Original Research Article

The transition to Post-Industrial BMI Values Among US Children

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ABSTRACT The trend in the BMI values of US children has not been estimated very convincingly because of the absence of longitudinal data. Our objective is to estimate time series of BMI values by birth cohorts instead of measurement years. We use five regression models to estimate the BMI trends of non-Hispanic US-born black and white children and adolescents ages 2–19 between 1941 and 2004. The increase in BMIZ values during the period considered was 1.3σ (95% CI: 1.16σ; 1.44σ) among black girls, 0.8σ for black boys, 0.7σ for white boys, and 0.6σ for white girls. This translates into an increase in BMI values of some 5.6, 3.3, 2.4, and 1.5 units, respectively. While the increase in BMI values started among the birth cohorts of the 1940s among black girls, the rate of increase tended to accelerate among all four ethnic/gender groups born in the mid-1950s to early-1960s. Some regional evidence leads to the conjecture that the spread of automobiles and radios affected the BMI values of boys already in the interwar period. We suppose that the changes in lifestyle associated with the labor saving technological developments of the 20th century are associated with the weight gains observed. The increased popularity of television viewing was most prominently associated with the contemporaneous acceleration in BMI gain. Am. J. Hum. Biol. 21:151–160, 2009.

Although descriptive statistics pertaining to the increasing prevalence of overweight and obesity of US children and adults have been very extensively reported (Freedman et al., 1997; Hedley et al., 2004; Ogden et al., 2002, 2004, 2006, 2008; Strauss and Pollack, 2001), trends have been identified less convincingly. For example, Troiano and Flegal (1998) suggest that "Overweight prevalence increased over time, with the largest increase between NHANES II and NHANES III," surveys, that is to say between those measured in the 1980s and early 1990s. Freedman et al. (2006) suggest in a similar vein that "the secular increases among black girls began during the 1970s, whereas increases among other children were not evident until the 1980s." Moreover, Anderson et al., (2003) suggest that "the increase in obesity began between 1980 and 1988." Hence, the implication seems to be that the epidemic began in the 1980s. [The upswing in overweight is said to have begun in Australia in the 1970s (Norton et al., 2006).] The drawback of these studies is that they all refer to period effects (measurement years) rather than birth cohorts. Insofar as a 13-year-old measured in 1988 was born in 1975, it is not at all clear from the cross-sectional evidence when that youth actually became overweight. It could have happened at anytime between 1975 and 1988. Another shortcoming of the conventional approach is that it generally does not control for earlier maturation. [The secular trend toward earlier maturation is well established (Rogol et al., 2002).] Thus, the 13-year-old might have a higher BMI value partly because of a tempo effect, i.e., she has now a bone-age of the 14-year-old of a generation ago.

Our hypothesis is that an examination of the trends using birth cohorts could provide insights that have hitherto eluded researchers. The extant studies tend to imply that the epidemic appeared rather suddenly without, however, being very precise about the beginning of the upswing. In contrast to the consensus in the literature, we use birth cohorts for our estimates insofar as BMI at the time of a particular measurement reflects the accumulated weight gains during the life course. Birth cohorts experienced similar social, economic, and technological changes, whereas the same cannot be said of measurement cohorts. For example, those measured in 1960 have been exposed to television viewing for different lengths of time during their lives. 19-year-olds, for example, would have been born in 1941 and would have spent much of their youth without watching television while 5-year-olds also measured in 1960 would have had the opportunity to watch TV all their lives. In contrast, considering the sample by birth cohorts implies that all those born in, say, 1955 were exposed to TV viewing all their lives regardless of when they were measured.

Another considerable advantage of the birth-cohort approach is that instead of having only five data points from the cross-sectional surveys (1959–2006) from which merely four differences can be calculated, we obtain data for almost every year continuously during the period under consideration (Fig. 1). Furthermore, the continuous approach also enables us to calculate the rate of change of BMI values which is not possible with the cross-sectional approach. Hence, in the main, our goal was to estimate the trend in BMI values in more detail and a bit more convincingly than has been done up to now.

DATA

We estimated for the first time the long-term trends in the BMI values (kg/m^2) of children and youth ages 2–19 years continuously for the birth cohorts 1941–2004 stratified by gender and ethnicity on the basis of surveys

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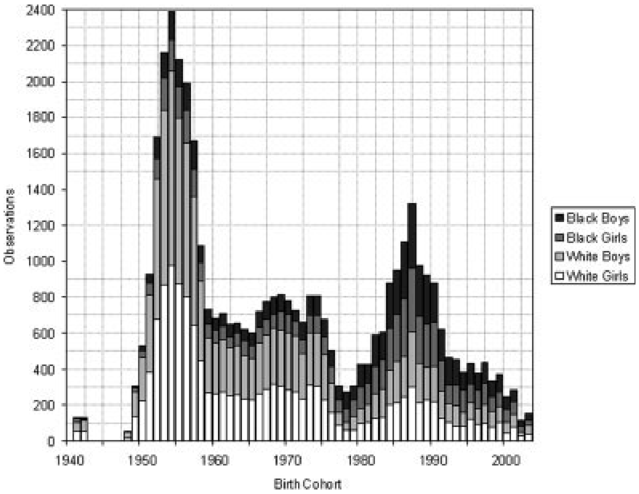


Fig. 1. Number of observations by birth cohort. Combined NHES and NHANES Samples, by birth cohorts 1941–2004.

