NBER WORKING PAPER SERIES

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Working Paper 19143
http://www.nber.org/papers/w19143

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
June 2013

Research for this paper was supported by grant #65068 from the Robert Wood Johnson Foundation to the National Bureau of Economic Research The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

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NBER Working Paper No. 19143
June 2013
JEL No. I1,I18

ABSTRACT

In this paper, we examine the effect of food prices on clinical measures of obesity, including body mass index (BMI) and percentage body fat (PBF) measures derived from bioelectrical impedance analysis (BIA) and dual energy x-ray absorptiometry (DXA), among youths ages 12 through 18. The empirical analyses employ data from various waves of the National Health and Nutrition Examination Survey (NHANES) merged with several food prices measured by county and year. This is the first study to consider clinically measured levels of body composition rather than BMI to investigate the effects of food prices on obesity among youths. We also examine whether the effects of food prices on body composition differ by gender and race/ethnicity. Our findings suggest that increases in the real price of one calorie in food for home consumption and the real price of fast-food restaurant food lead to improvements in obesity outcomes among youths. We also find that an increase in the real price of fruits and vegetables has negative consequences for these outcomes. Finally, our results indicate that measures of PBF derived from BIA and DXA are no less sensitive and in some cases more sensitive to the prices just mentioned than BMI.

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1. Introduction

The prevalence of childhood obesity has increased at an alarming rate in the U.S. over the last three decades. Since the mid-1970s, the proportion of children aged 12-19 who are obese has grown from 5.0 to 18.1 percent and has grown more rapidly among non-Hispanic black adolescents than among Hispanic or non-Hispanic white adolescents (Ogden et al. 2010). The growing prevalence and the racial/ethnic disparities in childhood obesity are of major concern to public health, given that obese children are more likely to develop health problems, such as high blood pressure, hypertension, gallbladder disease, and Type 2 diabetes as early as adolescence (Serdula et al. 1993; Freedman et al. 1999; 2007; Hill, Catenacci, and Wyatt 2006). Furthermore, obese children are more likely to experience negative long-term psychological and labor market outcomes ranging from poor self-esteem and depression to discrimination and lower wages (Daniels 2006; Mocan and Tekin 2011; Dietz 1998; Strauss 2000). Wang and Dietz (2002) estimate that hospital expenditures related to childhood obesity rose from $35 million in the late-1980s to $127 million (in 2001 constant dollars) in the late-1990s. Both the Institute of Medicine (2004) and *Healthy People 2020* (U.S. Department of Health and Human Services 2010) identify the prevention of childhood obesity, particularly among disadvantaged groups, as a top public health priority.

Public interventions for improving child and adolescent health typically take the form of policies that limit access and provide price incentives and disincentives (Grossman 2005). Raising price through taxation has been shown to be highly effective at reducing substance use among adolescents (e.g., Grossman 2005; Brownell and Frieden 2009; Engelhard, Garson, and Dorn 2009). Likewise, selective applications of taxation and subsidies may shift food consumption away from unhealthy food and towards healthier alternatives (Cawley 2010; Powell and Chaloupka 2009). The falling real cost of food has been suspected of being a contributing factor in the recent epidemic of obesity (e.g. Lakdawalla and Philipson 2002; Chou, Grossman, and Saffer 2004; Chou, Rashad, and Grossman 2008). In general, empirical studies that recently examined the effects of prices on obesity (e.g., Chou, Grossman, and Saffer 2004; Rashad, Chou, and Grossman 2006; Chou, Rashad, and Grossman 2008; Powell and Yanjun 2007; Powell 2009; Powell, Chaloupka and Bao 2007; Auld and Powell 2009; Sturm and Datar 2005) found larger and more significant effects than studies that examined the effects of food taxes (Powell, Chriqui, and Chaloupka 2009; Fletcher, Frisvold, and Tefft 2010).

