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Secular trends in the fat and fat-free components of body mass index in children aged 8–18 years born 1958–1995

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Abstract

Background—It is unknown whether the secular trend in childhood BMI reflects increases in fat-free mass as well as fat mass.

Methods—This study decomposed BMI trends in 488 participants in the Fels Longitudinal Study born between 1958–1995 and aged 8–17.99 years into their fat and fat-free components. Generalized estimating equations estimated birth year cohort (1958–1970–1971–1983–1984–1995) effects on 2208 observations of BMI, fat mass index (FMI = fat mass (kg)/height (m)^2) and fat-free mass index (FFMI = fat-free mass (kg)/height (m)^2).

Results—BMI in boys increased across cohorts, with those born between 1984–1995 being 2 kg/m^2 larger than those born between 1958–1970 ($p = 0.001$) and increases in FMI were highly significant (p-values < 0.001). FFMI did not differ by cohort. In girls, there was a significant advantage in BMI (1.2 kg/m^2) and FFMI (0.8 kg/m^2) of the 1984–1995 cohort compared to the 1971–1983 cohort (p-values < 0.05).

Conclusions—Because the long term trend in childhood BMI in boys appears to be driven by an increase in total body adiposity, evidence is provided to support current knowledge on the predicted deleterious long-term consequences of the childhood obesity epidemic in boys. Research is needed to confirm whether recent changes in BMI in girls are due to increases in fat-free mass resulting from changes in behaviour and lifestyle not yet manifest in boys.

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WJ had full access to all of the data in the study and takes responsibility for the integrity of the data and the accuracy of the analysis

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WJ designed research, performed statistical analysis and wrote the paper. EWD was responsible for project conception, development of the overall research plan and had primary responsibility for final content. WCC and SAC interpreted data and revised the manuscript. All authors provided critical revision for important intellectual content and approved the final version.

Keywords

Secular trend; obesity epidemic; body composition; hydrodensitometry

INTRODUCTION

The childhood obesity epidemic in the US became evident between the second and third National Health and Nutrition Examination Surveys (NHANES II and III) in 1976–1980 and 1988–1994, respectively (Troiano et al. 1995). It is assumed that this secular trend in increasing childhood body mass index (BMI) is due to increased fat mass, based on the strong correlation of BMI with fat mass (Freedman et al. 2005) and increases in the prevalence of adiposity-related co-morbidities (Gregg et al. 2005). In children, however, BMI is strongly associated with fat-free mass (Freedman et al. 2005), particularly in boys (Demerath et al. 2006). If secular trends in BMI reflect increases in both fat mass and fat-free mass, our interpretation of the long-term consequences of childhood obesity could be altered, because increases in fat-free mass may suggest that some part of the upward trend in BMI could have a protective effect on disease risk (Dulloo et al. 2010).

A recent paper (Sun et al. 2012) examined trends in body composition in children born between 1960–1999 in the Fels Longitudinal Study (Roche 1992). Differences in mean BMI, percentage body fat (%fat) and a fat-free mass index (FFMI = fat-free mass/height^2) between birth year decades were tested within yearly chronological age groups; a trend in increasing BMI and %fat was observed in the absence of a trend in increasing FFMI. In this analysis, body composition data were pooled within individuals from two different methods: hydrodensitometry and dual energy x-ray absorptiometry (DXA), beginning in 1990. It has previously been shown in the Fels Longitudinal Study that there is relatively poor agreement between %fat obtained from hydrodensitometry and %fat from DXA, which 'makes it unacceptable to interchange these methods for individuals' (Wellens et al. 1994, pp. 552). To ameliorate this issue, the recently published paper (Sun et al. 2012) used conversion equations to adjust the hydrodensitometry values to obtain similar means to the DXA values, but these equations have not been validated in any publication. The concern over pooling data from two different methods which were implemented over different periods of the study is that assessment of secular trends may be biased by the known intra-method differences.