collected between 1959 and 2006 by the National Center for Health Statistics (NCHS). We concatenated the following National Health Examination Surveys: [NHES I: 1959–1962 (ages 18–19), NHES II: 1963–1965 (ages 6–11), NHES III: 1966–1970 (ages 12–17).], and the National Health and Nutrition Examination Surveys: (NHANES I: 1970–1974, NHANES II: 1976–1980, NHANES III: 1988–1994, and Current NHANES 1999–2006). [Heights and weights in the surveys are actual measurements.] There are four surveys between 1959 and 1998, and another four surveys between 1999 and 2006 which are so close to one another (and each have fewer observations) that they count as one survey for purposes of this analysis, making a total of five "effective" surveys. To ensure comparability both over time and with studies in other countries and to reduce uncontrolled heterogeneity (Rosenbaum, 2005) (through immigration, for example), we confined our analysis to non-Hispanic blacks and non-Hispanic whites born in the United States. (Henceforth, we drop the designation non-Hispanic for sake of brevity.) [The US-born criterion cannot be applied for NHES I. For NHES II and III, we assume that those with a birth certificate were US-born. Information on Hispanic ethnicity is only available for NHANES III and Current NHANES. Lack of information in earlier surveys does not constitute a major problem, though, inasmuch as Hispanics were not intentionally oversampled before NHANES III.] We limited the analysis to individuals aged 24–240 months [For some surveys, we had to calculate age in months at examination and birth year as in NHES I, only age in years is available. We have subjects only aged 18 or 19 years in our sample from this survey and assume that these individuals were 222 and 234 months old, respectively. Year of birth cannot be exactly determined for NHES I (only up to a ca. 4-year interval), NHANES III (ca. 23-year interval), and Current NHANES (ca. 2-year interval). In these cases, we assumed that all subjects were measured at the midpoint of the examination interval.] (N = 6,653 black girls, 14,326 white girls, = 6,611 black boys, and 14,972 white boys). [Ages at the ends of the period discussed is also implicitly available in the figure insofar as the early cohorts pertain to 19-year-olds while the latest cohorts to 2-year-olds. Hence, one can see that the sample contains ca. 120 19-year-olds in the early 1940s and about as many 18-year-olds. The number of 2-year-olds is about 150 and of 3-year-olds is about 120.] While whites constitute the majority of observations well into the early 1970s, the ethnicity composition is more balanced in the second half of the period considered due to oversampling of minority groups (Fig. 1).

TABLE 1. Characteristics of the five models analyzed

Model no.	Dependent variable	Method	Special remarks
1	BMI	BayesX	–
2	BMIZ	BayesX	Sampling weights
3	BMIZ	BayesX	–
4	Overweight	BayesX	Logisic
5	BMIZ	OLS	Robust s.e.

We estimate five models: (1) using the body mass index (BMI) as dependent variable, (2) the standardized BMI-for-age *Z*-scores using the CDC 2000 reference values standardized for age and gender (Kuczmarski et al., 2002) (BMIZ) using sampling weights, [The BMIZ values were calculated using the *EpiInfo* software (Kuczmarski et al., 2002).] (3) we also analyze BMIZ without sampling weights, (4) a binary variable for overweight, defined as a BMI-for-age value above the 85th percentile of the reference value, (5) we use ordinary least squares regression (OLS) on the BMIZ values with dummy variables for age and time (Table 1). [Some statisticians argue that one does not need sampling weights if the sample is stratified by gender and ethnicity insofar as age is controlled for (DuMouchel and Duncan, 1983; Gelman, 2007). Nonetheless, we do this part of the analysis with and without weights in order to demonstrate this point. The sample weight for our effective sample (all four groups combined) was adjusted such that it had a mean of 1 within each underlying survey.] We use a flexible approach (discussed below) for the first four models. To explore the sensitivity of the results to the model choice, we also estimate the trend in BMIZ values with OLS regression using dummy variables for age and birth cohort groups, a more conventional approach (Model 5). Each of the five models is estimated four times, i.e., separately by gender and ethnicity.

Admittedly, all dependent variables have some limitations: (1) and (5) assume that a one unit increase in BMI has similar implications for nutritional balance at all ages. Similarly, 2(a) and 2(b) treat a one standard deviation (σ) change at all ages identically, whereas in actuality children's BMIZ values may not be uniformly responsive to environmental effects at all ages. Furthermore, the obesity prevalence pertains, in the main, to the right tail of the distribution. [The reference values pertain to the whole US population of children including Hispanics as well as legal immigrants. There is a problem with the reference values themselves (used with dependent variables 2 and 4) in as much as up to age six they include values measured in 1988–94 by which date BMI values have increased, whereas for ages 7–19 they were constructed on the basis of data from 1963 to 1980. Thus, the reference values are not time-consistent as BMI values were increasing during that time period.] Nonetheless, using these various approaches should enable us to provide a sufficient overview of the trends in BMI values among children and adolescents during the course of the second half of the 20th century.

Another limitation of the estimates is that ages are not evenly distributed during the period considered. During the beginning of the period we have only older ages, whereas toward the end we have mainly younger children in the sample. For instance, for the year 2004 we have only 2-year-olds measured in 2006. This sample composition implies that we should consider the estimates at the ends as preliminary and subject to revision as more data become available. Nonetheless, we chose to include them insofar as they do enable us to provide some conjectures of future developments, even if these are tentative.

METHODS

We first use non-parametric regression models which enable us to estimate the shape of the trend flexibly by the data rather than being determined ex ante (Yatchew, 1998). [Hence, there is no need to assume, for example, that the BMI values increased linearly or as a polynomial.] The changes in BMI are composed of an age and a birth cohort trend effect. That is to say, in addition to a trend that pertains to all ages, children in some ages might have experienced greater (or less) increases in BMI values perhaps because of earlier maturation. We assume that these two effects are additive and estimate them using penalized cubic spline functions, thus smoothing the functional form at adjacent values of the independent variables (Lang and Sunder, 2003). We use BayesX, a "freeware" computer program (Brezger et al., 2005; Brezger and Lang, 2006). It estimates an intercept term [The two spline functions are restricted to have an average value of zero.] (γ) as well as the functional relationships between the dependent variable on the one hand and age (f_1) and the year of birth [Interaction effects are not useful to estimate, insofar as the data do not exist for all age birth year combinations. Therefore, there would be too many missing values.] (f_2) on the other:

$$\mathrm{BMIZ}_i = \gamma + f_1(\mathrm{AGE}_i) + f_2(\mathrm{BYEAR}_i) + \varepsilon_i \qquad (1)$$

Here, AGE is the age of the subject at the examination in months, BYEAR is the year of birth, f_1 and f_2 are spline functions to be estimated, and ε is an error term. The age effect captures the impact of earlier maturation on average (earlier onset of puberty) on the BMIZ values (Hermanussen et al., 2007; Rogol et al., 2002). We estimate this model two ways: (Models 2, 3) as well as with BMI as the dependent variable (Model 1).