These studies typically attach location and year-specific prices or taxes to a variety of micro data sets and further include location and year fixed effects in regression analyses. The geographic unit of analysis is mostly states, but several studies focus on counties, cities, or even zip codes. These price variables usually include price indices of meals in fast-food and full-service restaurants, and an index of the price of food prepared at home. Prices of foods for consumption at home are decomposed into prices of foods with low energy densities, defined as low calories per pound of consumption (for example, fruits and vegetables) and those with high energy densities (for example, fast food). There is reasonably consistent evidence that fruit and vegetable prices, particularly non-starch variety, are associated with lower weight outcomes while fast-food prices are associated with higher weight outcomes for the adolescent population (Powell et al. 2013). These effects tend to be larger for minorities, children in lower-income families, and children whose mothers have less than a high school education. Some, but not all, of these results are based on BMI-measures of obesity calculated from self- or parental-reports of height and weight. An obese child or youth is classified as one having a BMI at or above the 95^{th} percentile based on age- and gender-specific growth charts.

However, none of these results are based on clinically measured levels of body fatness. BMI, defined as weight in kilograms divided by height in meters squared, is easy to calculate and readily available from many social science datasets but its reliability for use in epidemiological studies has come into question recently. Some of the weak or mixed results found by studies using BMI may be due to its limited ability to correctly distinguish body fat from lean body mass (e.g., Yusuf et al. 2004, 2005; Romero-Corral et al. 2006, 2007). Since it is body fat (and not fat-free mass) that is responsible for the detrimental health effects of obesity, several studies caution against a sole reliance on BMI and point to a need for using direct measures of body composition in obesity studies (e.g. Smalley et al. 1990; Romero-Corral et al. 2006).

These concerns of reliability are particularly relevant for children and adolescents due to the gender differences in physical growth as well as the gender and racial differences in the association of BMI with a child's body fatness (e.g. Daniels et al. 1997). Consequently, several studies tested whether measures other than BMI can be used as valid measures for the detection of the degree of obesity in obese children and adolescents. Widhalm et al. (2001) used regression methods to assess the relationship between the percentage body fat and BMI along with several demographic characteristics from a sample of 105 obese boys and 99 obese girls. The authors concluded that BMI provides only limited insight to the degree of obesity for children ages 10 and over. Skybo et al. (2003) recommend the use of body fat percentage for identification of overweight status in school-age children.

Another related limitation of the previous studies is using BMI calculated from self-reported values of height and weight may induce its own bias. Previous studies show considerable evidence of misreporting in weight and height (Rowland 1989; Gorber et al. 2007). In an effort to determine the degree of agreement between self-reported and measured values of height and weight, Gorber et al. (2007) reviewed 64 studies and conclude that there is evidence for under-reporting for weight and BMI and over-reporting for height that varies between men and women.

In this paper, we use clinically obtained body composition measures to conduct a comprehensive analysis of the effects of various food prices on body fatness among youths ages 12 through 18 and compare the sensitivity of our findings against results using BMI. This is the first study to consider clinically measured levels of body composition to investigate the effects of food prices on body fatness among youths. The body composition measure that we employ is the percentage body fat (PBF).[1] The PBF measure is derived from three separate sources, two of which rely upon bioelectrical impedance analysis (BIA) and one of which relies upon dual energy x-ray absorptiometry (DXA). We also employ clinically measured height and weight to estimate the effects of prices on BMI. We employ data from the restricted-use versions of NHANES to merge various county-level time-varying price variables.

Our findings have implications for the optimal targeting of public policies designed to prevent or reduce childhood obesity, including the extent to which changes in farm, tax, and subsidy policies might affect consumption patterns. Furthermore, the precision of alternative body fat measurements for the quantification of body fat and the correct identification of the degree of obesity in youths are important for the assessment of risk factors associated with obesity and its detection among youths. The rest of the paper is laid out as follows. We discuss our data and the way body fat measures are constructed in Section 2. We describe the empirical methodology in Section 3 and summarize our results in Section 4. We provide a brief conclusion in Section 5.

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[1] Note that $\text{PBF} \equiv 100 * \frac{\text{BF}}{\text{W}} \equiv 100 * \frac{\text{W} - \text{FFM}}{\text{W}}$, where W is weight, BF is body fat and FFM is fat-free mass.

