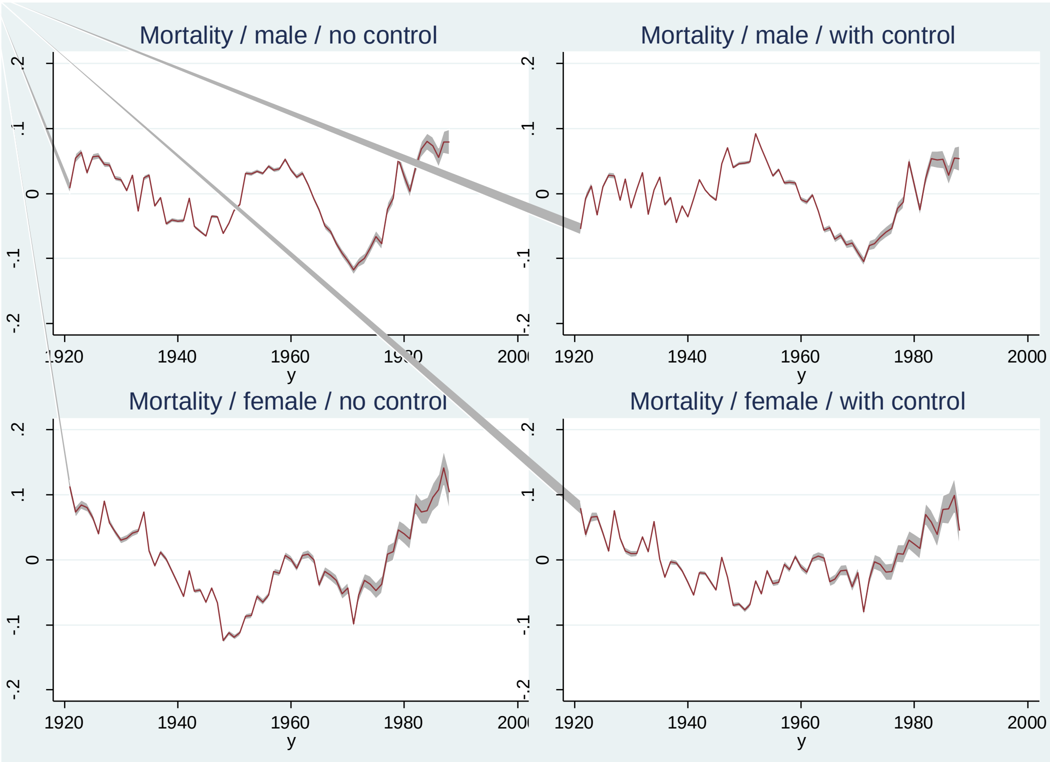

Figure 1. Cohort Fluctuation (APC-Detrended) for Mortality by Birth Cohort with and without Controls, Men and Women.

Note. American resident population, controls are education (BA & + owners, No BA), race (White, African-American, Other), marital status (married, not married); 95% confidence intervals are greyed.

Figure 1 shows the (de-trended) cohort fluctuations in mortality rates for both men and women, without control variates (left hand side) and with control variates (right hand side). For men, the 1950 cohort has a -5 % reduced risk, whereas the 1960 cohort faces a +5 % increased risk of death, relative to the linear trend in age and cohort. The control variables (see Table 1 in Annex) confirm what we know from Durkheim (1897) and his followers about suicide and also mortality: marriage is protective, all the more so for males. University education is a protection of the same magnitude of intensity as marriage. African American category is associated with a 27% higher risk of mortality.

Education has a protective role (Zajacova 2006), with a significantly stronger gradient for the male population (Montez et al. 2009). 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, which are deviations from the linear trend. Adjusting for micro level control characteristics does make a difference. Among males, the pattern looks similar, but note that the peak with controls is displaced further back into the 1950s. Among women, there is no similar displacement, but there is a substantial diminution in cohort variation. For both men and women, the more recent cohorts—those born after 1975 or 1980—face a new upsurge in mortality.