In addition, we also estimate a logit model specification that uses a binary indicator variable for being overweight ($Y_i = \{0,1\}$), defined as an age-specific BMI value above the 85th percentile of the reference values [BMIZ-value larger than 1.04.] (Model 3). The probability of being overweight is assumed to be a nonlinear additive function of age and year of birth:

$$\Pr(Y_i = 1) = \frac{1}{1 + e^{-(\gamma + f_1(\mathrm{AGE}_i) + f_2(\mathrm{BYEAR}_i))}} \qquad (2)$$

With this model, we obtain the predicted probability of being overweight. Finally, we estimate the trends using OLS using binary variables for birth cohorts in order to ascertain the extent to which the results obtained with the above models are sensitive to the choice of method.

RESULTS

We present the estimated trends with the four models (Table 1) (Eqs. 1 and 2) and three dependent variables in Top Panels of Figures 2–5. The trends reflect averages (of the particular dependent variables) of all ages of children in our sample. [In each case we added the constant, γ, and the effect of an age of 150 months to the estimated values of f_2.] The trends are derived using the above model and are therefore smoothed by the splines. They are, in fact, quite smooth, but their rate of change, calculated as the first derivative of the trend lines do fluctuate quite a bit (Bottom Panels, Figs. 2–5). [The considerable fluctuations in the first derivatives are not caused by over fitting as this graph is not derived directly from the estimated regression.] This is not surprising as slight changes in the trends can bring about considerable changes in their first derivative.

We do not report the estimated values of the function f_1 in Eq. 1 insofar as the age effects are peripheral to this article. The three secular trends estimated by function f_2 in Eq. 1 and the trend in prevalence estimated by Eq. 2 are presented in Figures 2–5. Models 2 and 3 are almost identical for three of the ethnic/gender groups, diverging somewhat from one another only among black boys. The use of the sampling weights makes little difference because of stratification. [The secular trends estimated here are in addition the age effects, which affect all ages and are available from the authors upon request.]

Among white girls, the four trends are quite similar for the period under consideration until ca. 1990 when the trend in the raw BMI values diverges substantially from the other three trends (Fig. 2, top panel). The prevalence

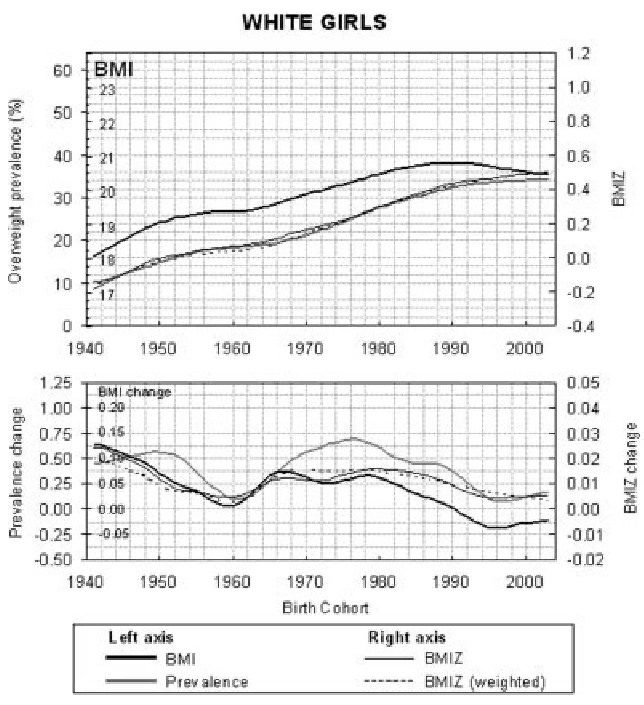


Fig. 2. Top panel: Trend in BMI and BMIZ values: function f_2 from Eq. 1; overweight prevalences estimated from Eq. 2; pertains to non-Hispanic US-born white girls ages 2–19 born 1941–2004. Bottom Panel: rate of change of the functions estimated in the top panel.

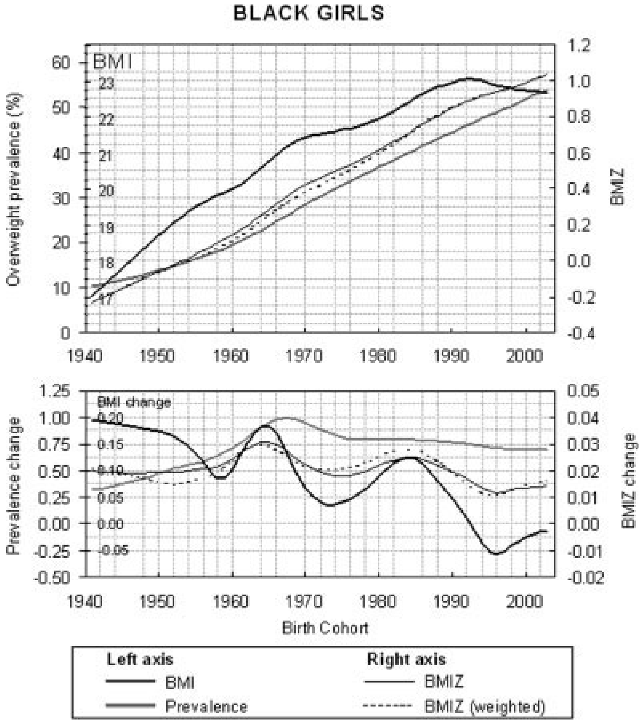


Fig. 3. Top panel: Trend in BMI and BMIZ values: function f_2 from Eq. 1; overweight prevalences estimated from Eq. 2; pertains to non-Hispanic US-born black girls ages 2–19 born 1941–2004. Bottom Panel: rate of change of the functions estimated in the top panel.

