

# Social Inequalities in Happiness in the United States, 1972 to 2004: An Age-Period-Cohort Analysis

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*This study conducts a systematic age, period, and cohort analysis that provides new evidence of the dynamics of, and heterogeneity in, subjective well-being across the life course and over time in the United States. I use recently developed methodologies of hierarchical age-period-cohort models, and the longest available population data series on happiness from the General Social Survey, 1972 to 2004. I find distinct life-course patterns, time trends, and birth cohort changes in happiness. The age effects are strong and indicate increases in happiness over the life course. Period effects show first decreasing and then increasing trends in happiness. Baby-boomer cohorts report lower levels of happiness, suggesting the influence of early life conditions and formative experiences. I also find substantial life-course and period variations in social disparities in happiness. The results show convergences in sex, race, and educational gaps in happiness with age, which can largely be attributed to differential exposure to various social conditions important to happiness, such as marital status and health. Sex and race inequalities in happiness declined in the long term over the past 30 years. During the most recent decade, however, the net sex difference disappeared while the racial gap in happiness remained substantial.*

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As the increase in Americans' life expectancy continues into the twenty-first century, there is a growing need for research and policy to examine both the quantity and the *quality* of life. A fundamentally important question for research communities, policy makers, and public authorities is: Are Americans living better as well as longer lives? Just as with objective life

conditions, subjective perceptions of well-being are increasingly adopted as useful social indicators to assess quality of life in the overall population and in population subgroups (George 2006). Measurements of quality of life in terms of subjective well-being are also useful for determining the extent to which societies meet the needs of their members and the degree to which citizens thrive (Veenhoven 1992). Previous research defines subjective well-being as a state of stable, global judgment of life quality and the degree to which people evaluate the overall quality of their present lives positively (Diener 1984). General happiness is the measure of subjective well-being examined most frequently (George 2006; Veenhoven 1992).

How do Americans fare in happiness? Previous research suggests that different demographic and socioeconomic groups do not fare equally well (e.g., Davis 1984; Easterlin 2001). Analyses of correlates of happiness, however, are mostly concerned with cross-sectional individual-level characteristics and attainments.

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We know relatively little beyond the stratification of subjective quality of life at a static point in time. Two additional limitations further prevent our thorough understanding of the dynamics of subjective well-being across the life course and over time. First, there is no simultaneous assessment of the age, time period, and birth cohort effects. These temporal sources of variations in subjective well-being need to be distinguished because they are crucial for attributions of explanatory mechanisms and generalizability of research findings. Second, there is little sociological understanding of the heterogeneity in life-course patterns, time trends, and birth cohort differences. The possibility that the age, period, and cohort effects may be modified by social status warrants a systematic investigation.

This study aims to address these limitations by using the longest time-series data available on happiness in the United States (spanning more than 30 years) from the General Social Survey (GSS), as well as recently developed methodologies of hierarchical age-period-cohort models (Yang and Land 2006, forthcoming). Specifically, this study assesses the following questions: What are the net age, period, and cohort effects on happiness? Do findings of life-course patterns and time trends hold when all three types of variation are taken into account? To what extent do life-course patterns and time trends actually reflect birth cohort differences? Furthermore, do social inequalities in happiness by sex, race, and education increase or decrease over the life course, over time, and across successive birth cohorts? This is the first study of subjective quality of life to conduct a systematic age, period, and cohort analysis, which allows us to see the various ways in which social stratification operates over the life course and historical time. This also enables us to test the extent to which aging and life-course theories apply to subjective well-being, and the results shed new light on the changes in the social distribution of quality of life in the United States.

### AGE, PERIOD, AND COHORT EFFECTS

Despite voluminous studies on correlates of subjective well-being across disciplines and countries, relatively little is known about

changes and social disparities in subjective well-being. Even less is known about the unique contributions of three types of time-related variations—age, period, and cohort effects—to these changes, and distinguishing these effects is crucial for attributions of individual or social mechanisms that produce inequalities in subjective well-being. Age effects represent aging-related developmental changes in the life course, whereas temporal trends across time periods or birth cohorts reflect exogenous contextual changes in broader social conditions. Period and cohort effects can also be distinguished from each other. Period effects occur due to cultural and economic changes that are unique to time periods, and they induce similar changes in well-being for individuals of all ages. Cohort effects are the essence of social change and represent the effects of formative experiences (Ryder 1965). They subsume the effects of early life conditions and the continuous exposure to historical and social factors that affect subjective well-being throughout the life course. This distinction can also be important for the generalizability of findings. For instance, in the absence of period and cohort effects, age changes in happiness are broadly applicable across individuals in different time periods and cohorts. But differences in periods and cohorts are indicative of the existence of period- and cohort-specific social forces that affect changes in happiness.

Extant research clearly indicates variations in happiness by age, historical time, or birth cohort. Most prior studies, however, examine these effects in isolation, based on the assumption that the effects omitted from the analyses are null. Findings are inconsistent regarding the directions and forms of these effects and not informative regarding their relative importance and confounding effects.

Do people become happier as they move through the life course? The increasing health problems and loss of important social relationships through mortality with increasing age lead to predictions of a decrease in quality of life over the life course (George 2006). The “role theory” of the aging process, on the other hand, suggests that self-integration, insight, and positive psychosocial traits such as satisfaction and self-esteem grow with age and that these signs of maturity in turn increase quality of life with age (Gove, Ortega, and Style 1989). Findings

from empirical studies are mixed. Age effects of happiness have been reported as negative (Rodgers 1982), positive (Charles, Reynolds, and Gatz 2001), constant (Costa et al. 1987), and U-shaped, with a minimum level of happiness occurring between the ages of 30 and 40 (Mroczek and Kolarz 1998).

The lack of representative samples, differences in measurements, age compositions of the samples, survey years, and adjustments of different sets of covariates may account for the inconsistency in findings on age differences in happiness. Inadequate research designs are also a key problem. Most relevant findings on life-course changes are based on one-point-in-time cross-sectional observations classified by age. Such designs ignore different age changes in happiness levels by historical time (Rodgers 1982) and confound age and cohort effects (Mason and Fienberg 1985). Even in extant longitudinal studies, age and cohort effects have not been explicitly distinguished and simultaneously estimated. Nevertheless, most researchers believe that, although historical and cohort changes could affect the association of age with happiness, the age effect is a basic and robust trend (Diener and Suh 1997; Shmotkin 1990). It may be true that the age effects of happiness are not simply period and cohort effects, but it must remain a possibility until formally tested.

Did overall levels of subjective well-being increase or decrease over the past few decades? The past 30 years can be characterized as a period of economic prosperity. Different theoretical perspectives on the relationship between economic growth and happiness predict different time trends of happiness. The social comparison, or reference group, hypothesis suggests that income affects happiness through ranks, not score levels, and higher relative status increases happiness (Davis 1984; Hagerty 2000). The adaptation hypothesis, in contrast, suggests that changes in income should relate to short-term changes but not long-term values in happiness (Davis 1984; Hagerty and Veenhoven 2003). Both of these relative-utility theories (the first relative to others; the second relative to one's past) predict no increases in happiness over time related to economic growth (Hagerty and Veenhoven 2003). In contrast, the theory of absolute utility states that income growth can fill more needs and increase happi-

ness, which suggests an increasing trend in happiness over time (Veenhoven 2005; Veenhoven and Hagerty 2006).

Existing studies of time trends in happiness generally support a lack of increase over time, despite ever-increasing economic prosperity (Easterlin 1995). Earlier analyses of U.S. surveys conducted between the 1940s and 1970s report significant but weak temporal changes in happiness over time (Davis 1984; Rodgers 1982; Smith 1979). A recent analysis of GSS data from 1972 to 1998 shows that Americans are becoming less happy over time (Blanchflower and Oswald 2004). The effects are constrained to be linear, very small in magnitude, and statistically insignificant when both age and other individual covariates are held constant.

Different conclusions can be drawn for the life-course changes in happiness in different calendar years. Similar problems can occur for the inference of period effects when studies of time trends are conducted without the consideration of age group differentials. Studies have found that older adults were less happy than younger adults before the 1970s, but they were happier in more recent years because younger adults experienced declines in happiness that adults over age 65 did not (Rodgers 1982). The differential time trends by age groups have been hypothesized to reflect cohort differences (Easterlin 2001; Rodgers 1982). Because birth cohort has explanatory power as a structural category and represents an important source of social variation in life-course outcomes (Ryder 1965), analyses of dynamics of happiness should formally test cohort effects.