Plotting the Age-Period-Cohort-Detrended (APCD) coefficients for cohort deviations from trend in rates of mortality against the APCD coefficient of other macro contextual variables :

- Relative size of birth cohort (variable: popsize)[6]
- Relative exposure of cohort to poverty (variable: pov)
- Relative level of education of the cohort (variable: educboom)
- Relative risk to live in a single household (variable: alo)

allows a better understanding of mortality rates fluctuations per cohort detected in APCD models for both genders (Figures 2a and 2b).

The APCD cohort coefficients for mortality (white males fig 2a) show complicated variations with two successive V shaped fluctuations where cohort born in 1947 and in 1971 had been minima in mortality, relative to the linear trend of health improvement by age and period. The APCD cohort coefficients for poverty rates present a somewhat equivalent shape. Cohorts born in the 1950s, and then in the late 1980s, had higher levels of celibacy. The APCD cohort coefficients show the strong acceleration of access to education for cohorts born near 1950 (Card and Lemieux, 2001), and a second minor acceleration of educational expansion benefitted to cohort 1970. Finally, cohort size shows the baby-boom of the period 1950-1970.

[6] 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.

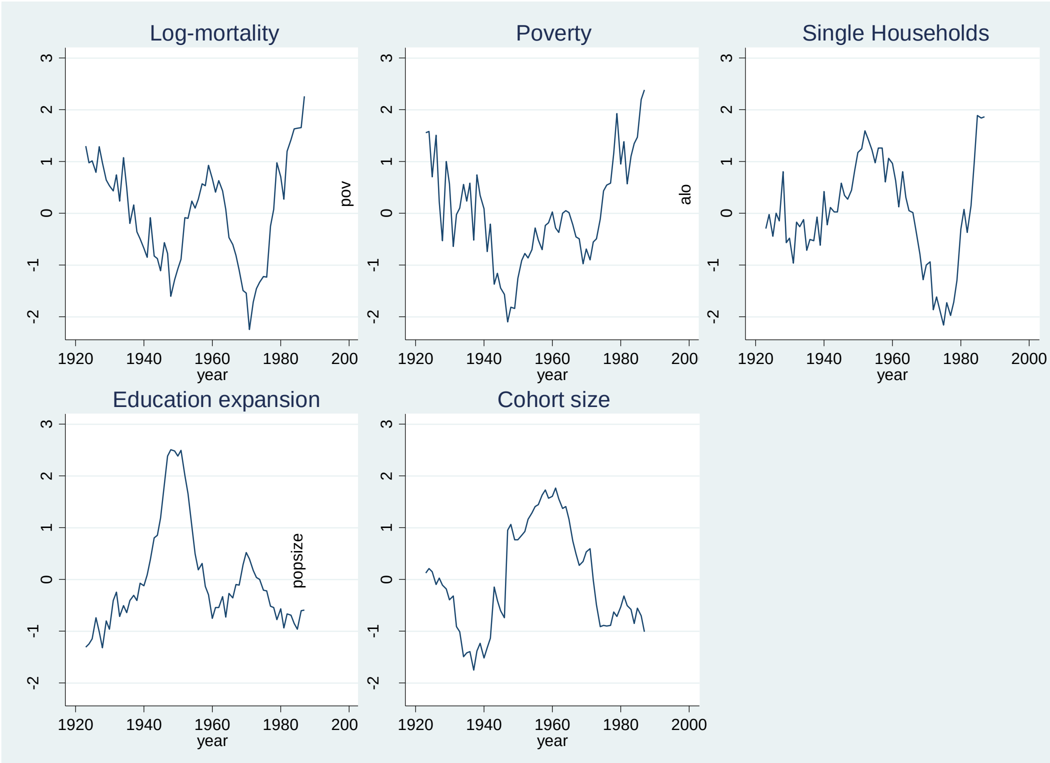

Figure 2a. Cohort Fluctuation (APC-Detrended, Standardized) Profiles for Logged Mortality, Poverty Rate, Share of Single Households, Education (BA+) and Logged Cohort Size, by Birth Cohort for White Men.