Our research group recently described trends in the BMI growth curve of children born between 1929–1999 in the Fels Longitudinal study (Johnson et al. 2012). Shifts in BMI growth in boys and girls were apparent starting in the 1970s and mainly affected after the BMI rebound (which occurs at ~5 years of age). Here, we use extensive body composition data from a single method (hydrodensitometry) to decompose the secular trend in childhood BMI into its fat and fat-free components. Further, we use an analysis approach that did not rely on extensive multiple testing as in the previous study, which may have inflated type-one error.

METHODS

The sample comprised 488 European-American participants (248 boys; 240 girls) in the Fels Longitudinal Study (Roche 1992) born between 1958–1995 and aged 8–17.99 years at assessment. Body composition was assessed using hydrodensitometry, from which body density was measured and total body fat mass and fat-free mass (kg) were computed (Lohman 1986) and weight and height measured. In total, 2208 assessments with complete data were available, with an average of 4.5 per participant (range 1–11) over an average of 4.4 years (range 0–10). All procedures were approved by Wright State University Institutional Review Board.

BMI was computed as weight (kg)/height $(m)^2$; fat mass and fat-free mass were also divided by height $(m)^2$ to obtain fat mass index (FMI) and FFMI, the fat and fat-free components of BMI. We initially modelled individual BMI, FMI and FFMI growth curves using fractional polynomial mixed effects modelling (Long and Ryoo 2010), but were not confident in the fit of the models when trying to incorporate birth year effects on the population average curves. For this reason and because this preliminary analysis showed that the birth year effect on each dimension did not differ significantly across the age range being studied (data not shown), we used a more parsimonious approach that can effectively incorporate serial data without having to unnecessarily model the trait (and the birth year effect on that trait) as a function of age.

Generalized estimating Equations (GEE) were fitted using GENMOD (SAS Institute, Cary, NC) specifying an autoregressive correlation structure to account for the non-independence of observations in the serial data (Hanley et al. 2003). The outcomes were BMI, FMI, FFMI and also height, thereby allowing us to comment on whether any trends in body composition variables occurred in the absence or presence of trends in height. BMI and height *Z*-score outcomes according to the CDC 2000 reference (Kuczmarski et al. 2000) were tested in additional models to allow the reader to interpret any observed secular trend relative to the growth chart currently used for this age range in the US. The exposure was birth year, categorized into tertiles to create three 12-year birth cohorts: 1958–1970 (referent group), 1971–1983 and 1984–1995. *P*-values from two degrees-of-freedom Chi-squared tests of the null hypothesis that both birth year cohort estimates (1971–1983 vs 1958–1970 and 1984–1995 vs 1958–1970) were equal to zero were used to assess the significance of the overall trends. In addition, models were the referent group was set to 1971–1983 were used to estimate betas for the 1984–1995 vs 1971–1983 comparisons. All models were stratified by sex and adjusted for exact age at measurement (centred to mean 13 years of age) and a birth cohort-by-age interaction term. Estimated least-square means are presented to adjust for the slight difference in mean age of children in each birth cohort. Model fit was assessed using the Quasi-likelihood under the Independence model Criterion (QIC), which is analogous to the Akaike's Information Criterion statistic used for comparing models' fit with likelihood based methods and also the QICu, which adds a penalty for the number of parameters and will approximate the QIC when the GEE is correctly specified.

RESULTS

Table I shows the parameter estimates from the GEEs. In all instances, the QICu approximated the QIC, thereby demonstrating that the GEEs were correctly specified to handle the serial observations. BMI in boys increased across birth cohorts, with those born between 1984–1995 being 2 kg/m^2 larger than those born between 1958–1970 ($p = 0.001$), representing an increase of 0.5 *Z*-scores on the CDC 2000 reference. Increases in FMI were highly significant (*p*-values < 0.001) and of only marginally smaller effect size. FFMI did not differ by birth cohort (*p*-values > 0.08). In girls, there were no significant associations of birth cohort with any outcome, with the exception of their being a significant advantage in BMI (1.2 kg/m^2) and FFMI (0.8 kg/m^2) of the 1984–1995 cohort compared to the 1971–1983 cohort. A stratified analysis by age group (8–12.99 and 13–17.99 years) produced similar results, as did using indices computed for each sex to be independent of height (data not shown).