To date, no study has explicitly examined birth cohort effects in a multivariate analysis of happiness. Several hypotheses suggest different directions for cohort changes. The "post-materialism" hypothesis predicts a smaller increase in happiness among more recent cohorts. This is because post-materialism induced a shift in values in the 1960s so that more recent cohorts, with more education and income, have become less concerned with basic survival needs and more concerned with higher-level needs such as political, ecological, and human relations problems (Rodgers 1982). On the other hand, Ryder's (1965) classic cohort analysis proposition emphasizes that individuals are particularly impressionable early in the life course. The difficult childhood and young

adulthood experiences associated with the Depression and the world wars of earlier cohorts may lead these cohorts to lower levels of general well-being (Elder 1974). In addition, it is important to consider the impact of the exogenous social-demographic environment into which cohorts were born and in which they came of age, as suggested by the Easterlin hypothesis that relates birth with fortune (Easterlin 1987). A large cohort creates more competition for schooling and jobs, leading to negative consequences for socioeconomic achievement and psychological well-being. The baby boomers, therefore, may exhibit lower levels of happiness than preceding and successive cohorts.

### **SOCIAL INEQUALITIES IN HAPPINESS: AGE AND TIME VARIATIONS**

Sociological analyses have long established the importance of social positioning to the distribution of happiness. The stratification hypothesis states that a higher rank along any evaluated dimension should produce greater happiness (Davis 1984). Empirical evidence supports this hypothesis and suggests that happiness is moderately stratified by sex and strongly stratified by race and socioeconomic status (SES). Women are, on average, happier and more content than men, with or without controls for variables in which women are underprivileged (Easterlin 2001; Shmotkin 1990). The relationship between gender and happiness is in the opposite direction than would be expected on the basis of stratification theory, but gender differences in happiness are generally small. The black-white difference is much larger than the gender difference, with blacks being less happy than whites, and this difference cannot be explained by SES or perceived racial discrimination (Davis 1984; Hughes and Thomas 1998). Higher educational attainment brings greater happiness, and the positive effect of education on happiness is found to be independent of the effects of earnings or income (Blanchflower and Oswald 2004; Easterlin 2001).

It is much less clear how social differentials in happiness change over the life course, time period, and birth cohort, or conversely, how age, period, and cohort effects on happiness vary depending on social status defined by sex,

race, and SES.<sup>1</sup> No previous study has tested these interaction effects when age, period, and cohort effects are simultaneously taken into account.

We can hypothesize that sex, race, and educational differentials exist in age variations in happiness. Because a host of other individual-level correlates that change with age—such as social solidarity in the form of marriage, health status, relative income level, labor force attachment, number of children, and religious attendance—all strongly affect subjective well-being (Easterlin 2003; Ellison 1991; Kohler, Behrman, and Skyttthe 2005; Waite 1995), we can further hypothesize that changes in social differentials over the life course can be largely attributed to differential exposures to these factors. The sex gap in life-course changes of happiness can be related to gender-related shifts of roles and behaviors during adult life. Some suggest that women are happier than men before old age, but less happy during old age. This shift happens because women suffer more from the adverse effects of widowhood and worse health in old age, whereas men benefit more from the positive effects of retirement and better health (Easterlin 2003; Shmotkin 1990). Cumulative advantage/disadvantage theory also suggests that socioeconomic gradients in subjective well-being vary over the life course (O’Rand 2003). The essence of the theory is that early inequalities in certain characteristics result in diverging life-course trajectories with the passage of time. Recent empirical tests of this theory are inconclusive with regard to health inequality—the effects of race and education on health outcomes have been found to increase, decrease, or remain constant over the life course (Ferraro and Kelley-Moore 2003; Kelley-Moore and Ferraro 2004). One can hypothesize that life-course patterns of happiness exhibit diverging race and education disparities because early life advantages or disadvantages accumulate with age and

<sup>1</sup> Education is the most commonly used indicator of socioeconomic status in life-course analysis of inequalities because it is usually stable across adulthood and does not exhibit the temporal variability of income and occupation. Income may not be an adequate indicator of lifetime social status because of the leveling effects of Social Security and the distribution of private pensions.



increase inequality in subjective well-being over the life course. That is, whites and the more educated have better access to all sorts of valuable resources such as money, social attachment, and healthcare that increase their sense of happiness, and these benefits accrue over the life course to produce even larger gaps in old age. Adjustments for these correlates should thus reduce the gaps.

Alternatively, the process of aging may level the effects of race and education through increasingly common life events that depress happiness—such as the loss of a spouse, withdrawal from the labor force, and declining health—and lead to less heterogeneity and a convergence in happiness levels in old age. It is also possible that the age effects on happiness reflect universal developmental changes and remain stable within each social stratum, which then translates to constant social gaps over the life course.

We can hypothesize that there have been period changes in recent decades in the gaps between men and women, blacks and whites, and the lower and higher educated, net of the effects of age and other individual-level covariates. Women, blacks, and people with less education could have experienced more favorable changes given the social and cultural trends in the United States during this period. Affirmative action during the post-1970 era drew ever increasing proportions of women into the labor force, and the civil rights movement brought more attention to minorities' welfare. The less well-off could have benefited more from improvements in the social welfare system, leading to diminishing gaps over time.

Extant analyses are constrained to the period from the 1970s to the 1990s, and they report mixed findings during this time. One study reports persistent black and white gaps in happiness (Hughes and Thomas 1998); another finds a narrowing racial gap, with blacks experiencing increases in happiness and whites experiencing downward trends (Blanchflower and Oswald 2004). In addition, there were possible decreases in educational gaps in happiness before the 1980s because the lowest education group experienced little decline in happiness and slight increases in the 1970s (Rodgers 1982), although there was also a report of constant gaps (Blanchflower and Oswald 2004). These findings concern marginal/gross period effects

that may be confounded with age or cohort effects. In addition, a lack of statistical tests of interaction effects in a multivariate framework and different time periods may contribute to the substantive differences in findings.

One can also expect that different birth cohorts with different formative experiences will show distinct patterns of changes in social disparities in happiness. Cohorts that came of age during a period of extreme socioeconomic adversities (during the Great Depression and World War II) exhibit the greatest social inequalities in physical health (Elder 1974; House, Lantz, and Herd 2005). Similar cohort patterns may occur in emotional well-being, although there is no study that tests this cohort by social status interaction effect.

In sum, this review of prior studies suggests there is an absence of identification of different components of temporal changes in happiness and unclear patterns of sex, race, and education inequalities in happiness by age and period. The enduring conceptual importance of cohort effects also suggests that a formal test of cohort changes in happiness is needed. Data limitations and descriptive analyses in previous studies have not allowed a systematic assessment of net age, period, and cohort effects. Whether the age, time, and cohort trends in overall happiness and social disparities in happiness hold when their confounding effects are delineated remains a question and merits further analysis. This study uses age-specific data for a sufficiently long time period, an appropriate research design, and improved statistical models to test hypotheses that can corroborate or dispel the existence of true age, period, and cohort changes in happiness and changing social inequalities in happiness across these temporal dimensions.