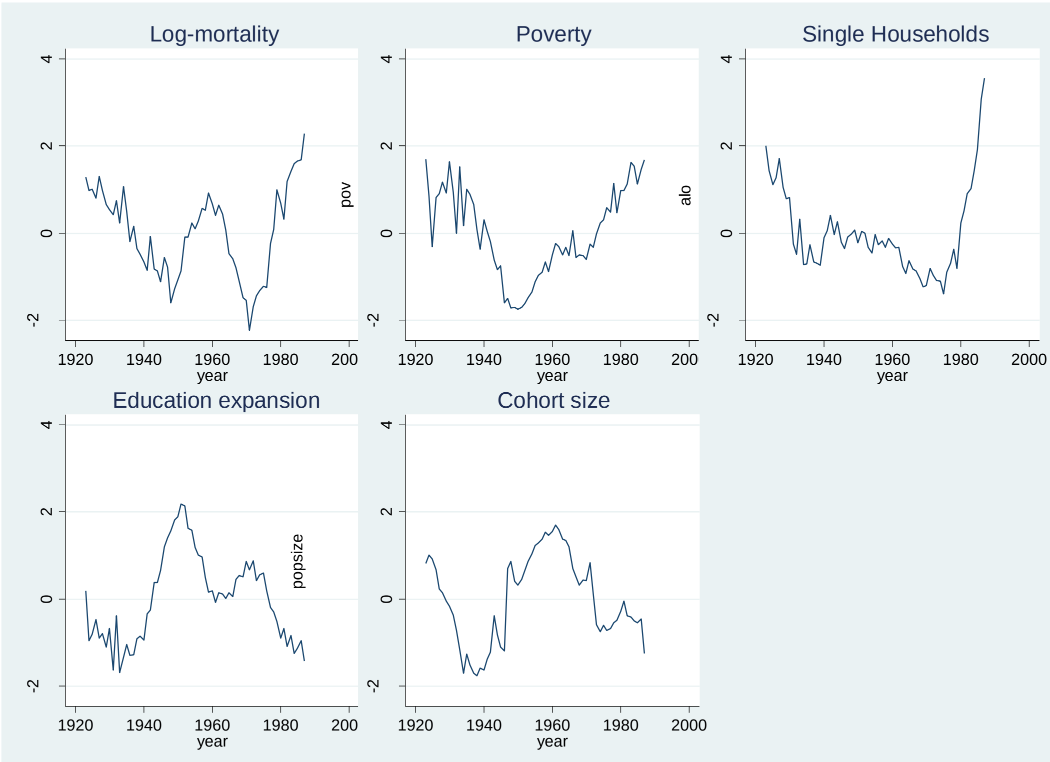

Figure 2b. Cohort Fluctuation (APC-Detrended, Standardized) Profiles for Logged Mortality, Poverty Rate, Share of Single Households, Education (BA+) and Logged Cohort Size, by Birth Cohort for White Women.

For the white female population, the fluctuations are broadly similar, with a higher acceleration of education for the median baby-boomers, and a strong development of the non-married population in the cohorts born after 1980.

We correlate these cohort coefficients with those previously obtained for the APCD model for mortality rates (Annex Table 1). The R^2 of this correlation suggests that, for men, a total of 47 % of the variance in mortality rates can be explained by the APCD coefficients of poverty rates 35 % by education expansion (upper left panel of Figure 3a), and 17% by non-marriage. For women, associations of birth cohort size and mortality fluctuation are respectively 31, 34 and 49%, smaller for education and poverty, higher for non-marriage, but size of birth cohort contributes nothing to the variance in mortality fluctuations ($R^2 = 0$ %; lower right panel of Figures 3). APCD Models for mortality were also run including individual-level variables. Once the APCD of mortality is net of individual education, marriage, race, we observe no role of contextual (aggregated) level of poverty or education, but the APCD coefficients of the rate of non-married individuals still have a significant relation with mortality, for both male and female population (lower left panel of Figure 3b and 3d). This means that the gross correlation between cohort education and cohort mortality come entirely from individual resources, not of cohort context; the same for poverty. This is to say that “lucky” generations in terms of education are “lucky” in terms of health because of better individual level resources. By the contrary, non-marriage at the contextual level of cohorts still have a role, even after the control by individuals' marital status.