2. Data

The data for the empirical analyses are drawn from several waves of the National Health and Nutrition Examination Survey (NHANES), a program of the National Center for Health Statistics (NCHS).[2] The NHANES is a series of cross-sectional nationally representative surveys designed to assess the health and nutritional status of adults and children in the U.S. The surveys are unique in that they combine interviews and physical examinations administered by trained nurses and technicians. The NHANES is ideal for the purposes of our paper because it is the only set of nationally representative health surveys that contains measures of body composition (body fat and fat-free mass) or information necessary for calculating these measures. In the NHANES surveys, these measures were obtained by using readings from BIA and DXA. In this paper, we use data from NHANES 1999-2000, NHANES 2001-2002, and NHANES 2003-2004 since all three measures of obesity outcomes we use are available only in these rounds.

2.1 Construction of Measures of Percentage Body Fat and Obesity

Body composition has been used for a long time by nutritionists and physiologists for the purpose of studying nutrition, physical growth, and physical performance (Forbes 1999). However, recent improvements in clinical measurements and the rising tide of obesity have led to a renewed interest in body composition. In multivariate analyses, body composition has been shown to be significantly better at explaining individual variations in strength, health, and physical performance than body size (Björntorp 2001; Institute of Medicine 2004). Furthermore, it has been demonstrated that fat-free mass (FFM) has a positive effect on health and physical performance, while body fat (BF) has a negative effect (e.g., Heitmann et al. 2000; Allison et al. 2002). One potential obstacle to wide adoption of body composition in studying obesity is that it is considerably more difficult and costly to obtain than BMI. However, advancements in various

[2] See http://www.cdc.gov/nchs/nhanes.htm for more information.

measurement technologies have reduced the cost of obtaining measures of body composition considerably, which has led to their incorporation in the NHANES.

Advancements in measurement technology have resulted in the development of alternative methods of measuring obesity and physical fitness, including body composition derived from bioelectrical impedance analysis (BIA) and dual energy x-ray absorptiometry (DXA).[3] Although clinical researchers have been using these alternative measures for some time, social scientists have only recently begun to take advantage of them in studying obesity, primarily due to the lack of clinical data in large-scale social science datasets. Exceptions include Burkhauser and Cawley (2008), Johansson et al. (2009), and Wada and Tekin (2010), who studied labor market outcomes by using direct measurements or developing a method for imputing body composition in social science datasets. In a recent paper, Grossman, Tekin, and Wada (2012) examine the effects of fast-food restaurant advertising on television on the body composition of adolescents as measured by percentage body fat and assess the sensitivity of these effects compared to using conventional measures of obesity based on BMI.

One particular methodology developed by clinical investigators to measure body composition is based on bioelectrical impedance analysis (BIA) (Kushner et al. 1990; Roubenoff et al. 1995; Sun et al. 2005; Chumlea et al. 2002). In BIA, body composition is estimated by measuring the electrical resistance of a body to a weak electrical current (National Institutes of Health 1994). FFM registers a lower electrical resistance due to its high water content, but in contrast, BF does not conduct electricity well (Chumlea et al. 2002; Sun et al. 2005). The observed electrical resistance is then converted into FFM by entering it into a predetermined equation obtained from a multiple regression analysis along with a set of easily acquired

[3] See http://www.cdc.gov/nchs/data/nhanes/bc.pdf for more information.

characteristics of individuals such as weight, height, age, and gender. Once FFM is predicted, BF is computed from the following identity:

$$BF \equiv W - FFM. \tag{1}$$

We make use of two alternative measures of FFM contained in the NHANES 1999-2000, NHANES 2001-2002, and NHANES 2003-2004. One of these measures was obtained from a prediction equation developed by Chumlea (2002) for particular use with the NHANES III, while the other was obtained from a prediction equation developed by Boileau (1996). For males, Chumlea's prediction equation is

$$FFM = -10.678 + 0.262\ \text{weight} + 0.652\ \text{height}^2/\text{resistance} + 0.015\ \text{resistance}, \tag{2}$$

where resistance is obtained from BIA. The corresponding equation for females is

$$FFM = -\ 9.529 + 0.168\ \text{weight} + 0.696\ \text{height}^2/\text{resistance} + 0.016\ \text{resistance}. \tag{3}$$

Boileau's equation, which he developed specifically to measure FFM in children ages 8 through 16, is

$$FFM = 4.138 + 0.657\text{height}^2/\text{resistance} + 0.16\ \text{weight} - 1.31\ \text{male}, \tag{4}$$

where male is a binary indicator that takes on the value of 1 if the individual is a male, and 0 otherwise.