## DATA AND METHODS

### SAMPLES AND MEASURES

This study uses data from General Social Surveys (GSS) conducted over the 33-year period from 1972 to 2004. The GSS, an ongoing survey that has monitored the attitudes and behaviors of adults in the United States since 1972 (Davis and Smith 2005), is among the best sources of national data on happiness in the country. It spans a longer time period than any other survey of the United States, and it is part

of the World Database on Happiness (WDH) (Veenhoven 1992). Each survey uses multistage stratified probability sampling and includes a nationally representative sample of non-institutionalized adults ages 18 and older. Happiness is assessed as a single-item scale reported from respondents. The sample sizes range from about 1,500 to 3,000 across survey years. The data on happiness are available annually from 1972 to 1994 (except for 1979, 1981, and 1992) and biannually from 1994 to 2004. In all years, the GSS item on overall happiness is: "Taken all together, how would you say things are these days—would you say that you are very happy, pretty happy, or not too happy?" The responses are coded as 1 = very happy; 2 = pretty happy; and 3 = not too happy.<sup>2</sup>

Despite the simplicity of the happiness measure, there is considerable evidence of its psychometric adequacy. The measure has adequate validity and reliability (Veenhoven 1996). Cross-national studies have demonstrated substantial cross-group stability of similar subjective well-being measures (Chamberlain 1988). There is strong evidence that similar referents (i.e., the areas of life upon which judgments of happiness rest) are used both within and across nations (Veenhoven 1992). It has also been shown that measures of subjective well-being remain stable over time. The sources of happiness are stable across individuals and over time because the dominant concerns for most people are making a living, family life, and health (Easterlin 2001). Methodological studies that use data spanning a 16-year interval further enhance confidence in the usefulness of the measures over time. Factor analyses show high stability in the way the domain-specific measures of subjective well-being (e.g., concerns about oneself, fun, standard of living, job, and leisure) relate to

each other. Multiple classification analyses also show high stability in the way these measures contribute to or predict global well-being (life-as-a-whole or happiness). Analysts have concluded that the meaning and structure of measures of subjective well-being are remarkably constant over time (Andrews 1991). In fact, overall happiness is a core variable in "Quality of Life Surveys" used in many developed nations since the 1970s, and it provides the basis for comparative studies of subjective well-being in international research and time trend analysis (Rodgers 1982; Veenhoven 2005). Based on its widespread use in previous studies, I use the happiness variable to provide evidence that can be compared with other studies.

Key individual-level variables include respondent's age (reported in single years at last birthday, grand-mean centered, and divided by 10), sex (female = 1; male = 0), race (black = 1; white = 0), and education (years of education completed), which is coded as two dummy variables (less than a high school degree and college education or more; reference = 12 to 15 years). The analysis also adjusts for other characteristics that are shown to be correlated with subjective well-being. Relative income is coded as two dummy variables indicating the lowest quartile and the highest quartile (reference = middle two quartiles) in family income (converted to 1986 dollars and adjusted for family size). Marital status categories include married (reference), divorced, widowed, and never married. Health status is indicated by self-rated health and includes four categories: excellent, good (reference), fair, and poor. Work status includes full-time (reference), part-time, unemployed, and retired. Number of children is dichotomized as having no children (= 1) versus having one or more children. Religious involvement is measured by frequency of attending religious services, which ranges from never (0) to several times a week (8). As will be clear in the methods section below, survey years and birth cohorts are level-2 contextual variables in hierarchical models. The operational definitions and descriptive statistics of all variables used in the analysis are included in the Appendix, Table A.

There were 41,886 black and white respondents with data on reported happiness. The analysis focuses on black and white disparities and excludes a small number of respondents of other races (less than 3 percent). Regression

<sup>2</sup> There are scales tapping life satisfaction in the GSS for fewer survey years. The life satisfaction question has also been subjected to this model for age-period-cohort analysis. Because life satisfaction is measured using an index of domain-specific items that does not evaluate life-as-a-whole and that may relate to different social groups differently, this measure is not accepted as a valid indicator in the World Database of Happiness. For these reasons, my analysis does not include life satisfaction (details are available on request).

analyses show no significant differences between respondents of other races and black respondents in subjective well-being. The listwise deletion of missing values yields a final sample of 28,869 respondents.<sup>3</sup> The difference between the final sample and those excluded due to missing values is not statistically significant ( $t = 1.74, p = .082$ ). Although the exclusion of these cases substantially reduces the sample size, the final sample provides a sufficient number of observations for the subsequent regression analysis. Using the same sample across models facilitates the comparison of model fit.

### ANALYTIC METHODS

The GSS uses a repeated cross-sectional survey design. This design is increasingly available to social scientists, and it provides unique opportunities for age-period-cohort (APC) analysis. Pooling data from all survey years, one can formulate a rectangular age by period array of respondents, where columns correspond to age-specific observations collected in each survey year and rows are observations from each age across years. Linking the diagonal cells of the array yields the observations that belong to people born during the same calendar years who age together. Although only a longitudinal panel study design provides data from true birth cohorts that follow identical individuals over time, this design allows for a classic demographic analysis using the synthetic cohort approach (Mason and Fienberg 1985; Preston, Heuveline, and Guillot 2001), which traces essentially the same groups of people from the same birth cohorts over a large segment of the life span.<sup>4</sup> Compared to a longitudinal design,

which usually spans a short time period, the synthetic cohort approach has the advantage of facilitating the simultaneous tests of age and period effects because it is based on representative national surveys of all ages conducted regularly from one period to the next and covering more than three decades. It suffers less from the difficulty of locating sample respondents across time in panel studies, although it is not exempt from attrition due to mortality.

In spite of its theoretical merits and conceptual relevance, empirical APC analysis suffers from methodological problems. The fundamental question of determining whether the process under study is due to some combination of age, period, and cohort phenomena points to the necessity of statistically estimating and delineating these effects. The conventional statistical APC analysis focuses on modeling age-by-time period tables of *aggregate population data* (such as vital rates). The major challenge of estimating separate age, period, and cohort effects is the "identification problem" induced by the exact linear dependency among age, period, and cohort:  $\text{period} - \text{age} = \text{cohort}$ . The resulting regression coefficient estimates are not unique and cannot be used for statistical inference (Mason and Fienberg 1985).

The sample survey research design distinguishes itself from aggregate rates data because it yields individual-level observations, that is, *micro-level data*, and provides clues to solving the methodological problem. Note that the APC identification problem for aggregate population data does not necessarily transfer directly to the research design of repeated cross-sections. The aggregate data are usually confined to age and period measured in fixed interval lengths (commonly in one or five years), which creates the linear dependency problem. In a repeated cross-section survey, however, we can use different temporal groupings for the age, period, and cohort variables to break the linear dependency. Specifically, we can use single years of age, time periods corresponding to years in which the surveys are conducted, and birth cohorts defined by five-year intervals, which are conventional in demography. It is not possible to determine the exact age of each respondent from knowledge of the period of survey and the birth cohort in which each sample respondent is a member. Although we can determine the general five-year category in which each respondent is a member, there is no exact linear

<sup>3</sup> There are missing cases for different variables in each survey. Variables on health and income constitute the majority of missing cases. For instance, the question on health was not asked in 1978, 1983, or 1986, nor was it asked in the full sample in surveys from 1988 to 2004. This yields close to 9,600 missing cases. The income variable has more than 3,000 missing cases for all survey years combined. The listwise deletion results in a final sample of respondents who have data on all the variables in the analysis.

<sup>4</sup> The composition of cohorts in this setting will be affected somewhat by international migration.

**Table 1.** Percentage of “Very Happy” Respondents Cross-Classified by Period and Cohort: United States 1972 to 2004<sup>a</sup>

Cohort	Period									
	1972 to 1978		1980 to 1989		1990 to 1998		2000 to 2004		Total	
	Percent	N <sup>b</sup>	Percent	N	Percent	N	Percent	N	Percent	N
Pre-1899	36.9	377	41.9	86					37.8	463
1900	41.3	346	35.8	176	31.3	16			39.2	538
1905	35.5	501	38.5	257	26.2	84			35.5	842
1910	38.7	545	38.4	417	30.6	170	21.7	23	37.1	1,155
1915	36.5	647	35.9	518	32.9	289	17.9	67	34.8	1,521
1920	35.9	708	34.9	616	35.6	365	27.6	105	35.0	1,794
1925	34.6	673	34.1	602	39.6	399	42.5	127	36.1	1,801
1930	35.5	647	32.6	530	33.4	404	30.1	146	33.7	1,727
1935	35.4	775	32.0	596	34.6	468	40.2	174	34.6	2,013
1940	36.3	867	30.6	722	31.4	602	37.6	213	33.5	2,404
1945	30.2	982	28.7	969	29.6	798	31.9	279	29.7	3,028
1950	28.9	895	27.1	1,120	26.1	952	28.0	325	27.4	3,292
1955	28.1	295	29.0	1,116	30.2	1,082	33.6	372	30.0	2,865
1960			27.2	830	29.6	1,021	28.6	346	28.5	2,197
1965			31.0	306	29.6	878	36.1	366	31.4	1,550
1970			25.0	20	26.0	653	29.8	322	27.2	995
1975					23.8	235	33.1	293	29.0	528
1980					.0	4	30.3	152	29.5	156
Total	34.5	8,258	31.3	8,881	30.2	8,420	32.3	3,310	32.0	28,869

Source: The U.S. General Social Surveys conducted by the National Opinion Research Center (Davis and Smith 2005).