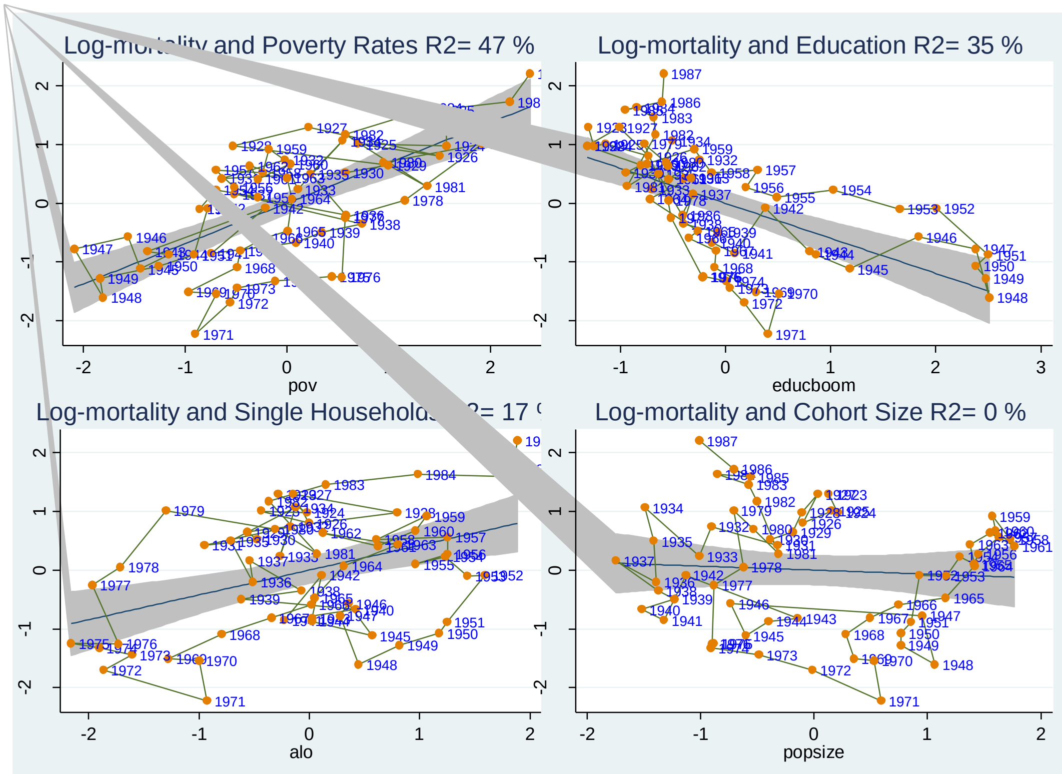

Figure 3a. Associations of APCD Log-mortality (Vertical) with Four Explanatory Factors (APCD Coefficients of Poverty Rate, Share of Single Households, Education (BA+) and Logged Cohort Size) for White Men (No Controls).

Figure 3b. Associations of APCD Log-mortality (Vertical) with Four Explanatory factors (APCD Coefficients of Poverty Rate, Share of Single Households, Education (BA+) and Logged Cohort Size) for White Men (with Controls).

Figure 3c. Associations of APCD Log-mortality (Vertical) with Four Explanatory Factors (APCD Coefficients of Poverty Rate, Share of Single Households, Education (BA+) and Logged Cohort Size) for White Women (no Controls).

Figure 3d. Associations of APCD Log-mortality (Vertical) with Four explanatory Factors (APCD Coefficients of Poverty Rate, Share of Single Households, Education (BA+) and Logged Cohort Size) for White Women (with Controls).