Finally, we constructed PBF measures of obesity from dual-energy x-ray absorptiometry (DXA) as an alternative to the two BIA-based measures of PBF described above. As opposed to BIA, DXA does not rely on a specific prediction equation and yields direct measures of body fat and fat-free mass. It is also one of the most widely adopted methods of measuring body composition (Centers for Disease Control, 2008). DXA has long been used as a method to measure bone mineral content and bone mineral density and considered to be a highly reliable method because of its precision, accuracy, and low radiation exposure (Njeh et al. 1999; Wahner et al. 1994; World Health Organization 1994; Genant et al. 1996).[4] It is also increasingly being used as a criterion method for body composition assessment for children (Cameron et al. 2004; Eisenmann et al. 2004; Elberg et al. 2004; Frisard et al. 2005; Okasora et al. 1999; Lazzer et al. 2008; Eisenkolbl et al. 2001). Recent scientific developments in the DXA hardware along with fan-beam technology have led to new software development for a body composition assessment, which has allowed its incorporation into large surveys (Tylavsky et al. 2003).

In DXA, a complete body scan is administered with two low dose x-rays absorbed at different rates of energies by bone and soft tissue mass. The participants are positioned supine on the tabletop with their feet in neutral position and hands flat by their side. Each administered scan of the NHANES subjects was analyzed by the Radiology Department at the University of California, San Francisco using special software, standard radiologic techniques, and specific protocols developed for the NHANES to produce assessments of various body components, such as bone mineral content and density, fat mass, lean mass with and without bone mineral content, and percentage body fat.[5] Because DXA was collected multiple times per person, all estimations associated with DXA were carried out using a multiple-imputation methodology as recommended by the NHANES technical documentation (NCHS 2008).

In addition to these three PBF measures developed by body composition analyses based on BIA and DXA, we also estimate all of our models using clinically measured BMI. Typically, an indicator of obesity is employed as an additional outcome. The indicator is equal to one if a youth's BMI is at or above the 95th percentile based on age-gender specific CDC growth charts.

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[4] See http://www.cdc.gov/nchs/data/nhanes/dxa/dxa_techdoc.pdf for more information on DXA.

[5] Hologic software version 8:26:a3 is used to administer the scans (Centers for Disease Control and Prevention, 2008). More detail on the NHANES DXA examination and protocol features can be found at http://www.cdc.gov/nchs/data/nhanes/dxa/dxx_c.pdf.

We do not employ that indicator because PBF is a continuous measure and because obesity cutoffs based on PBF are not well developed.[6]

2.2 Food Prices

We make use of three separate measures of county- and year-specific food prices in our analysis. These are the real price of a calorie of food consumed at home, the real price of fast food consumed in restaurants, and the real price of fruits and vegetables consumed at home. These prices are obtained from the Council of Community and Economic Research (C2ER) Cost of Living Index, which has been published quarterly since 1968 for between 250 and 300 cities.[7] Researchers have made extensive use of the C2ER prices in studying obesity (Chou, Rashad, and Grossman 2008; Chou, Grossman, and Saffer, 2004; Powell 2009).

The C2ER collected prices of 44 different items during the period spanned by our NHANES data (1999-2004) and also reported the weight of each item in the typical budget of a household whose head holds a mid-management position. Included were the prices of 21 foods for consumption at home and three food items sold by fast-food restaurants for consumption on or off the premise.[8] We computed annual averages of each price and developed an algorithm to assign a relevant set of prices to each of the 3,114 counties in the U.S. In a number of cases, counties were assigned prices based on the prices of the geographically nearest within-state county for which price data were available. The measurement was based on the distance from the geographic center point of one county to another.

[6] See Section 5 for more details.
[7] C2ER was formerly referred to as ACCRA and before that as the American Chamber of Commerce Researchers Association.
[8] The 21 prices of foods for consumption at home pertain to the following items: a pound of t-bone steak, a pound of ground beef, a pound of Jimmy Dean or similar sausage, a pound of whole chicken, a 6 ounce can of tuna, a half gallon of milk, a dozen large eggs, a 1 pound tub of butter, 8 ounces of Kraft parmesan, a 10 pound sack of potatoes, a pound of bananas, a head of iceberg lettuce, a 24 ounce loaf of white bread, a 12 ounce can of coffee, a 4 pound bag of sugar, an 18 ounce box of Kellogg's corn flakes, a 16 ounce can of peas, a 30 ounce can of peaches, a 3 pound can of Crisco shortening, a 16 ounce can of corn, and a 2 liter bottle of Coke. The fast-food restaurant prices are specified in the text.