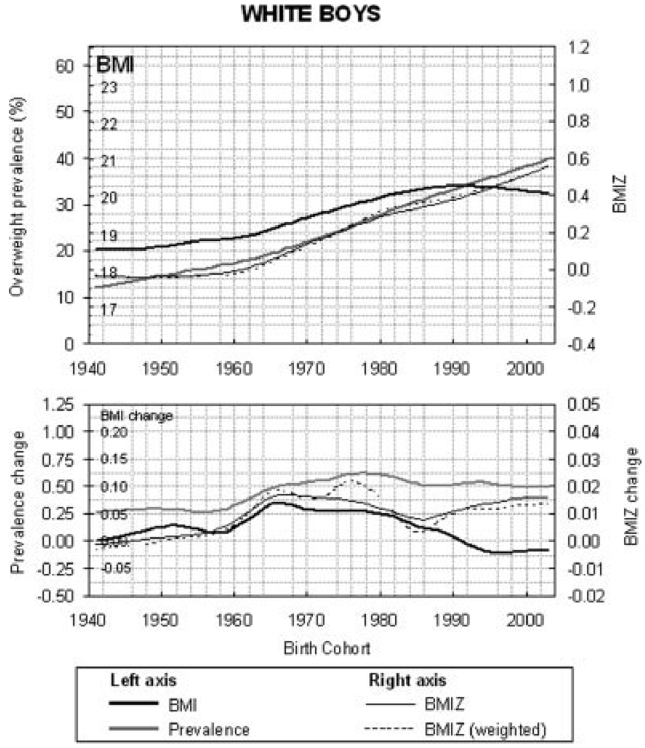


Fig. 4. Top panel: Trend in BMI and BMIZ values: function f_2 from Eq. 1; overweight prevalences estimated from Eq. 2; pertains to non-Hispanic US-born white boys ages 2–19 born 1941–2004. Bottom Panel: rate of change of the functions estimated in the top panel.

trend also flattens out somewhat compared to the BMIZ values (both weighted and unweighted) during the final decade considered. The increase in BMI was already under way at the very onset of the period considered, i.e., in the 1940s. The relatively high rate of change of BMI values at the outset tended to decelerate during the 1950s reaching nearly zero by ca. 1960 (Fig. 2, bottom panel). Thereafter, it increased according to all four models reaching a plateau in the 1970s. After ca. 1980 the rate of increase decreased, but the estimates diverge from one another somewhat. The rate of change of BMI and of the prevalence decreased most rapidly. The BMIZ values show the least decrease after 1980 while the rates of change of the raw BMI values actually become negative in the early 1990s.

Among black girls the two BMIZ estimates (Models 2 and 3) are virtually indistinguishable, as among white girls. The trends for black girls also suggest an early start in the increase in BMI values (Fig. 3, top panel) with the raw BMI values showing the largest rate of change in the 1940s (Fig. 3, bottom panel). Thereafter, there are some fluctuations in the rate of change in both the raw BMI values as well as the BMIZ values with peaks in the mid-1960s and mid-1980s. There was a tendency for the rate of change to decrease until the mid 1990s when it began to increase again slightly as among white girls. However, the rate of change of the prevalence peaked in the late-1960s and then declined but very slightly thereafter, not showing the cycles evinced by the other three estimates. It also does not indicate a decrease after 1990. The raw BMI values alone indicate a negative rate of change after ca. 1993.

The BMI values of white boys began to increase later than that of the girls. There were no meaningful changes until the late-1950s except for a slight increase in the prevalence (Fig. 4, top panel). However, by 1960 an upswing began for all indicators. The rate of change in the BMI and BMIZ (unweighted) values peaked in the mid-1960s and then declined slightly with the BMIZ values (weighted) showing a bit of a rebound in the mid-1980s (Fig. 4, bottom panel). For the overweight prevalence estimates the acceleration that began in the late-1950s continued until the late 1970s. The estimate for the raw BMI value alone declines after ca. 1990.

The four estimates for black boys differ among themselves in the beginning of the period under consideration, with BMIZ values changing very little while BMI and overweight prevalence increasing (Fig. 5, top panel). But the rate of change in all indicators increased by the 1950s and peaked in the mid-1960s (Fig. 5, bottom panel). The rate of change began to decrease by the late 1960s with a slight rebound in the mid 1980s. In the early 1990s, the four indicators diverged from one another again with the BMI and the BMIZ values (weighted) decreasing while the overweight prevalence continued to increase. Two of the rate of change indicators even reached negative values (BMI and BMIZ weighted).

BMI increased substantially between 1941 and 2004. The BMI values for white girls increased by 2.4, those of black girls by 5.6, of white boys by 1.5, and of black boys by 3.3 (without the age effects). Insofar as the confidence intervals were on the order of ca. ±0.3 all of the values are significantly different from one another. In terms of preva-

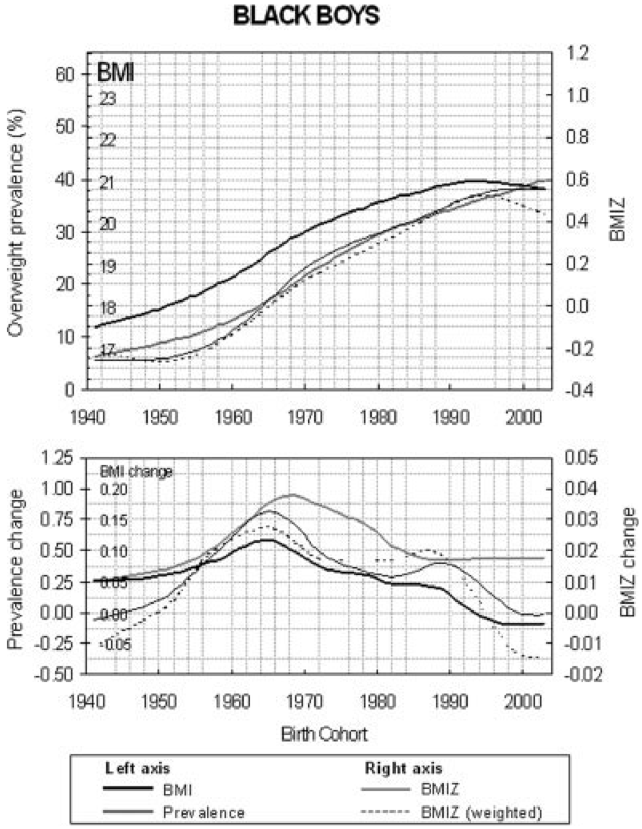


Fig. 5. Top panel: Trend in BMI and BMIZ values: function f_2 from Eq. 1; overweight prevalences estimated from Eq. 2; pertains to non-Hispanic US-born black boys ages 2–19 born 1941–2004. Bottom Panel: rate of change of the functions estimated in the top panel.