Table 2a. Correlation (R) between Cohort Context Variables and Mortality for White Men (with Controls).

	popsize	educboom	pov	alo	logmorta
popsize	1.0000				
educboom	0.2998*	1.0000			
pov	-0.3864*	-0.7674*	1.0000		
alo	0.4197*	0.2667*	-0.0662	1.0000	
logmorta	0.0093	0.0216	0.2701*	0.6000*	1.0000

Table 2b. Correlation (R) between Cohort Context Variables and Mortality for White Women (with Controls).

	popsize	educboom	pov	alo	logmorta
popsize	1.0000				
educboom	0.4965*	1.0000			
pov	-0.4113*	-0.8560*	1.0000		
alo	-0.0437	-0.3904*	0.4991*	1.0000	
logmorta	0.0004	-0.1330	0.1090	0.6212*	1.0000

Finally, consider the relative contribution of size of birth cohort (*It's the demography, stupid!* Easterlin, 1987), educational expansion (*It's the education, stupid!* Dowd & Hamoudi, 2014), poverty rates (*It's the economy, stupid!*) to explaining fluctuations in mortality rates. Tables 2a and 2b give the crude associations between the cohort indicators.

The regression analysis of the male and female APCD fluctuations (without and with controls, Table 3) shows the role of non-marriage rate in a generation to understand the variations of mortality, unexplained by individual characteristics. Education has a strong role before control by individual variables, which disappears after taking individual variables into account.

Table 3. Linear Regression of the APCD Parameters of Log-mortality (with Controls) and APCD of Poverty Rate, Share of Single Households, Population Size and Education.

	Men, no controls	Men, with controls	Women, no controls	Women, with controls
Regression coefficients				
Educboom	-0.478***	0.227	-0.648***	-0.223
Alo	0.593***	0.636***	0.521***	0.778***
Popsize	-0.045	-0.161	0.335***	-0.058
Pov	0.343**	0.428**	-0.111	-0.494*
Fit of Regression Model				
R^2	0.775	0.502	0.699	0.458
Adj R^2	0.760	0.469	0.679	0.422

Note. *p<0.05. **p<0.01. ***p<0.001. Educboom - Relative level of education of the cohort; Alo - Relative risk to live in single household; Popsize - Relative size of birth cohort; Pov - Relative exposure of cohort to poverty.

Conclusion

Our analysis demonstrates that, before control by individuals' education, non-marriage and race, the cohort variations of mortality are highly correlated with cohort variation of education expansion, poverty rate, non-marriage (but not cohort size), at least with data of the United States at the time between 1989 and 2015. As the overall explained variance suggests, size of birth cohort is actually not the main driver of mortality rates. Once individual variables are controlled, non-marriage rate of a cohort is the last contextual variable still related with mortality. This has been demonstrated for general mortality. One limitation is that we work on general mortality and not by specific causes of death, but for the final full paper, we will compare ten different causes of death. This could help in the contrast between avoidable and non-avoidable causes, and in stress-based or lack of care causes.

Our contribution shows the important role birth cohort in the explanation of mortality: contemporary acceleration of mortality of the young seniors (close to 60 years old) is indeed related to a cohort effect of a cohort born near to 1955-1965, that experienced public health problems in their young adulthood in the context of AIDS/HIV, crack epidemic, etc. This means a specific cohort with intrinsic frailties accumulated over life course, from early adulthood. The public health problem analyzed in terms of birth cohort by Case and Deaton (2017) has found here its correct APCD identification as a cohort non-linearity.

Our method helped in the detection of individual protecting factors: higher educated, married, white women benefit from lower risk of early death. In this respect education is of particular importance in terms of birth cohorts: some cohorts benefitted from accelerations in educational investment (the cohort born near to 1948) and others faced scarcer and more expensive opportunities in universities. Those ones faced a relative progression of mortality. Analysis of mortality patterns over time for high- and low-educated men and women showed that white women with low education face a sharp increase, significantly above the linear time trends, in mortality rates among cohorts born after 1960. No such pattern is found for African-Americans, whether high- or low-educated, for both males and females. This accords with the observation that whatever privations have been occurring in the white U.S. population, the extent of change is small relative to the base privations associated with mortality that have historically afflicted the African-American sub-population. The distinct mortality patterns of Blacks and Whites we find join a growing number of studies on the distinct life experiences of Blacks and Whites (see e.g. for self-rated health, Assari, Lankarani & Burgard 2016).