The 21 foods for consumption at home were used to construct a Laspeyres index of the price of one calorie for home consumption. The index is given by

$$L_{ijt} = \frac{\sum_{i=1}^{21} \pi_{ijt} C_{ib}}{\sum_{i=1}^{21} \pi_{ib} C_{ib}}. \tag{5}$$

In this equation π_{ijt} is the price of one calorie in food type i in county j in year t, C_{ib} is total calories consumed from food type i in the base year (b) in the U.S. as a whole, and π_{ib} is the price of one calorie in food type i in the base year in the U.S. as a whole. The index can be rewritten as

$$L_{ijt} = \frac{\sum_{i=1}^{21} \left(\frac{\pi_{ijt}}{\pi_{ib}}\right) \pi_{ib} C_{ib}}{\sum_{i=1}^{21} \pi_{ib} C_{ib}}. \tag{6}$$

Let p_{ib} be the price of one gram of food type i in the base year in the U.S. as a whole, let q_{ib} be the number of calories in one gram of that food type, and let X_{ib} be the number of grams consumed. Then

$$\pi_{ib} = (p_{ib}/q_{ib}) \tag{7}$$

$$C_{ib} = q_{ib}X_{ib} \tag{8}$$

$$\pi_{ib}C_{ib} = p_{ib}X_{ib} \tag{9}$$

$$\frac{\pi_{ib} C_{ib}}{\sum_{i=1}^{21} \pi_{ib} C_{ib}} = \frac{p_{ib} X_{ib}}{\sum_{i=1}^{21} p_{ib} X_{ib}} = k_{ib}. \tag{10}$$

Note that k_{ib} in the last equation is the fraction of total food outlays spent on food item i in the base year in the U.S. as a whole.

Given equation (10), equation (6) can be rewritten as

$$L_{ijt} = \sum_{i=1}^{21} k_{ib}\left(\frac{\pi_{ijt}}{\pi_{ib}}\right) = \sum_{i=1}^{21} k_{ib}\left(\frac{p_{ijt}}{p_{ib}}\right)\left(\frac{c_{ib}}{c_{it}}\right). \tag{11}$$

Equation (11) contains the reasonable assumption that the number of calories in one gram of each food item does not vary among counties in a given year. The U.S. Department of Agriculture National Nutrient Database reports the number of calories in one gram of each food item at a moment in time for the U.S. as a whole. There are, however, no data on variations in c_i over time. Therefore, we make a second reasonable assumption that c_i is the same in each year. This yields the final formula for L_{ijt}:

$$L_{ijt} = \sum_{i=1}^{21} k_{ib}\left(\frac{p_{ijt}}{p_{ib}}\right) = \sum_{i=1}^{21} k_{ib} r_{ijt}\,. \tag{12}$$

In equation (12), $r_{ijt} = p_{ijt}/p_{ib}$ is a simple price relative: the price of item i in county j in year t relative to the price of that item in the base period in the U.S. as a whole.

Equation (12) indicates that a Laspeyres index of the price of one calorie of food for home consumption coincides with a Laspeyres index of the price of food for home consumption. We compute it by first expressing each of the 21 food items as a simple price relative in which the denominator is the average nationwide nominal price of that item in 2000.[9] We take the fractions of food outlays spent on each food item (k_{ib}) for the same year. We then deflate by the annual Bureau of Labor Statistics Consumer Price Index (CPI, 2000 = 1) for all goods and services for the U.S. as a whole to convert the price of a calorie into real terms.