<sup>a</sup> Two other response categories are “pretty happy” and “not too happy.”

<sup>b</sup> Base N in each cell. Number of respondents is reported from the final sample.

dependence at the level of the individual respondent. Including other characteristics or covariates of individuals in the samples also lets us test explanatory hypotheses.

In addition, the nonlinear transformations approach, a conventional strategy used to solve this problem, applies a parametric nonlinear transformation, such as polynomials, to at least one of the age, period, or cohort variables so that its relationship to the others is nonlinear (Fienberg and Mason 1985). Following this strategy, and noting prior findings of curvilinear age effects from both exploratory data analysis and extant studies, this study specifies a model of happiness as a quadratic function of age.

The use of either or both of the above two specifications solves the underidentification problem, and fixed effects models could be estimated by adding variables to control for the age, period, and cohort effects in a conventional multiple linear regression analysis. As Yang and Land (2006:84–85) demonstrate, however, the assumption of fixed-period and cohort effects ignores the multilevel structure of the data design and may not be adequate. The mul-

tilevel data structure of the study samples is evident in Table 1, where respondents in the “very happy” state are nested in, and cross-classified by, the two higher-level social contexts defined by time period and birth cohort. That is, individual members of any birth cohort can be interviewed in multiple replications of the survey, and individual respondents in any particular wave of the survey can be drawn from multiple birth cohorts. It is possible that sample respondents who were surveyed at the same historic point and who belonged to the same cohort group may have similar responses because they share random error components unique to their period and cohort. Adequate models must take into account this level-2 heterogeneity for valid statistical inference.<sup>5</sup>

<sup>5</sup> Although ignoring the complicated error structure may not seriously affect the estimates of regression coefficients, it may lead to underestimated standard errors, inflated t-ratios, and Type I errors that are much larger than the nominal alpha level (Hox and Kreft 1994).



It is important to note that the APC identification problem is inevitable only under the specification of linear models of fixed age, period, and cohort effects that are assumed to be additive. Additivity, however, is only one approximation of the process of how social change occurs. The multilevel data structure suggests the use of models that can capture the contextual effects of period and cohort and reveal the social process by which individuals' well-being is shaped by their social environment. Such contextual effects of period and cohort can be well estimated as random effects in hierarchical or mixed models, which have been increasingly employed in social sciences. Furthermore, methodological comparisons of the fixed and random period and cohort effects formulations show that the random-effects specification is more statistically efficient in cases where the research design is "unbalanced" (e.g., the GSS data illustrated in Table 1), in the sense that there are unequal numbers of respondents in each year-by-cohort cell (Yang and Land forthcoming).

This analysis uses recently developed hierarchical APC models for repeated cross-section surveys (Yang and Land 2006) to test hypotheses about age changes in happiness that are independent of the effects of historical period and cohort membership. A cross-classified random-effects model (CCREM) is specified to take into account the embeddedness of respondents in the GSS surveys within a time period by birth-cohort cross-classified matrix (Raudenbush and Bryk 2002; Yang and Land 2006). Specifically, the model estimates fixed effects of age and other individual-level covariates and random effects of period and cohort and takes the following form:

Level-1 within-cell model:

$$Y_{ijk} = \alpha_{jk} + \beta_{1jk}A + \beta_{2jk}A^2 + \beta_{3jk}F + \beta_{4jk}B + \beta_{5jk}E + \sum_{p=6}^P \beta_p X_p + e_{ijk} \quad (1)$$

where  $Y_{ijk}$  stands for the ordinal response outcome of happiness of the  $i$ th respondent for  $i = 1, \dots, n_{jk}$  individuals within the  $j$ th period for  $j = 1, \dots, J$  time period and the  $k$ th cohort for  $k = 1, \dots, K$  birth cohort;  $A$  and  $A^2$  denote age and age-squared, respectively;  $F$  denotes being female;  $B$  denotes being black;  $E$  denotes level of education; and  $X_p$  denotes the vector of other individual-level variables such as age by

sex, age by race, age by education interaction terms, and control variables.  $\alpha_{jk}$  is the intercept indicating the cell mean for the reference group at mean age who were surveyed in year  $j$  and belong to cohort  $k$ ;  $\beta$  denotes the level-1 coefficients where  $P$  is the maximum number of covariates; and  $e_{ijk}$  is the random individual effect or cell residual.

Level-2 between-cell random intercept and coefficients model:

$$\alpha_{jk} = \pi_0 + t_{0j} + c_{0k} \quad (2.1)$$

$$\beta_{3jk} = \pi_3 + t_{3j} + c_{3k} \quad (2.2)$$

$$\beta_{4jk} = \pi_4 + t_{4j} + c_{4k} \quad (2.3)$$

$$\beta_{5jk} = \pi_5 + t_{5j} + c_{5k} \quad (2.4)$$

The level-2 between-cell models test the hypotheses about period and cohort effects through the specifications of random variance components for the random intercept and coefficients. Equation 2.1 is the model for the random intercept  $\alpha_{jk}$ , which specifies that the overall mean varies from period to period and from cohort to cohort.  $\pi_0$  is the expected mean at the zero values of all level-1 variables averaged over all periods and cohorts;  $t_{0j}$  is the overall *period effect* in terms of residual random coefficients of period  $j$  averaged over all birth cohorts with variance  $\sigma_{t_0}$ ; and  $c_{0k}$  is the overall *cohort effect* in terms of residual random coefficients of cohort  $k$  averaged over all time periods with variance  $\sigma_{c_0}$ . Similarly,  $\pi_3, \dots, \pi_5$  are the level-2 fixed-effects coefficients in Equations 2.2 to 2.4 that represent the fixed effects of sex, race, and education. To test whether sex, race, and education stratifications of happiness vary by time or birth cohort, the equations specify that their coefficients have period effects,  $t_{3j}, \dots, t_{5j}$ , and cohort effects,  $c_{3k}, \dots, c_{5k}$ , whose corresponding random variance components are  $\sigma_{t_3}, \dots, \sigma_{t_5}$ , and  $\sigma_{c_3}, \dots, \sigma_{c_5}$ . Other level-1 covariates are modeled as fixed across level-2 units. Both the period and cohort random variance components for the intercept and coefficients are assumed to have multivariate normal distributions (Raudenbush and Bryk 2002).

Based on the combined models in Equations 1 and 2, I estimate ordinal logit CCREMs of happiness using SAS PROC GLIMMIX (Littell et al. 2006).<sup>6</sup> The modeled probabilities are

<sup>6</sup> The sample SAS codes are available at [http://home.uchicago.edu/~yangy/apc\\_sectionE](http://home.uchicago.edu/~yangy/apc_sectionE).

cumulated over the lower ordered values based on the assumption of proportional odds, and the coefficients indicate the effects for the probability of being very happy versus others and the probability of being very happy or pretty happy versus not too happy. I use Bayesian Information Criterion (BIC) statistics to compare nested and non-nested models with respect to goodness of fit.<sup>7</sup>

Because the samples include young-adult respondents whose education may not be complete, I replicated the analysis starting at ages 25 and 30 so the majority of people in the samples would have finished their education. The results show that the inclusion or exclusion of younger adults does not substantively change the findings. Therefore, the analyses reported here use adults of all ages.

## RESULTS AND FINDINGS

Table 2 presents estimates of fixed effects coefficients in the form of odds ratios of being happy and random-effects variance components from the ordinal logit CCREMs. I first establish the overall trends and social differentials in happiness in an age-period-cohort modeling framework. Next, I examine the social differentials in happiness across the life course, over time, and across birth cohorts. Predicted probabilities of happiness are displayed in graphs from selected models to illustrate key findings.

### OVERALL TRENDS AND DIFFERENTIALS IN HAPPINESS

Models 1 to 4 in Table 2 estimate the bivariate associations of level-1 independent variables—age, sex, race, and education—with happiness.