Our study design meant that the summation of the beta estimates for FMI and FFMI for any given birth cohort approximated the respective beta estimate for BMI. The results therefore show the partitioning of changes in BMI into their fat and fat-free compartments. For example, we can be fairly confident that the 2 kg/m^2 larger BMI of boys born between 1984–1995 compared to boys born between 1958–1970 was more than 80% due to an increase in fat-mass (i.e. FMI beta 1.652/BMI beta 1.970 * 100 = 83.9%) and less than 20% due to an increase in fat-free mass (i.e. FFMI beta 0.202/BMI beta 1.970 * 100 = 10.3%). This change occurred despite the boys born between 1984–1995 being more than 2 cm taller than boys born between 1958–1970 ($p = 0.023$).

DISCUSSION

In this sample, a positive long term secular trend in BMI in boys was primarily due to increasing adiposity, despite the presence of a trend in increasing height. In girls, the trend in BMI was only apparent in more recent birth years and was primarily due to an increase in fat-free mass not any change in fat-mass. It is noteworthy that the magnitude of the differences in BMI presented here are not unusual compared to published NHANES data (Odgen et al. 2004). Other publications investigating how childhood body composition has changed during the second half of the 20th century have typically compared data at one time point to some earlier born reference population (Ruxton et al. 1999; Wells et al. 2002); this approach cannot assess a graded birth year or cohort effect on body composition and is subject to findings attributable to differences between the study sample and reference population. The present paper, for the first time, decomposes trends in BMI over a period of time encompassing the immediate, the emerging and the established obesity epidemic environments into their fat and fat-free components using directly measured body composition data.

Diet and physical activity levels in these Fels Longitudinal Study children were not systematically examined, thus it is not clear what factors were responsible for the observed secular trends. Our understanding of the effects of how environmental changes over time affect underlying processes of body composition needs greater attention. Another limitation

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is a relatively small sample (after stratification by sex and birth year cohort) composed exclusively of European-American children born in southwestern Ohio which limits generalizability of the results to other ethnic groups and the nation at large. Because the long term trend in childhood BMI in boys appears to be driven by an increase in total body adiposity, we provide evidence to support current knowledge on the predicted deleterious long-term consequences of the childhood obesity epidemic in boys. Research is needed to confirm whether recent changes in BMI in girls are due to increases in fat-free mass resulting from changes in behaviour and lifestyle not yet manifest in boys.

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Table I

Generalized estimating equations[a] (GEE) testing differences by birth cohort in childhood body mass index, fat mass index, fat-free mass index and height.

	Boys (n = 248)			Girls (n = 240)		
	Birth year 1958–1970	Birth year 1971–1983	Birth year 1984–1995	Birth year 1958–1970	Birth year 1971–1983	Birth year 1984–1995
n children	81	88	79	85	78	77
n observations	283	576	294	265	483	307
Body mass index (kg/m^2)						
Least-squares mean (SE)	18.50 (0.23)	19.31 (0.38)	20.49 (0.47)	19.67 (0.30)	19.39 (0.28)	20.65 (0.44)
Beta (SE) *p*-value	referent	0.820 (0.431)	1.970 (0.507)	referent	−0.264 (0.413)	0.954 (0.525)
Beta (SE) *p*-value		*0.057*	*< 0.001*		*0.523*	*0.069*
		referent	1.150 (0.586)		referent	1.218 (0.514)
			0.050			*0.017*
p-value for overall trend[b]	0.001			0.064		
GEE fit criteria[c]: QIC, QICu	1075.5, 1159.0			1074.9, 1061.0		
Body mass index *Z*-score[d]						
Least-squares mean (SE)	−0.23 (0.09)	−0.17 (0.13)	0.27 (0.12)	0.14 (0.09)	0.06 (0.09)	0.31 (0.12)
Beta (SE) *p*-value	referent	0.061 (0.159)	0.506 (0.154)	referent	−0.076 (0.133)	0.166 (0.151)
Beta (SE) *p*-value		*0.700*	*0.001*		*0.569*	*0.274*
		referent	0.443 (0.173)		referent	0.229 (0.144)
			0.010			*0.111*
p-value for overall trend[b]	0.003			0.267		
GEE fit criteria[c]: QIC, QICu	1176.1, 1159.0			1076.5, 1061.0		
Fat mass index (kg/m^2)						
Least-squares mean (SE)	2.38 (0.14)	3.43 (0.27)	4.08 (0.37)	4.71 (0.21)	4.73 (0.21)	5.13 (0.31)
Beta (SE) *p*-value	referent	1.045 (0.302)	1.652 (0.385)	referent	0.046 (0.296)	0.422 (0.371)
Beta (SE) *p*-value		< *0.001*	*< 0.001*		*0.876*	*0.255*
		referent	0.607 (0.446)		referent	0.376 (0.370)
			0.174			*0.310*
p-value for overall trend[b]	< 0.001			0.498		
GEE fit criteria[c]: QIC, QICu	1176.3, 1159.0			1075.8, 1061.0		