In addition to the price just described, we include a fast-food restaurant price in all of our models. That price is computed from the prices of a McDonald's Quarter-Pounder with cheese,

[9] Each simple price relative is adjusted so that the trend in it between 1988 and 2006 is the same as the trend in the comparable item in the Bureau of Labor Statistics Consumer Price Index.

a thin-crusted cheese pizza at Pizza Hut or Pizza Inn, and fried chicken at Kentucky Fried Chicken or Church's. In each case, the price is expressed as a simple relative as in equation (12). Then a weighted average of these simple price relatives is computed, where the weights are the shares of outlays on each item in total outlays on the three combined in 2000. The resulting price is deflated by the CPI. Finally, we compute a fruits and vegetables price from a subset of the prices of six items purchased for consumption at home: potatoes, bananas, lettuce, peas, peaches, and corn. The methodology is the same as that employed for the fast-food price. Simple price relatives are obtained and then averaged using as weights the shares of each item in total expenditures on fruits and vegetables in 2000. Finally the resulting price is deflated by the CPI.

3. Empirical Implementation

Our goal is to estimate the effects of various types of food prices on obesity outcomes of youths. To accomplish this goal, we specify a regression equation in the following form:

$$Y_{ijt} = \alpha_0 + \alpha_1 \text{Realcala}_{ijt} + \alpha_2 \text{Realfasta}_{ijt} + \alpha_3 Z_{ijt} + \mu_t + \nu_j + \varepsilon_{ijt}. \tag{13}$$

In equation (13), Y_{ijt} is one of the obesity outcomes for youth i in county j surveyed in year t. The key regressors are the price of a calorie in food for home consumption (Realcala_{ijt}) and the real price of food sold by fast-food restaurants (Realfast_{ijt}). The vector Z_{ijt} consists of youth and household specific characteristics. These include indicators for race and ethnicity, age in months, indicators for the ratio of income to the family's appropriate poverty threshold (less than 1.35, between 1.35 and 3, and between 3 and 5, and income missing) indicators for the living situation of the youth (in a married household, in a household headed with a female, in a household headed by a male, and living situation missing), household size, indicators for the education of the household head (less than 8^{th} grade, 9^{th} to 11^{th} grade, high school or GED, some college, college or higher).

The equation also includes controls for year fixed effects, μ_t, that account for unmeasured variables that vary over time at the national level and that are correlated with obesity outcomes and their determinants. Finally, the model includes county fixed effects, ν_j, to account for time-invariant, area-specific unmeasured factors that are correlated with the prices and weight. For example, locations with a high proportion of minority and low-income populations may have a higher concentration of fast-food restaurants, which may push the price of food in these restaurants down. These areas may also have a higher concentration of obese individuals if low socio-economic status and poverty are positively associated with obesity. Then a failure to control for county fixed effects may result in a biased estimate of the effect of food prices on youth obesity. Finally, the variable ε_{ijt} is an idiosyncratic error term.

We predict that α_1 and α_2 should be negative. We realize, however, that the effects of increases in the two food prices are not unambiguous. For example, an increase in the price of one calorie in food consumed at home can lead to an increase in food consumed at fast-food restaurants. To cite another example, dense convenience food prepared and consumed at home could rise in response to an increase in the price of fast food obtained in restaurants. Since we employ body composition rather than the consumption of certain foods as outcomes, our estimates will reflect these types of substitutions. In that sense, they are more informative than a regression in which the only outcome is fast-food restaurant consumption or consumption of convenience food. Note that the identification of the two price coefficients in the regression comes from within county changes in prices over time.

Obviously not all foods are unhealthy in the sense that increases in their consumption lead to weight gains. Therefore, we also estimate a version of equation (13), in which the calorie price is replaced by the fruits and vegetables price. We continue to include the fast-food restaurant price in that specification. Here our prediction is that the coefficient of the fruits and vegetables price should be positive. This reflects a substitution away from "healthy" foods and towards "unhealthy" foods in response to an increase in the price of the former. The specification takes into account an important point made by Auld and Powell (2009). They show that an increase in the price of foods with high energy densities (for example, food consumed at fast-food restaurants), with the price of foods with low energy densities (for example, fruits and vegetables) held constant, will result in a reduction in total calories consumed if the price of dense food is cheaper per calorie consumed than the price of non-dense food also per calorie consumed.