Another cohort issue that our paper underlines is the fantastic acceleration of mortality below age 30 for the cohorts born after 1980. This new set of cohorts must be taken into consideration in the future: is it a temporary situation or the new coming of age of a frail generation? The story of cohort analysis is that the Cassandrae are recognized when it is too late.

Acknowledgments

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Annex Table 1. Coefficients with Individual-Level Controls for Men and Women.

Variable	Value	male/no control	male/w control	female/no control	female/w control
Education	>=BA		-0.879***		-0.771***
Marital status	married		-0.814***		-0.591***
Ethnicity	Af.Amer		0.270***		0.275***
Ethnicity	Other		-0.357***		-0.486***
age	25	0.153***	0.317***	0.0591***	0.193***
age	26	0.130***	0.259***	0.0461***	0.142***
age	27	0.146***	0.234***	0.0350***	0.109***
age	28	0.0937***	0.156***	-0.0244***	0.0249***
age	29	0.106***	0.133***	-0.0478***	-0.00902
age	30	0.0560***	0.0744***	0.0108*	0.0326***
age	31	0.0520***	0.0472***	-0.0225***	-0.0230***
age	32	0.0334***	0.0130***	-0.0135**	-0.0284***
age	33	0.0210***	-0.00332	-0.0206***	-0.0375***
age	34	0.0113***	-0.0286***	-0.0390***	-0.0563***
age	35	-0.0600***	-0.0931***	-0.0417***	-0.0625***
age	36	-0.0632***	-0.103***	-0.0235***	-0.0568***
age	37	-0.0584***	-0.106***	-0.0188***	-0.0541***
age	38	-0.0739***	-0.124***	-0.0217***	-0.0545***
age	39	-0.0740***	-0.130***	-0.0263***	-0.0577***
age	40	-0.118***	-0.163***	-0.0290***	-0.0663***
age	41	-0.0964***	-0.141***	-0.0173***	-0.0539***
age	42	-0.0990***	-0.141***	-0.00869**	-0.0433***
age	43	-0.0989***	-0.144***	0.00962**	-0.0313***
age	44	-0.0847***	-0.135***	0.00397	-0.0323***
age	45	-0.0907***	-0.138***	0.000165	-0.0388***
age	46	-0.0800***	-0.118***	-0.0052	-0.0360***
age	47	-0.0773***	-0.107***	0.0124***	-0.0201***
age	48	-0.0697***	-0.103***	0.0249***	-0.00835**
age	49	-0.0706***	-0.0897***	0.0344***	0.000609
age	50	-0.0898***	-0.105***	0.00772**	-0.0235***
age	51	-0.0338***	-0.0600***	0.0361***	0.00398
age	52	-0.0219***	-0.0392***	0.0440***	0.0232***
age	53	-0.00469**	-0.0268***	0.0391***	0.0160***
age	54	0.00655***	-0.0198***	0.0279***	0.00585**
age	55	-0.0228***	-0.0251***	0.0191***	0.00118
age	56	-0.0126***	-0.00660***	0.0273***	0.00723***
age	57	0.0203***	0.0186***	0.0330***	0.0123***
age	58	0.0194***	0.0261***	0.0209***	0.0179***
age	59	0.0265***	0.0351***	0.0239***	0.0247***