	Boys (n = 248)			Girls (n = 240)		
	Birth year 1958–1970	**Birth year 1971–1983**	**Birth year 1984–1995**	**Birth year 1958–1970**	**Birth year 1971–1983**	**Birth year 1984–1995**
Fat-free mass index (kg/m^2)						
Least-squares mean (SE)	16.09 (0.17)	15.80 (0.19)	16.25 (0.22)	15.03 (0.16)	14.67 (0.15)	15.52 (0.21)
Beta (SE) *p*-value	referent	−0.275 (0.247)	0.202 (0.266)	referent	−0.371 (0.219)	0.458 (0.261)
Beta (SE) *p*-value		*0.266*	*0.447*		*0.090*	*0.079*
		referent	0.477 (0.278)		referent	0.829 (0.257)
			0.086			*0.001*
p-value for overall trend[b]	0.224			0.007		
GEE fit criteria[c]: QIC, QICu	1173.7, 1159.0			1072.5, 1061.0		
Height (cm)						
Least-squares mean (SE)	157.76 (0.73)	158.55 (0.68)	160.11 (0.72)	153.52 (0.79)	152.23 (0.78)	155.18 (0.81)
Beta (SE) *p*-value	referent	0.817 (0.994)	2.302 (1.013)	referent	−1.390 (1.126)	1.380 (1.140)
Beta (SE) *p*-value		*0.411*	*0.023*		*0.217*	*0.226*
		referent	1.485 (0.972)		referent	2.770 (1.111)
			0.127			*0.013*
p-value for overall trend[b]	0.075			0.053		
GEE fit criteria[c]: QIC, QICu	1173.2, 1159.0			1075.7, 1061.0		
Height Z-score[d]						
Least-squares mean (SE)	0.21 (0.10)	0.35 (0.10)	0.43 (0.09)	0.16 (0.10)	0.27 (0.11)	0.37 (0.11
Beta (SE) *p*-value	referent	0.139 (0.140)	0.224 (0.135)	referent	0.111 (0.147)	0.213 (0.145)
Beta (SE) *p*-value		*0.321*	*0.100*		*0.449*	*0.141*
		referent	0.090 (0.134)		referent	0.113 (0.156)
			0.505			*0.467*
p-value for overall trend[b]	0.147			0.296		
GEE fit criteria[c]: QIC, QICu	1173.9, 1159.0			1074.5, 1061.0		

[a]GEE models were fitted using GENMOD (SAS Institute, Cary, NC), specifying an autoregressive correlation structure, and were adjusted for exact age at measurement (centred to mean 13 years of age) and a birth cohort-by-exact age at measurement interaction term;

[b]The *p*-value for the overall trend is a two degrees-of-freedom test of the null hypothesis that both birth year cohort estimates (1971–1983 vs 1958–1970 and 1984–1995 vs 1958–1970) are equal to zero;

[c] GEE fit criteria: the Quasi-likelihood under the Independence model Criterion (QIC) is analogous to the Akaike Information Criteria statistic used for comparing models fit with likelihood based methods; the model with the smaller statistics is preferred. QICu adds a penalty to the quasi-likelihood for the number of parameters in the model and will approximate the QIC when the generalized estimating equation is correctly specified;

[d] *Z*-scores were calculated according to the CDC 2000 reference (Kuczmarski et al. 2000).