The Auld-Powell model highlights another reason why the signs of the price coefficients in equation (13) are ambiguous. In that equation, the price of one calorie in food for home consumption is held constant when the price of food consumed in fast-food restaurants varies. But that does not guarantee that the prices of dense and non-dense food consumed at home are fixed. Moreover, when the price of one calorie consumed at home varies, with the price of fast-food restaurant food held constant, prices of dense and non-dense food may be changing. In theory, one would want to include the calorie prices of many different types of food in the regression. Multicollinearity among these prices makes that approach infeasible, however. Intercorrelations among the three prices that we do use are fairly high. That is another reason for the two alternative specifications and also is a reason for interpreting the results with caution.

We estimate all of our models separately for males and females because Chumlea's measure is based on formulae developed separately for each gender. Furthermore, body structures may exhibit major differences in size, shape, composition, and function during puberty. For example, most girls begin puberty between the ages of 9 and 13, while most boys experience puberty later between the ages of 10 and 16. This suggests that a pooled specification may not fully account for gender-specific differences in body growth. Finally, in light of the well-known health disparities between whites and minorities, we also present estimates from regressions that are estimated separately for each gender and race/ethnicity combination. Doing so will allow us to assess the extent to which the food price-body composition relationship differs by race/ethnicity.

4. Results

Definitions, means and standard deviations of all variables used in the empirical analysis are reported in Table 1. Means and standard deviations are weighted using the medical examination sampling weights provided in the NHANES. Note that nonwhites are oversampled, which is why there are more observations for nonwhites in the columns that contain data by gender and race/ethnicity.

All three measures of PBF indicate that females possess a higher percentage of their total weight as body fat than males. Furthermore, the means for PBF from the measures developed from the prediction equations of Boileau and Chumlea are very close to each other. Those generated by DXA are somewhat larger, but not far from the other two measures of PBF. Note that both Boileau and Chumlea are based on BIA, while DXA is based on x-ray imaging. While the PBF figures are indicative of a higher percentage body fat among females than males, this pattern does not hold if we focus on BMI. Nonwhites have higher values of PBF than whites for both males and females. One exception to this pattern is DXA measured for males, where the difference is quite small. While the differences are perhaps not as striking as one might expect, nonwhite adolescents and especially adults exhibit higher rates of obesity than whites (Flegal et al. 2012; Ogden et al. 2012). Since obese youths are likely to become obese adults, this motivates, in part, the gender and race/ethnicity-specific regression analysis later in this section.

Additionally, the means for control variables follow patterns that are usually consistent with one's expectations. For example, heads in households where minority children live are more likely to have lower education, more likely to be in poverty, less likely to be married, and more likely to be female.

Table 2 presents pairwise correlations among our outcome measures. Consistent with the means presented in Table 1, there is a high degree of correlation among our three body composition measures. For example, the four gender and race/ethnicity-specific correlation coefficients between PBF from Boileau and Chumlea are at least as large as 0.98. The eight pairwise correlations between the PBF measure from DXA and the PBF measures from Boileau and Chumlea are somewhat lower, ranging from 0.83 to 0.92. The twelve pairwise correlation coefficients between BMI and each measure of PBF are smaller, falling into an interval from 0.67 through 0.88. Similar conclusions hold within each of the four gender-majority/minority-specific groups.

We present our regression results in Tables 3-5. In Table 3, Columns 1 and 2 show the results for all males and columns 3 and 4 show them for all females. Columns 1 and 3 present estimates from the model specifications that include the real price of one calorie of food for home consumption and the real price of food sold by fast-food restaurants. Columns 2 and 4 show estimates from the model specifications that replace the real price of one calorie of food for home consumption with the real price of fruits and vegetables. Finally, Table 2 contains four panels, each presenting the price estimates for one of the outcomes: PBF based on Boileau, PBF based on Chumlea, PBF based on DXA, and BMI.[10] Note that all the regressions are weighted using the appropriate sampling weights and the standard errors are clustered by county.[11] This allows the error term to be correlated among different youths in the same county both at a moment in time and over time.

Focusing on the first panel (PBF-Boileau), we find that the price of one calorie of food consumed at home is associated with a decrease in PBF for both genders, but the male estimate is smaller in magnitude than the female estimate and is not statistically significant. Evaluated at sample means, the male elasticity of -0.87 is slightly larger in absolute value than the female elasticity of -0.76. That is a 10 percent increase in the price of one calorie in food for home consumption lowers PBF for males by approximately 9 percent and lowers PBF for females by approximately 8 percent.