age	60	0.0302***	0.0330***	0.00328	0.00553**
age	61	0.0427***	0.0481***	0.0183***	0.0240***
age	62	0.0423***	0.0678***	0.0174***	0.0312***
age	63	0.0658***	0.0887***	0.00862***	0.0246***
age	64	0.0665***	0.0947***	-0.00561***	0.0232***
age	65	0.00362**	0.0413***	-0.0543***	-0.00650***
age	66	0.0298***	0.0723***	-0.0307***	0.0199***
age	67	0.0646***	0.111***	-0.0240***	0.0271***
age	68	0.0892***	0.137***	-0.0405***	0.0143***
age	69	0.0917***	0.141***	-0.0479***	0.0163***
per	1989	-0.0266***	-0.0117***	-0.0397***	-0.0303***
per	1990	-0.0252***	-0.0142***	-0.0413***	-0.0348***
per	1991	-0.0285***	-0.0128***	-0.0363***	-0.0290***
per	1992	-0.0157***	-0.00793***	-0.0342***	-0.0323***
per	1993	0.0219***	0.0287***	-0.00171	-0.00377*
per	1994	0.0515***	0.0478***	0.0298***	0.0164***
per	1995	0.0578***	0.0470***	0.0387***	0.0280***
per	1996	0.0203***	0.0160***	0.0266***	0.0213***
per	1997	-0.0129***	-0.0125***	0.00344	0.0122***
per	1998	-0.0130***	-0.0199***	0.00459*	0.0111***
per	1999	-0.000909	-0.0111***	0.0225***	0.0229***
per	2000	0.00291	-0.0106***	0.0355***	0.0293***
per	2001	-0.00199	-0.0203***	0.0292***	0.0182***
per	2002	0.000579	-0.0105***	0.0189***	0.0155***
per	2003	0.0182***	0.00541***	0.0214***	0.0245***
per	2004	0.0012	-0.00799***	0.0108***	0.00992***
per	2005	-0.00367*	-0.00187	0.0154***	0.0169***
per	2006	-0.00876***	-0.00223	0.00912***	0.00667***
per	2007	-0.00879***	-0.0105***	-0.00864***	-0.00629***
per	2008	-0.00297*	-0.00683***	-0.00034	-0.000967
per	2009	-0.00895***	-0.0141***	0.00237	-0.00482**
per	2010	-0.0366***	-0.0242***	-0.0284***	-0.0284***
per	2011	-0.0167***	-0.00871***	-0.0288***	-0.0257***
per	2012	-0.0161***	-0.00571***	-0.0303***	-0.0241***
per	2013	0.00981***	0.00958***	-0.0191***	-0.0152***
per	2014	0.0115***	0.0187***	0.000556	-0.00164
per	2015	0.0315***	0.0403***	0.0000131	0.00429**
rescacoh	-0.446***	-0.604***	0.0225***	-0.296***	
rescaage	1.539***	1.223***	1.683***	1.618***	
coh	1921	-0.0551***	0.00934*	0.0805***	0.112***
coh	1922	-0.00652	0.0543***	0.0392***	0.0740***
coh	1923	0.0118***	0.0637***	0.0655***	0.0845***
coh	1924	-0.0328***	0.0325***	0.0667***	0.0810***
coh	1925	0.0107***	0.0563***	0.0409***	0.0649***
coh	1926	0.0286***	0.0580***	0.0136***	0.0399***