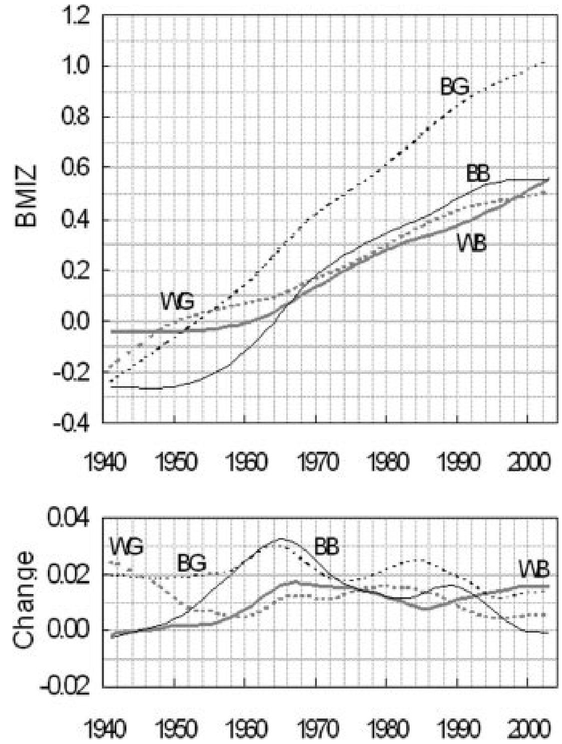


Fig. 6. Top panel: Trend in BMIZ values: function f_2 from Eq. 1; of non-Hispanic US-born black and white boys and girls ages 2–19 born 1941–2004 (Model 2b); Bottom Panel: rate of change of the functions estimated in the top panel.

lences, the increase was 24% points among white girls, 43% points among black girls, 28% points among white boys, and 34% points among black boys. BMIZ values increases among black girls by 1.3σ (95% CI: 1.16σ; 1.44σ) for black boys by 0.8σ, among white girls by 0.7σ, and among white boys by 0.6σ. The 95% confidence intervals around the point estimates of the BMIZ values are ~0.05σ for whites and about 0.07σ for blacks except at the ends of the period under consideration on account of the smaller number of observations. Hence, the BMIZ values of the black girls are statistically different from those of the other three groups and those of black and white boys are also statistically different from each other, although the difference pertains, in the main to the early part of the period under consideration. [While the "fit" of the models is low, considering that we do not control for genetic components (e.g. with BMI of parents), eating habits, or level of physical activity. The squared correlation between predicted and actual BMIZ scores ranges from 0.02 (white boys) to 0.10 (black girls). This might seem like a small amount but note that the proximate determinants of BMI values are omitted. Nonetheless the trends are significant.]

The BMIZ trends seem most plausible as the estimates based on the raw BMI values appear to be more influenced by the unbalanced nature of the sample in the beginning and end of the period. Moreover, the overweight prevalences pertain, in the main, to information in the right tail of the distribution. The comparison of the BMIZ scores for the four ethnic-gender groups shows how much more the BMI of black girls increased compared to the other groups (Fig. 6). The annual rate of increase was almost consistently above 0.02σ per annum and it did not fluctuate as much as that of the other groups (Fig. 6, bottom panel). The rate of change of BMIZ values of black females and white boys experienced an acceleration from the late-1950s to the mid-1960s, while that of black boys started somewhat earlier and that of white girls a bit later. After the mid-1960s, the rate of increase tended to be generally somewhat slower with a slight increase among white boys after the mid-1980s. After 1990, the rate of increase was slower among white girls and black boys than among black girls and white boys. The fluctuation in the rate of change was the greatest among black boys increasing from near zero to 0.03σ per annum in the 15-year period 1950–1965 and then declining back to zero again by 1999.

To explore the extent to which the results are sensitive to the choice of technique, we also analyzed the data using OLS regressions with dummy variables for age and for quinquennium of birth with BMIZ values as the dependent variable (Model 5, Fig. 7). This approach yields qualitatively similar results to the ones reported above. We report the results only for girls, though those of boys lead to identical conclusions. Circles indicate the estimates of the cohort effects and they are obviously quite similar to those obtained with Model 2. Admittedly, the dummy variable estimates fluctuate somewhat more than the values estimated with Model 2. [Age is in completed years. One could have defined these dummy variables in narrower or

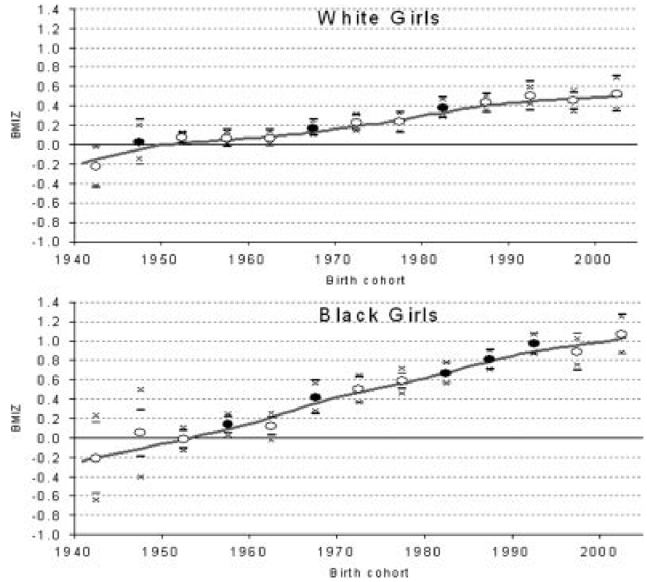


Fig. 7. Estimated trends of BMIZ scores of US-born girls by race, and birth cohort using OLS regression.