<sup>7</sup> The BIC test adjusts the impact of model dimensions on model deviances and is a more generalized test of model fit than likelihood ratio test. The smaller the BIC, the better the model fit (see, e.g., Raftery 1986). I estimate the hierarchical ordinal logit models of happiness based on the residual pseudo-likelihood estimation. There is no overall goodness-of-fit statistic available in the literature that can be used to directly compare models estimated using pseudo-likelihood (Littell et al. 2006). Therefore, the model fit comparison of the ordinal logit models is based on the BIC statistics obtained from the normal HLM.

Model 5 estimates the main effects of all level-1 independent variables jointly.

Model 1 shows a significant quadratic age effect net of the random period and cohort effects. It suggests that, adjusting for time period and birth cohort variations, the odds of being happy increase 5 percent with every 10 years of age (odds = 1.049, CI = [1.03, 1.069]), but the increase declines at the rate of 1 percent every decade over the life course (odds = .989, CI = [.981, .997]). Figure 1 presents the overall trends of happiness in terms of predicted probabilities of being very happy, estimated from Model 1.<sup>8</sup> Figure 1a shows the curvilinear and concave age effects. The estimates of random effects in terms of residual variance components at level-2 indicate significant period and cohort effects controlling for the age effect, as reported in the lower panel of Table 2. Thus, net of the age effects, the mean odds of being happy vary significantly by time period and birth cohort.

Figure 1b displays the estimated period effects, namely, the predicted probabilities of being very happy for each year at the mean age and averaged over all birth cohorts, which is calculated as  $\hat{\alpha}_j = \hat{\pi}_0 + t_{0j}$ , where  $\hat{\pi}_0$  is the intercept or estimated overall mean and  $t_{0j}$  is the period-specific random-effects coefficients estimated from Model 1. The period effects show a general weak downward trend across the first two decades, which is consistent with findings from previous studies based on GSS data, but they also exhibit clear nonlinear declines over time. There were continuous declines in happiness after 1985, followed by a gradual rebound since 1995 to a level similar to that in the 1970s and early 1980s. The magnitude of such period changes is relatively small: the predicted probabilities of being very happy largely fell between .30 and .35 over the past three decades.

Figure 1c shows the estimated cohort effects in terms of the predicted probabilities of being very happy at the mean age and averaged over all periods. Similar to the period effects, it is calculated as  $\hat{\alpha}_k = \hat{\pi}_0 + c_{0k}$ , where  $c_{0k}$  is the cohort-specific random-effects coefficient estimated from Model 1. As is the case for period effects, the cohort effects are small relative to the age

<sup>8</sup> The predicted log-odds or logits, denoted as  $\beta$ , are converted to predicted probabilities by computing  $probability = 1/(1 + \exp(\beta))$  (see, e.g., Raudenbush and Bryk 2002).

**Table 2.** Odds Ratio Estimates from Ordinal Logit Cross-Classified Random Effects Age-Period-Cohort Models of Happiness

Fixed Effects	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Intercept <sup>a</sup> , $\pi_0$	.495***	.472***	.525***	.509***	.524***	.523***	.520***	.376***
Age, $\pi_1$	1.049***				1.056***	1.097***	1.098***	1.084***
Age <sup>2</sup> , $\pi_2$	.989**				.995	.997	.997	1.027***
Female, $\pi_3$		1.036			1.073**	1.069**	1.068	1.191***
Black, $\pi_4$			.481***		.516***	.523***	.529***	.665***
Education (ref. = 12-15)								
Education 1 (< 12), $\pi_5$				.681***	.721***	.700***	.701***	.957
Education 2 ( $\geq$ 16), $\pi_6$				1.368***	1.314***	1.294***	1.299***	1.006
Age $\times$ Female						.897***	.898***	.968*
Age $\times$ Black						1.117***	1.119***	1.151***
Age $\times$ Education 1						1.033*	1.032	1.039*
Age $\times$ Education 2						.974	.973	.967*
Income (ref. = 2-3 Qs)								
1st Quartile								.867***
4th Quartile								1.264***
Marital Status (ref. = married)								
Divorced								.389***
Widowed								.333***
Single								.474***
Health (ref. = good)								
Excellent								1.945***
Fair								.551***
Poor								.299***
Work Status (ref. = full time)								
Part-time								.892**
Unemployed								.516***
Retired								1.144**
No Children								1.113***
Religious Attendance								1.088***

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Table 2. (continued)

Random Effects—Variance Components	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Period Effect								
Intercept, $\sigma_0$	.007*	.006*	.004*	.006*	.008*	.007*	.003	.005*
SEX effect, $\sigma_3$							.013*	.006
RACE effect, $\sigma_4$							.030	.037*
Cohort effect, $\sigma_{t0}$	.003*	.009*	.007*	.030*	.007*	.007*	.006*	.009*
Goodness-of-fit (BIC)	214819	214829	214348	214440	214056	213996	213943	210197

Note: Fixed effects logit coefficients, random effect coefficients, and their standard errors are omitted in the interest of space.

<sup>a</sup> Intercept of being very happy (1) versus others (2 and 3). The intercept of being very or pretty happy (1 and 2) versus not too happy (3) is omitted. \*  $p \leq .05$ ; \*\*  $p \leq .01$ ; \*\*\*  $p \leq .001$  (two-tailed tests).

effects and show no linear increases or declines. The cohort trends are flat for the earlier and later cohorts. The most intriguing finding is the dip in the predicted probabilities of being very happy for cohorts born between 1945 and 1960 (i.e., the baby boomers).

Results from Models 2 to 4 show that, as expected, women, whites, and college graduates on average have higher odds of being happy relative to men, blacks, and people with less education. Whereas the sex effect is small and not statistically significant, the race and education effects are substantial and highly significant. Being black is associated with an over 50 percent reduction in the odds of being happy (odds = .481, CI = [.450, .514],  $p < .001$ ). Having a college degree increases the odds of being happy by about 37 percent (odds = 1.368, CI = [1.290, 1.451],  $p < .001$ ), whereas having less than a high school degree decreases the odds by about 32 percent (odds = .681, CI = [.643, .721],  $p < .001$ ). These results largely hold in Model 5 when all attributes are considered, with two exceptions. Net of the effects of sex, race, and education, the quadratic age coefficient is not statistically significant, indicating a positive linear age effect. And the positive effect of being female increases and becomes statistically significant when holding race and education constant. The smaller BIC statistic indicates a better model fit for Model 5 than any of the previous models. These results are not surprising in light of previous studies of social correlates of happiness. But they also suggest two new findings. One is that the individual-level effects hold when level-2 heterogeneity, represented by the period and cohort effects, are taken into account. Second, net of age and other social status indicators, there are significant variations in odds of happiness that can be attributed to period- and cohort-specific factors.

Models 6 to 8 are additive models to Model 5 that include significant interaction effects at levels 1 and 2. The final model (Model 8) yields the smallest model deviance adjusted by degrees of freedom as measured by the BIC statistic and, therefore, significantly improves the model fit over all previous models in Table 2 and alternative models.<sup>9</sup>

#### AGE VARIATIONS IN SEX, RACE, AND EDUCATION EFFECTS ON HAPPINESS

Model 6 shows that the sex, race, and educational gaps in happiness vary significantly with age. Model 7 further shows that these interaction effects remain significant when the ran-

<sup>9</sup> I also examined alternative model specifications such as fixed-period and cohort effects (see the Methods section). These produced fixed-effect coefficients and linear trends that are largely consistent with the results presented above, but such specifications did not reveal nonlinear changes over time and across cohorts.



Figure 1. Overall Age, Period, and Cohort Effects on Happiness: GSS 1972 to 2004



dom-period variations in sex and race coefficients are considered. Figure 2 shows the predicted probabilities of being very happy from Model 7 and indicates the very different life-course patterns of happiness for men and women, blacks and whites, and the more educated and the less educated. Figure 2a compares the trajectories of age changes in the probability of being very happy by sex and race. For both blacks and whites, men and women show crossovers of happiness trends by age, with women being happier before middle age and less happy afterward. The racial gap in happiness is larger than the sex gap for a large segment of the life course. Whites' advantage over blacks starts off large and gradually decreases with age. White men over the age of 50 appear to be the happiest of all, whereas black women of the same ages appear to be the least happy. Although the age trajectories trend upward for most groups, white women experience a slight decline in happiness with age. Figure 2b presents the age and education interaction effects. Adjusting for other factors, a large educational gradient in predicted probabilities of happiness remains. In addition, the mean education effects also show some sign of convergence across the three categories of attainment during old age, but to a much less degree than for the race effects.