coh	1927	0.0269***	0.0453***	0.0757***	0.0896***
coh	1928	-0.0100***	0.0448***	0.0327***	0.0572***
coh	1929	0.0229***	0.0234***	0.0137***	0.0435***
coh	1930	-0.0217***	0.0216***	0.00913***	0.0304***
coh	1931	0.00709***	0.00495*	0.00958***	0.0338***
coh	1932	0.0317***	0.0281***	0.0355***	0.0408***
coh	1933	-0.0316***	-0.0266***	0.0121***	0.0444***
coh	1934	0.00588**	0.0236***	0.0588***	0.0739***
coh	1935	0.0250***	0.0284***	0.000878	0.0146***
coh	1936	-0.0176***	-0.0193***	-0.0267***	-0.00863***
coh	1937	-0.00557**	-0.00648**	-0.00315	0.0118***
coh	1938	-0.0445***	-0.0466***	-0.00461*	0.000727
coh	1939	-0.0191***	-0.0410***	-0.0164***	-0.0181***
coh	1940	-0.0361***	-0.0421***	-0.0345***	-0.0368***
coh	1941	-0.00816***	-0.0418***	-0.0539***	-0.0561***
coh	1942	0.0218***	-0.00707***	-0.0198***	-0.0165***
coh	1943	0.00564***	-0.0506***	-0.0206***	-0.0483***
coh	1944	-0.00317	-0.0580***	-0.0330***	-0.0466***
coh	1945	-0.00996***	-0.0653***	-0.0463***	-0.0654***
coh	1946	0.0461***	-0.0354***	0.00353	-0.0436***
coh	1947	0.0702***	-0.0356***	-0.0269***	-0.0663***
coh	1948	0.0404***	-0.0610***	-0.0696***	-0.124***
coh	1949	0.0463***	-0.0451***	-0.0680***	-0.113***
coh	1950	0.0471***	-0.0258***	-0.0771***	-0.119***
coh	1951	0.0485***	-0.0169***	-0.0689***	-0.112***
coh	1952	0.0917***	0.0320***	-0.0330***	-0.0871***
coh	1953	0.0688***	0.0301***	-0.0522***	-0.0855***
coh	1954	0.0487***	0.0346***	-0.0173***	-0.0567***
coh	1955	0.0273***	0.0308***	-0.0362***	-0.0647***
coh	1956	0.0373***	0.0416***	-0.0342***	-0.0544***
coh	1957	0.0167***	0.0355***	-0.00664*	-0.0181***
coh	1958	0.0179***	0.0381***	-0.0151***	-0.0206***
coh	1959	0.0158***	0.0526***	0.00525	0.00663*
coh	1960	-0.00882***	0.0361***	-0.0107***	0.000455
coh	1961	-0.0131***	0.0258***	-0.0190***	-0.0125***
coh	1962	-0.00179	0.0318***	0.00202	0.00635*
coh	1963	-0.0269***	0.0136***	0.00594	0.00819*
coh	1964	-0.0561***	-0.00791**	0.00334	-0.000692
coh	1965	-0.0527***	-0.0253***	-0.0337***	-0.0384***
coh	1966	-0.0699***	-0.0497***	-0.0295***	-0.0181***
coh	1967	-0.0643***	-0.0579***	-0.0165***	-0.0250***
coh	1968	-0.0787***	-0.0768***	-0.0151***	-0.0317***
coh	1969	-0.0760***	-0.0918***	-0.0413***	-0.0521***
coh	1970	-0.0902***	-0.104***	-0.0200***	-0.0439***
coh	1971	-0.104***	-0.117***	-0.0796***	-0.0985***

coh	1972	-0.0795***	-0.106***	-0.0285***	-0.0539***
coh	1973	-0.0774***	-0.100***	-0.00269	-0.0322***
coh	1974	-0.0667***	-0.0841***	-0.00635	-0.0374***
coh	1975	-0.0593***	-0.0656***	-0.0179**	-0.0478***
coh	1976	-0.0532***	-0.0771***	-0.0176**	-0.0383***
coh	1977	-0.0223***	-0.0238***	0.0101	0.00879
coh	1978	-0.0132**	-0.00724	0.00984	0.0127
coh	1979	0.0486***	0.0540***	0.0306***	0.0458***
coh	1980	0.0135**	0.0274***	0.0247**	0.0401***
coh	1981	-0.0250***	0.0041	0.0177*	0.0318***
coh	1982	0.0230***	0.0406***	0.0704***	0.0873***
coh	1983	0.0533***	0.0681***	0.0587***	0.0724***
coh	1984	0.0528***	0.0803***	0.0373***	0.0754***
coh	1985	0.0539***	0.0730***	0.0777***	0.0967***
coh	1986	0.0279***	0.0560***	0.0794***	0.109***
coh	1987	0.0559***	0.0792***	0.0999***	0.142***
coh	1988	0.0526***	0.0795***	0.0423**	0.105***
coh	1989	0.139***	0.160***	-0.0304	0.0481*
	cons	-4.451***	-5.283***	-5.219***	-5.892***

* p<0.05, ** p<0.01, *** p<0.001