broader categories. It is clearly not desirable to have "too many" categories, as this would overfit the data and result in unrealistically abrupt changes; "too few" categories would also be a problem inasmuch as this might average out interesting patterns. This trade-off between flexibility and smoothness also exists with higher-order spline functions, so that one would have to specify a number of inner knots that determine the flexibility of the function (e.g., with the "GAM" routine available for STATA). If there is more than one such function, there is no straight-forward method to obtain an optimal combination of degrees of freedom or an optimal amount of smoothing for each function. An alternative to choosing a number of knots is to specify a relatively large number of inner knots in the first place (20 in our models) and to impose a penalty for abrupt changes of the slope of the function (Eilers and Marx, 1996). BayesX allows the joint estimation of coefficients for the spline basis functions and a term that governs the smoothness of the spline-modeled as the variance of the second derivative with weakly informative inverse gamma hyperpriors (with both parameters set to 0.001). As a result of the smoothing process, the effective degrees of freedom used range from 4.7 (black boys) to 7.0 (white girls) for the age spline function and from 4.2 (black girls) to 5.0 (white girls) for the birth year spline function. Hence, this specification could be considered more parsimonious than the approach with dummy variables in Figure 7. Brezger and Lang (2006) provide details on Bayesian inference with penalized splines.] This is, in fact, the reason why we prefer not to rely on ad-hoc dummy-variable specification in which, e.g., the definition of the start and end points of the cohort dummy variable can determine the size of the change. [If the number of observations had been greater, we could have used single year dummy variables for the birth cohort effects.] Hence, we favor the other models, as we believe that the assumption of a "smooth" trend is quite plausible in this application.

However, the OLS method does have the advantage of being able to calculate robust standard errors that account for clustering within the primary sampling units (PSU). [We use the "SVYREG" routine in STATA. PSU numbers were changed such that each PSU number in the combined dataset can originate only from one of the constituent surveys.] The robust standard errors are also reported in Figure 7 along with the usual OLS standard errors (i.e. under the random sampling assumption). To be sure, the PSU-adjusted confidence intervals are not meaningfully larger than the "standard" ones, so that we believe that this design effect does not play a major role in our context. Nonetheless, we can test if successive BMIZ-scores estimates are statistically significant from one another (using PSU-adjusted robust standard errors). In several cases we find significant differences (at the 10% level); these are indicated by filled circles (Fig. 7). In sum, the OLS method does not affect the conclusions crucially even if the fluctuations are somewhat greater than with the other models. However, our results are not artefacts of the method of analysis.

HISTORICAL COMPARISON

There are no national samples prior to the ones analyzed above, but there is some regional evidence on BMI values of white male youth going back to the birth cohorts of the 1850s in case of the West Point Cadets (Cuff, 1993; Komlos, 1987). Although obviously not representative of the population at large, they are the only such early source available on youth. These indicate that average BMI values were quite low in the middle of the 19th century and increased very little during the course of the second half of the century even among military cadets who were surely among the better situated members of the society (Fig. 8). In fact, about 90% of the cadet sample was below today's median reference value (Fig. 9). Another historical sample from The Citadel Military Academy in Charleston, SC, indicates that BMI values continued to be low for the remainder of the 19th century. In fact, BMI values tended to decline slightly toward the turn of the 20th century [Evidence from the military corroborates the decline in BMI values toward the end of the 19th century (Costa and Steckel 1997, p. 55).] (Coclanis and Komlos, 1995). This was also the case among a sample of convicts

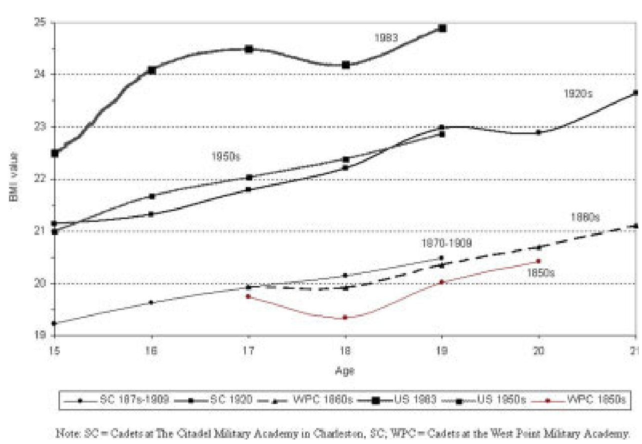
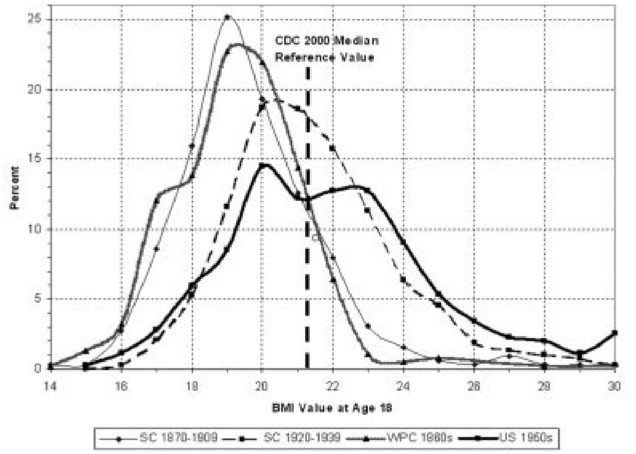


Fig. 8. BMI values of US-born white male youth, born ca. 1850–1983. [Color figure can be viewed in the online issue, which is available at www.interscience.wiley.com.]



Note: SC = Cadets at The Citadel Military Academy in Charleston, SC. WPC = Cadets at the West Point Military Academy.

Fig. 9. Distributions of BMI Values of White US 18-year-olds, 1860s–1950s.

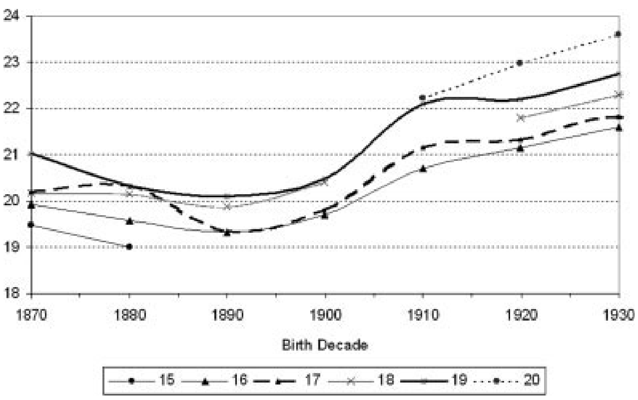


Fig. 10. BMI values of students attending The Citadel Military Academy in Charleston, SC.

from Pennsylvania (Carson, 2008, Racial differences in body mass indices at the time they were imprisoned in 19th Century Pennsylvania, unpublished manuscript). Not until the birth cohorts of the 1910s–1920s did BMI values increase substantially, at least in the South, but still remained well within the normal range (Fig. 10). The next substantial increase in BMI values occurred among boys in the mid-1950s. The increase in the 20th century came about not only by shifting the distribution to the right, but also by increasing the right tail of the distribution substantially (Fig. 9).