Model 8, the full model, adjusts for other correlates of happiness. As expected, the effects of relative income, marital status, health, work status, children, and religious involvement all play a significant role in affecting happiness. The negative effect of being poor is twice as large as the positive effect of financial abundance on happiness. Relative to the middle two income quartiles, being in the lowest income quartile decreases the odds of happiness by about 26 percent and being in the highest income quartile increases the odds of happiness by about 13 percent. Marital status and health status have by far the strongest influences on one's sense of happiness. Compared to the married, those who are widowed, divorced, and single are about 70, 60, and 50 percent less likely to be happy, respectively. Compared to those in good health, people in excellent health are almost twice as likely to be happier, whereas those in poor health are 70 percent less likely to be happy. The results also show that lower levels of labor-force attachment decrease hap-

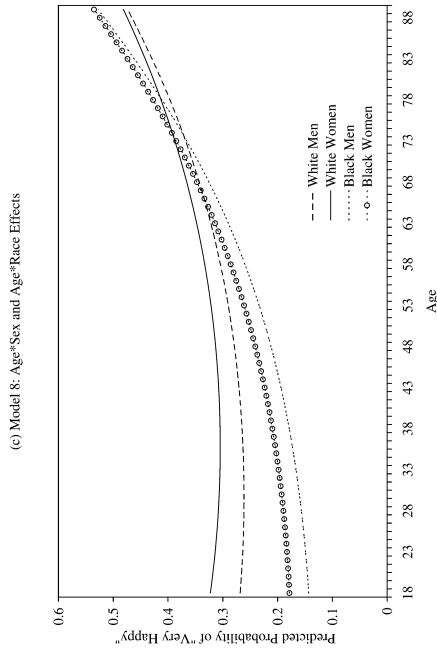
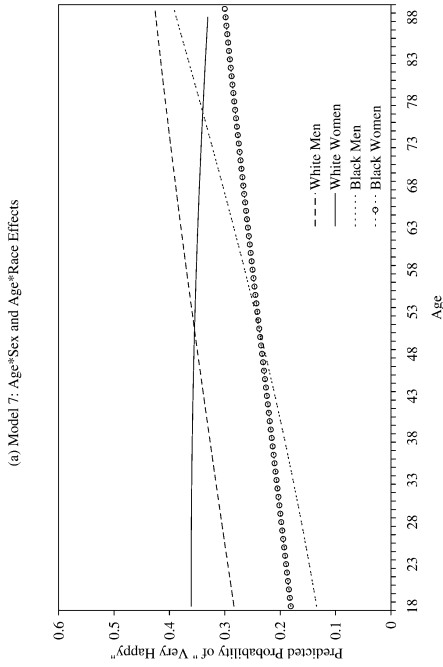
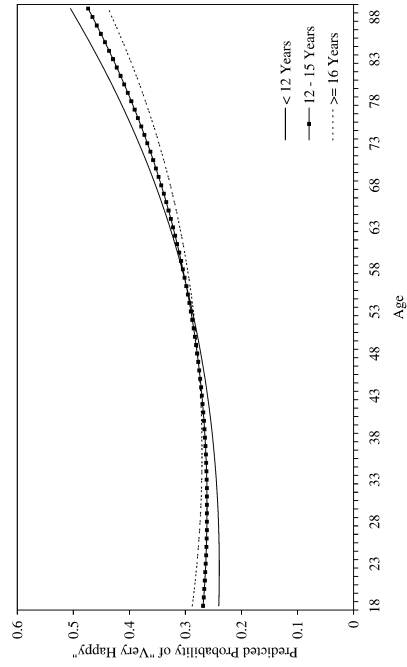
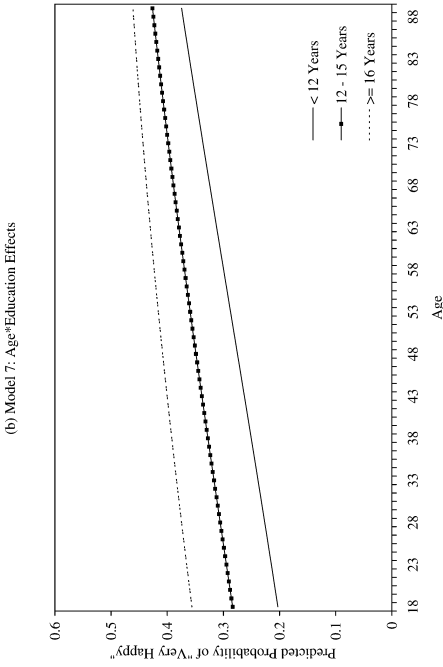
piness. Having no children increases the odds of being happy.<sup>10</sup> More frequent religious attendance also slightly increases happiness, net of all other factors.

To what extent can life-course patterns of happiness be accounted for by differential exposures to the above conditions? Adjusting for all of the above factors, Model 8 suggests that the main age effect remains highly significant and has changed direction. Figure 2c compares the predicted probabilities of being very happy by age, estimated from the full model for various social groups. The shapes of the age curves become quadratic and convex. The J-shaped net age effects suggest that the average happiness level bottoms out in early adulthood and increases at an increasing rate as one moves through the life course.<sup>11</sup> Model 8 also shows that salient sex and race differentials in happiness persist even when major correlates of happiness are held constant. But the race effect decreases in size, and the effects of educational attainment are no longer significant. This means that some of the effect of being black and all of the effect of education are mediated by other covariates.

All the interaction terms remain significant when all things are considered, so the life-course trajectories of happiness still show appreciably different trends. When potential explanatory factors are controlled, however, the interaction effects of sex, race, and education with age shrink in size. The adjustment for marital status, health, and work status decreases the age by sex interaction effect. This supports the hypothesis that the changes in the gender effect on

<sup>10</sup> Additional analysis of bivariate associations suggests that having no children is associated with less happiness. But childbearing does not always have positive effects on happiness. When marital partnerships are considered together with fertility, as the full model suggests, the effect becomes negative. This corroborates Davis's (1984) finding that "two's company, three's a crowd"—namely, married partners with children are less happy than those living as pairs.

<sup>11</sup> The age corresponding to the minimum predicted happiness, for example, is 29.6 for the reference group. The age of minimum happiness can be calculated by setting the derivative of Y with regard to age variables to zero and solving the equation for age.



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**Figure 2.** Predicted Age Variations in Sex, Race, and Education Effects on Happiness

Notes: Model 7 includes all independent variables (age, age<sup>2</sup>, sex, race, education, and their interaction effects). Model 8 also adjusts for control variables. Both models are graphed for the reference groups.

happiness with age are partly due to increases in widowhood and worsening health among women (which decrease happiness) and to retirement and better health among men (which increases happiness). Specifically, the sex gap in happiness changes from a crossover (Figure 2a) to a convergence at old age (Figure 2c). Everything else being equal, women are happier than men throughout the life course, but they are not much happier in late life. When other factors are held constant, the race differentials show a stronger form of convergence, with reversals occurring around age 70, after which blacks become slightly happier than whites. The educational gap also decreases in magnitude in the final net effects model, as shown in Figure 2d. There appear to be crossovers in age effects of happiness among the three educational groups at age 50, after which the highly educated are not particularly better off than the less educated. These convergences, which indicate a loss of advantages for women, whites, and the highly educated with age, do not seem to be completely explained by marital status, health, or other factors.

### **TIME TRENDS OF SEX AND RACE DIFFERENCES IN HAPPINESS**

Model 7 shows significant period changes in sex and race inequalities net of age effects, cohort changes, and other factors. Figure 3 shows that both sex and race disparities decreased during the past 30 years in the United States. Figure 3a shows that the reduction in the sex gap is due to women's decreasing levels of happiness and men's stable trend as of 1995 and similar increases for both afterward. The sex differentials are relatively small even at the initial time period, and they are not statistically significant in the final model when other factors are taken into account.

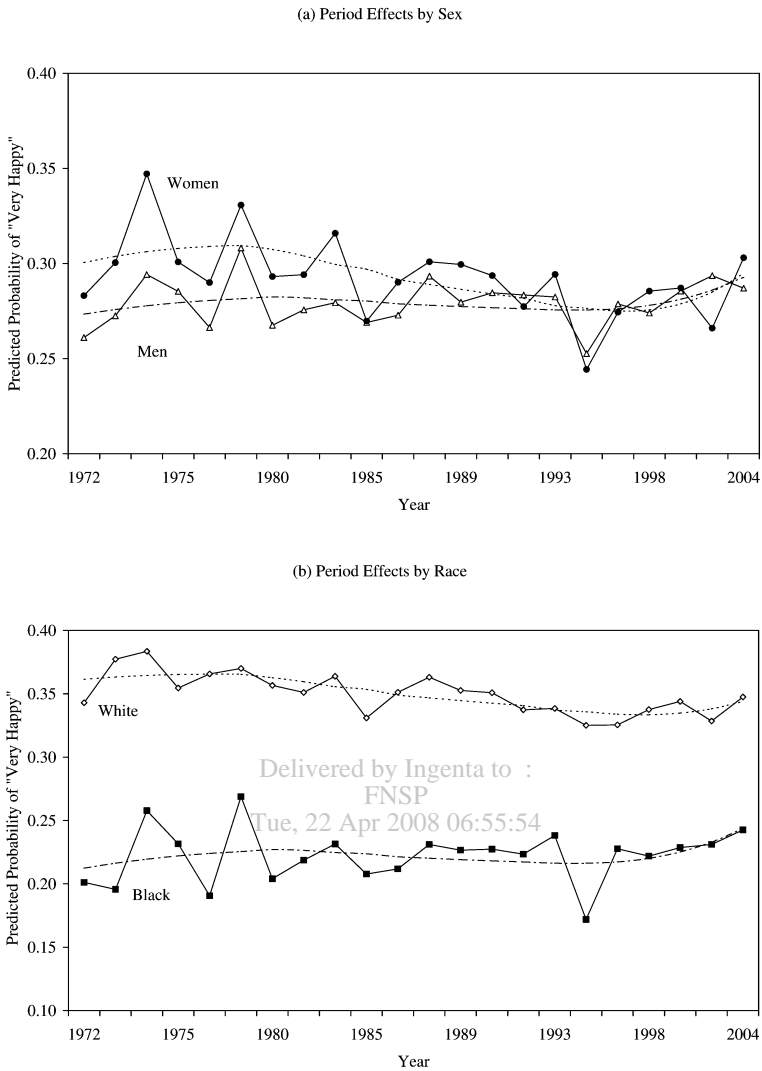
Figure 3b shows that the black and white gap was much more pronounced throughout the 30-year period, but there were small declines in more recent years. From the 1970s to the mid-1990s, whites experienced a slight downward trend in happiness and blacks experienced some fluctuations early on and a stable trend after 1980. Both groups fared better after 1995. Although gross/marginal period effects with no controls for anything suggest a convergence in black and white mean levels of happiness in

the last two survey years (2002 and 2004) due to large increases for blacks, the net period effects suggest a lack of convergence. Blacks experienced more pronounced improvement, but this did little to close the gap. Because gross period effects are estimated for all age and cohort groups present in different survey years, changes in age and cohort composition would strongly affect the estimates of time trends. For instance, blacks' higher total fertility rates across the twentieth century produced a younger age structure for blacks than whites in most periods. But decreases in the fertility rates of blacks, bringing them closer to the rates for whites in recent years, are associated with shifts in blacks' age structure so that it more closely resembles whites' older structure (Yang and Morgan 2004). Because younger adults are estimated to be less happy than older adults, decreases in the proportion of young adults could lead to increases in happiness. In addition, because the baby-boomer cohort is estimated to have the lowest levels of happiness, larger decreases in the proportional representation of baby boomers in blacks in recent survey years could also account for the convergence. Adjusting for the age and cohort effects in the model, therefore, yields period trends purged of the effects of compositional changes. Such period changes in racial differences remain statistically significant in the final model.

None of the three stratifying effects show any significant variations by birth cohort, meaning the advantages of women and the highly educated and the disadvantages of blacks and the less educated are constant across successive cohorts.

### **DISCUSSION**

The vast majority of research on subjective well-being to date focuses on its cross-sectional individual-level determinants. Prior to this study, there was no comprehensive temporal model for understanding the distinct effects of internal characteristics and external circumstances on perceptions of life quality. Using time-series data on happiness that span 33 years, and improved statistical models that disentangle the confounding effects of age, period, and birth cohorts, this study provides new evidence of life-course and temporal changes in subjective quality of life in the United States that can



**Figure 3.** Predicted Period Variations in the Sex and Race Disparities in Happiness

*Notes:* The solid lines are random period effects estimated from Model 7 at mean age of the reference group and averaged over all birth cohorts. The dotted lines imposed on the period effects indicate the trends fitted by regression models using linear, quadratic, and cubic functions of year.

be attributed to the processes of aging and social change. The analysis also reveals previously unknown patterns of changes in social inequalities in happiness across the adult life course and over historical time. Five major findings emerged.

First, the results show life-course patterns, time trends, and birth cohort differences in happiness that are distinct and independent of each other. The significant net age, period, and cohort effects suggest it is important to test variations

formally in all three time-related dimensions in studies of changes in subjective well-being to arrive at adequate interpretations and valid inferences with regard to these effects.

Second, with age comes happiness. That is, overall levels of happiness increase with age, net of other factors. This supports the “age as maturity” hypothesis suggested by the role theory of aging. The age effects are strong and independent of the time period and cohort effects. Adjusting for relevant social correlates changes

the shapes of the age trajectories of happiness but does not alter their positive slopes. It is important to note that in models of happiness where all three temporal factors are considered, the age effects dominate and the period and cohort variations are small. This suggests that studies of temporal changes in subjective well-being that ignore life-course changes may be misleading in giving the impression that time trends and cohort differences of happiness are more substantial than they actually are.

Third, there are life-course variations in the social disparities of happiness. All three stratifying effects examined—sex, race, and education—show decreases with age, suggesting that the cumulative advantage/disadvantage model, which predicts diverging social inequalities with age, does not apply to life-course changes in happiness. Instead, the present results resemble findings of convergence in SES differences in health-related studies (Ferraro and Kelley-Moore 2003; House et al. 2005). Several mechanisms could be at work. My analysis shows that adjusting for factors strongly associated with happiness largely reduces the sex and race gaps and diminishes the education gaps in happiness in later life. Therefore, the decreasing sex, race, and education gaps in happiness with age can be substantially explained by the differential exposures of these groups to various social correlates of happiness, especially in middle to old age. Eligibility for social welfare benefits such as Medicare and Medicaid, retirement from the paid labor force, and increasing stressful life events such as widowhood and deaths of relatives and friends can minimize gender, race, and socioeconomic differences in happiness in old age. These events equalize access to health care, and the loss of social support and social integration among the elderly erodes the advantages that some experience in earlier life stages. This leveling of protective and harmful factors can attenuate social differences and lead to a reduction in disadvantages for men, blacks, and people with low education in the latter part of the life course.

Decreasing social disparity over the life course could also be related to changes in the impact of social correlates on happiness with age. It is possible that resources and social status have less impact on happiness in old age because as one matures one becomes more immune to life stresses. This analysis assessed

this possibility by including a series of interaction terms of age and each of the other correlates in the models. No significant age differences in the impact of these factors were found, so they do not explain the group convergences in life-course patterns.

Selective survival also may play a role in explaining the positive age effect and the leveling of stratification effects. Higher levels of happiness in older adults may result from the selective survival of respondents who are happier (Danner, Snowdon, and Friesen 2001). This study cannot test the effects of selective survival, due to the lack of follow-up data on mortality. This analysis provides preliminary evidence that selective survival is not solely responsible, because the full model suggests that the positive age effect remains even after adjusting for health status. Selective survival of happier respondents may also reduce heterogeneity in happiness in later life and give the appearance of convergence in social disparities across the life course. Additional analyses show that cohort means of factors that strongly predict survival (such as SES and health) did not universally increase and converge over time. Therefore, selective survival may have contributed to the convergence in age patterns of happiness, but it is unlikely that it acts as the only or dominant force that drives such changes. Nonetheless, future studies should use longitudinal panel data to sort out which of the above processes are more plausible in accounting for stratifications of aging and happiness.