DISCUSSION

Our goal has been to identify the long-run trends in adiposity of US-born (white and black) children and youth born 1941–2004 (net of effects that are due to earlier maturation) rather than to analyze in detail their proximate determinants. Such a long-run perspective has not yet been estimated as the various NHANES samples have not been concatenated as in this analysis and have been analyzed exclusively using period effects rather than birth

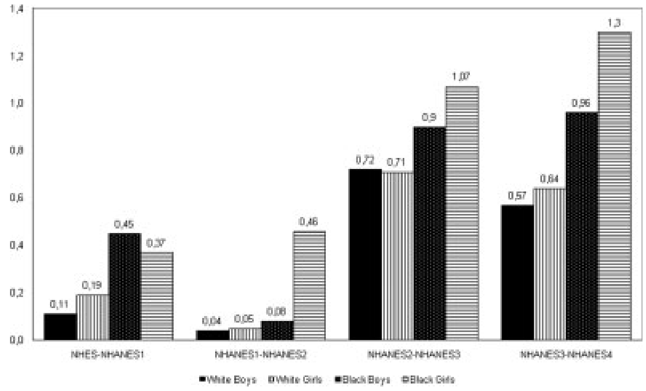


Fig. 11. Increases in BMI values between surveys using conventional methods.

cohort effects as above. Our approach seemed useful for interpreting the path of the obesity pandemic insofar as so ordered data exist for practically every year in the second half of the 20th century whereas analysis by measurement years provide only five cross-sectional observations from which four differences can be calculated so that important turning points are concealed. Indeed, using conventional analysis one could easily conclude that the major increases in BMI values started suddenly among those measured in the 1980s (Fig. 11). To be sure, there is some indication among blacks even using the conventional methodology that the adiposity epidemic was underway somewhat earlier, but this, too, has not been emphasized sufficiently. Furthermore, the conventional analysis generally disregards age effects, i.e., changes in the tempo of growth. Hence, using birth cohorts we can estimate the trend probably more accurately and in considerably greater detail insofar as we obtain ca. 58 estimates of BMI values instead of five. Consequently, there is a higher chance of identifying changes in the rate of increase as well as some factors associated with the trend.

Moreover, the ethnic differences in the spread of the BMI epidemic have also not been stressed unambiguously. For example, Troiano and Flegal (1998) did not find systematic variation of overweight with "race-ethnicity, income, or education," whereas Freedman et al., (2006) report that "Overall, black children experienced much larger secular increases in BMI, weight, and height than did white children." Hence, our findings support more the latter assertion rather than the former. [BMI values of black men exceeded that of whites also among Union Army soldiers (Costa, 2004).] Furthermore, that the increase in BMI of black females and possibly also that of white females was already under way to some extent among the 1940s birth cohorts seems to have completely eluded researchers hitherto. Admittedly, why the effect was greatest among black females in the 1940s is unclear (Fig. 6). To be sure, the small number of observations as well as the unbalanced age composition of the sample in the 1940s renders the results for this decade quite tentative and preliminary. [The result for the black females is slightly more convincing for this decade insofar as the rate of change in BMI remains similar for the rest of the century while that for white females first declines temporarily in the 1950s before rising again.] Moreover, it is uncertain just when during the life cycle the weight gain took place among those born, for example, in 1942 and measured as 19-year-olds in 1961. [We are not able to estimate period and cohort effects simultaneously with the current data set as we have only 5 surveys.] The increase in the 1940s might be associated with the war, insofar as the early cohorts were born at a time when nutritional availability was limited and could have gained in weight after the war's conclusion, but then why did the war not affect boys at all? [Cindy Fitch was kind enough to point out that the National School Lunch Program started in 1946 and could have improved the nutritional status of girls.] And why did the increase in BMI values continue among black girls in the 1950s? Thus, it would be premature to suppose that the increase in the BMI values among the 1940s female birth cohorts was brought about primarily by the end of war effort.

Nonetheless, it appears highly unlikely that the obesity pandemic appeared suddenly in the 1980s among American children as conventional analysis would suggest (Fig. 11) but has rather manifested itself slowly and persistently for an extended period of time beginning at least with the appearance of television and fast-food culture in the 1950s, but possibly earlier. In fact, the transition to post-industrial BMI values might well have been underway already in the 1920s with the introduction of radio broadcasting and the rapid spread of automobiles. Admittedly, this latter assertion is based on a regionally limited and socially selective sample of military cadets before nationally representative surveys became available. Nonetheless, the conjecture is worth entertaining for future research particularly as some scattered collaborating evidence does exist for boys [This is indicated by the BMI values of 6- and 10-year-old (black and white) boys, as well as those of 15-year-old white boys.] (Meredith, 1963). The "creeping" nature of the epidemic as well as its long-term persistence suggests that its fundamental roots are multifaceted and are arguably anchored in similarly slowly changing but persistent and irresistible impersonal socio-economic forces experienced by the US population in the 20th century.

The most obvious continuous forces exerted on individuals without countervailing support were associated with the major labor saving technological changes of the 20th century which included (inter alia) the industrial processing of food, the spread of fast-food culture, the use of automobiles, the introduction of radio and television broadcasting, [Television viewing has an additional effect through food and drink advertising that also affects food intake and obesity rates (Chou et al., 2007; Powell et al., 2007).] the increasing labor-force participation of women, and the IT revolution (Anderson et al., 2003; Bleich et al., 2007; Lakdawalla et al., 2005; Popkin, 2004). Combined with increasing affluence, these developments reinforced one another and led to the cultural transformation associated with the post-industrial nutritional revolution and a sedentary lifestyle (Cutler et al., 2003; Philipson and Posner, 2003; Lin et al., 2001). For example, the share of food expenditures spent on eating outside of the home increased from 24% in 1950 to 45% in 1995 (Offer, 2001, 2006; Guthrie, et al., 2002). ["The per-capita number of fast-food restaurants doubled between 1972 and 1997" (Chou et al. 2004, p 568), and the calories available for consumption increased by some 20% in the late 1980s and 1990s. In turn, the consumption of such energy-dense foods was associated with the increase in the number of hours worked by mothers (Anderson et al., 2003).]