The fourth finding is that levels of happiness also changed over time periods. The use of different analytic foci and time periods, together with insufficient consideration of confounding effects, made it difficult for previous empirical studies to discern true period effects. My results, from a multivariate age-period-cohort analysis, reveal that net of other factors, Americans were happier in some years than in others. In support of the relative utility theory, the positive cross-sectional effects of economic conditions on subjective well-being did not translate into continuous period increases of happiness. Instead of a linear decline in happiness as shown in some previous studies using the same data, I find a nonlinear trend that was downward first and upward later. These period effects are statistically significant albeit small. The seemingly stable time trends in happiness



mask substantial subgroup differences. Both sex and race inequalities declined over the past 30 years, showing signs of social progress. From 1972 to 1995, happiness levels remained stable for men and blacks but worsened slightly for women and whites. After 1995, all groups experienced improvement. It is an important finding that minorities such as women and blacks did not experience less improvement than the general population. The other major trend is that, while the male and female gap largely disappeared in the most recent decade, the black and white gap persisted. These findings corroborate and extend earlier findings of trends in racial disparity in quality of life in the United States (Blanchflower and Oswald 2004; Hughes and Thomas 1998). They suggest that such disparity, although having decreased moderately with time, continued into the first decade of the twenty-first century.

What could have contributed to period changes is not well understood. Cross-national analyses of happiness data collected from Europeans and Americans from the 1970s to the 1990s suggest that the changes during this period could be correlated with changes in macroeconomic variables such as gross domestic product (GDP) and levels of joblessness (Di Tella, MacCulloch, and Oswald 2003). There is evidence that a higher GDP buys extra happiness and higher unemployment rates decrease happiness, but the effects of macroeconomic conditions are generally small. If period changes are induced by economic changes, then the results for period changes in sex and race differentials could mean that the effects of economic improvement on one's sense of well-being were more pronounced for men and blacks compared to women and whites. The apparent decline in the happiness level of women relative to men across time could also reflect the effects of increasing female labor force participation and rising divorce rates. It is possible that women suffered more from these cultural changes due to increasing levels of social stress associated with balancing work and family.

Last, but not least, cohort changes in happiness exist that are not confounded with the aging and period effects. There is barely any evidence in prior studies for cohort differences in happiness, though there are speculations that more recent cohorts have experienced lower

levels of happiness. My results do not strongly support such a hypothesis when confounding effects are controlled. This study does not find the monotonic declines in levels of happiness in successive birth cohorts that previous research had speculated might be engendered by cohorts' value shifts under the influence of "post-materialism." I find that baby boomers have experienced less happiness on average than both earlier and more recent cohorts. Therefore, fortunes do seem to be more closely related to early life conditions and formative experiences. Larger cohort sizes increase the competition to enter schools and the labor market and create more strains to achieve expected economic success and family life. The unique experiences of these cohorts during early adulthood can have a lasting impact on their sense of happiness.

The advantage of a hierarchical APC regression modeling approach is that it allows the inclusion of period-level and cohort-level covariates to test the above hypotheses of contextual effects. Future analyses should expand the level-2 models to test the relationships of economic prosperity and cohort characteristics and period and cohort changes in happiness.

There are several other questions this study has not resolved and that merit additional research. First, higher levels of subjective well-being in old age are regarded as a paradox. That is, despite physiological declines, the onset of frailty, and social losses such as widowhood, older adults are able to appraise their quality of life positively and sustain high levels of well-being. The results from the full models suggest that the residual age effects remain significant even when relevant social conditions are held constant and confounding temporal factors are controlled. Social psychological perspectives offer additional explanations to the role theory of aging. They relate the positive age effects to age differences in social comparisons and goal discrepancies. Studies of British and Hong Kong elderly samples suggest that older adults tend to be happier because they are more likely to use downward social comparisons (i.e., compare themselves to the less advantaged) (Gana, Alaphilippe, and Bailey 2004) or because they have smaller discrepancies between aspiration and achievement, especially in the domains of material resources and social relationships, than their younger counterparts (Cheng 2004). Further study using data on social-psychological meas-

urements is needed to determine whether or how these aging-related mechanisms work to explain the positive age effects in the American population.

Second, although the repeated cross-sectional design facilitates age-period-cohort analysis, only longitudinal panel data can provide the strongest and most conclusive evidence of intra-individual age changes. Unfortunately, there are few nationally representative longitudinal surveys that include more than two or three waves of measures on subjective quality of life to allow for inferences about life-course patterns across adulthood. I therefore interpret the age effects in this study in the context of synthetic cohorts rather than real birth cohorts that were followed prospectively. This analysis minimizes the possible difference in two ways. Because the GSS are nationally representative in every survey year, the constructed birth cohorts are seen as consisting of the same individuals from year to year and, in this sense, synthetic—this assumption underlies life table models in conventional demographic studies. Controlling for cohort changes in the regression models also helps to disentangle aging effects from cohort effects. That is, the results show significant age changes within any given synthetic birth cohort. In all, my analyses provide findings that largely improve our understanding of the relationships between age and cohort differentials in happiness. These findings undoubtedly need to

be corroborated and supplemented with those from future panel studies.

Third, it is surprising that average levels of happiness show very small changes over the past 30 years and across cohorts. There have been no previous sociological discussions about stability in cohort trends of happiness, but the stability and the lack of increase in happiness over time across a number of developed countries have been objects of attention and even debates among economists (Veenhoven and Hagerty 2006). No agreement about the direction of the change in happiness over time has been possible in these debates, partly due to the lack of age-period-cohort analysis that could explicate the true period effects. This study contributes stronger evidence of period changes in one of the wealthiest countries in the world. Based on this, we can ask: How can such stability be explained when the society has gone through tremendous economic growth and changes in the social and political environment? Cross-national comparative studies are needed to understand to what extent social structures and processes produce changes in happiness at the population level.

*Yang Yang is Assistant Professor of Sociology and research associate of the Population Research Center and the Center on Aging at NORC at the University of Chicago. Her research interests include sociology of aging and the life course, cohort analysis, mathematical and medical demography, and Bayesian statistical methods.*

## APPENDIX

**Table A.** Summary Statistics for All Variables in the Analysis: GSS 1972 to 2004 (N = 28,869)

Dependent Variable	Description and Coding	Mean	SD	Min	Max
Happy	Level of happiness: 1 = very happy; 2 = pretty happy; 3 = not too happy	2.20	.64	1	3
<b>Level-1 Variables</b>					
Age	Respondent's age at survey year Centered around grand mean and divided by 10	44.84 0	17.12 1.71	18 -2.68	89 4.42
Female	Respondent's sex: 1 = female; 0 = male	.55	.50	0	1
Black	Respondent's race: 1 = black; 0 = white	.14	.34	0	1
Education	Respondent's years of schooling	12.55	3.16	0	20
Education 1	1 = 0 – 11 years; 0 = otherwise	.26	.44	0	1
Education 2	1 = 16 years or more; 0 = otherwise	.20	.40	0	1
Income	Family income in 1986 dollars/household size (in thousands)	13.23	13.22	.05	162.61
1st quartile	1 = lowest 25 percent; 0 = otherwise	.25	.43	0	1
4th quartile	1 = highest 25 percent; 0 = otherwise	.25	.43	0	1
Marital Status	Respondent's marital status				
Divorced	1 = divorced or separated; 0 = otherwise	.15	.36	0	1
Widowed	1 = widowed; 0 = otherwise	.09	.29	0	1
Single	1 = never married; 0 = otherwise	.18	.39	0	1
Health Status	Respondent's self-rated health				
Excellent	1 = excellent; 0 = otherwise	.32	.47	0	1
Fair	1 = fair; 0 = otherwise	.18	.39	0	1
Poor	1 = poor; 0 = otherwise	.05	.23	0	1
Work Status	Respondent's work status				
Part-time	1 = working part-time; 0 = otherwise	.10	.30	0	1
Unemployed	1 = unemployed; 0 = otherwise	.03	.17	0	1
Retired	1 = retired; 0 = otherwise	.12	.32	0	1
No Children	Number of children: 1 = no children; 0 = 1 or more children	.27	.44	0	1
Religious Attendance	Frequency of attending religious services: 0 = never; . . . ; 8 = several times a week	3.89	2.67	0	8
<b>Level-2 Variables</b>					
Period	Survey year		N	Min	Max
			22	1972	2004
Cohort	Five-year birth cohort		18	1899	1986

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