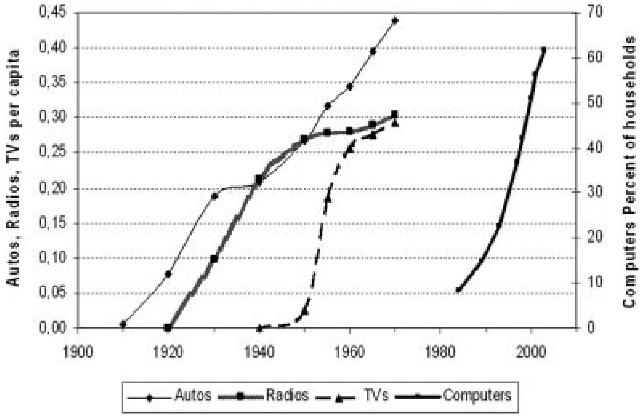


Fig. 12. The spread of major technologies in the 20th century in the US.

It is evident that the BMI values of all four (gender/ethnic) groups considered here in all of the four models estimated (in total 16 estimates, Figs. 2–5) accelerated in the mid-to-late 1950s at the time when the introduction and spread of TV-viewing in US households increased very rapidly [The acceleration is not evident in the OLS models (Figure 7) insofar as those are quinquennial averages that conceal the annual trends. However, Germaine Cornelissen-Guillaume has pointed out that exposure to magnetic fields of TV sets could be associated with hormonal imbalances (Salti et al. 2006).] (Chou et al., 2007, Fast-food restaurant advertising on television and its influence on childhood obesity, unpublished manuscript) (Fig. 12). Black boys were affected the most (Fig. 5). The rate of increase peaked in the mid-1960s and then decelerated somewhat although staying at a high level throughout the 1970s. Insofar as the trends in the BMI values of the four groups considered here are distinct to some extent, the technological changes mentioned above seems to have affected them somewhat differently. Nonetheless, this is the first time to our knowledge that the contemporaneous upswing in TV ownership and in the acceleration in children's BMI has been vividly identified.

On the other hand, the association of the IT revolution of the 1980s and 1990s on the rate of increase of BMI values was not as consistent as the spread of TV sets, only making a slight impact on white boys and to a limited extent also on black boys (of short duration). This is consistent with the fact that fewer black households own computers than white ones (black: 46%; white: 83%) (US Bureau of the Census, 2001). There might have been a substitution from one leisure activity to another both of which contributed to a sedentary life style without having a net effect. After ca. 1990, the rate of change of BMI values was slowing among white girls and black boys but remained higher among black girls and white boys. Ogden et al (2006) also report that the prevalence of overweight children ages 2–5 among white females remained unchanged between measurement years 1999 and 2004 and among black males between 2001 and 2004. We thereby corroborate this finding.

To be sure, in some details the patterns are somewhat model dependent primarily at the two ends of the period discussed probably because of the uneven age coverage in the sample. Hence, the deceleration in BMI values at the end of the period should be considered preliminary even though Ogden et al., (2008) suggest that there was "no statistically significant trend" in obesity between 1999 and 2006. [Especially since Ogden et al. (2006) conclude that "The prevalence of overweight among children and adolescents and obesity among men increased significantly during the 6-year period from 1999 to 2004."] At the beginning of the period the increase in BMI values and in prevalence tend to be faster among boys than those estimated by the BMIZ values, while at the end of the period the BMI values tend to decline faster in contrast to the results obtained by the other methods. Nonetheless, the five models basically do corroborate the substantial increase in adiposity during the period considered, its persistent nature and the special impact of TV viewing. All models also indicate that the BMI of black girls increased faster than those of the other groups and all of the estimates tend to corroborate the acceleration in BMI gain in the mid-1950s and early-1960s when in most cases the greatest rates of increases are found. This was also the time when the height of children tended to stagnate or decline implying that there might have been an association between these two phenomena (Komlos and Breitfelder, 2008). The rate of increase in the BMI gain tended to decline after ca. 1970. The estimated time trends remain essentially unchanged if we use a standard OLS regression instead of the non-parametric models.

Nationally representative data on BMI values were not collected prior to the samples reported above. However, regional evidence indicates that at least among a sample of mostly Southern white boys the increase in BMI values actually began among the 1910s and 1920s birth cohorts (Figs. 8–10). According to this evidence, the transition between early-industrial and post-industrial BMI values could well have begun with the cohorts born around World War I. The transition would have occurred at least among US white boys in two steps of approximately equal size: in the 1910s–1920s which was actually a healthy jump, i.e., an improvement in biological well being, and then another shift began among the mid-1950s birth cohorts which soon brought too many BMI values into the danger zone (Fig. 9). [However, there is also contradictory evidence: the BMI values of a sample of Union Army soldiers was the same as of those measured just after World War II (Costa 2004). Yet, this evidence is difficult to reconcile with data on the BMI values of West Point cadets who were to a considerable degree underweight (Figures 8 and 9). Gould (1869, vol 2, p. 403) gives the BMI values of native soldiers as 21.8 which is more in line with the values reported for the West-Point Cadets and the Citadel (Figures 7 and 9).] The conjecture appears plausible even if somewhat speculative that the incremental and persistent effects of changes in lifestyle associated with the multifaceted labor saving technological developments of the 20th century combined with cultural and nutritional revolutions might well be the underlying cause of the "creeping" obesity pandemic among US-born children and youth (Cutler et al., 2003; Philipson and Posner 2008) (Fig. 12). [Admittedly, we are unable to explain many of the patterns found. For example, we do not know why the rate of change of white girls' weight decreased in the early 1950s or why the weight trend among white girls was more similar to that of black boys than to that of the other groups after 1990.]

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