The fast-food restaurant price coefficient in the first panel is sensitive to the specific model that is being estimated. When the price of a calorie is held constant, the coefficient is negative and significant for both genders (see columns one and three). On the other hand, the coefficient is smaller and not significant but still negative when the price of fruits and vegetables is held constant (see columns two and four). These results are due in part to multicollinearity, and we interpret them with caution. Here and in the remainder of the paper, we summarize the magnitude of the fast-food price effect based on an average elasticity implied by the two

[10] In the interest of space, we only provide a discussion of the price coefficients in the paper. However, the estimates on other covariates are usually consistent with those documented in the relevant literature. The full results for the models with each of the outcomes are available from the authors upon request.
[11] Note that the sample sizes in the four panels of Table 3 vary with the outcome variables, depending on the availability of these outcome measures in various NHANES rounds. In order to assess whether the differences in the price effects in the table are due to the differences in the sample sizes, we investigated the results for all these models limiting the analyses to the same sample size. These results are very similar to those presented here and are available from the authors upon request.

specifications. For males the figure is -1.73 and for females, it is -0.76. When PBF is measured by the Boileau formula, the coefficient on the price of fruits and vegetables is positive for both genders (see columns 2 and 4 of panel 1) but significant only for females. The female elasticity of 0.88 is larger than the male elasticity of 0.72.

Taken as a whole, the results in the first panel of Table 3 suggest that the price of healthy food, measured by the price of fruits and vegetables, is a more important determinant of female PBF than of male PBF. When this price rises by 10 percent, PBF rises by 9 percent for females but by 7 percent for males, and the estimate is significant only for females. On the other hand, the price of a calorie in food consumed at home or in fast-food restaurants plays a more important role, as reflected by elasticities, in male than in female body fatness. For example, a 10 percent reduction in the price of fast-food restaurant food is associated with a 17 percent increase in PBF for males, which is much larger than the 8 percent reduction for females. This result and the corresponding result for the food at home price may reflect a greater willingness of girls to substitute towards healthy food when the price of unhealthy food rises.

In general, the results for the other three outcomes (PBF-Chumlea, PBF-DXA, and BMI) tell a similar story. Male calorie price elasticities range from -0.68 in the case of PBF-DXA to -1.00 in the case of BMI. The corresponding range for female elasticities is -0.52 for BMI to -0.84 for PBF-Boileau. Male fast-food price elasticities range from -0.76 (PBF DXA) to -1.73 (PBF-Boileau). Female fast-food price elasticities are all negative only for the Boileau-PBF and Chumlea-PBF (-0.76 for the former and -0.50 for the latter). The only negative and significant effect of this variable for females is obtained in the model that includes the price of a calorie in food consumed at home. The price of fruits and vegetables never has a significant effect on male

PBF. For females, the effect always is positive and significant, and the elasticities lie in a fairly tight range from 0.79 for BMI to 0.88 for PBF-Boileau.

In Tables 4 and 5, we investigate whether the effects of food prices on PBF differ between adolescents from different racial and ethnic backgrounds. To do this, we estimate our gender-specific models for whites in the former table and for nonwhites (blacks and Hispanics) in the latter table. Before these results are discussed, it is useful to point to some factors that might generate differences.

An overarching consideration is that nonwhite youths come from families with lower incomes than white youths. To be sure, such proxies for family income, and the education of the head of the household, and indicators for income to poverty ratios are included as regressors. But these variables are held constant at lower levels for nonwhites than for whites in the estimation of food price effects. Hence, in a fundamental sense, differences in price effects can be traced to interactions between family income and price.[12]

Components that contribute to the "full" price of consuming food in addition to the money price may also contribute to differences in observed money price elasticities between youths from low-income families and other youths. One of these components is the time price. Shopping time (the sum of the time spent traveling to and from a food outlet and waiting in line to pay for the purchase) is required to obtain food for consumption at home. Travel and waiting time also are required to consume food at restaurants. The time price of food consumption, defined as time required to purchase and consume it multiplied by the value of time, is negatively related to the per capita availability of food outlets. A one percent change in the

[12] We do not estimate separate regressions by gender, race, and income because the resulting estimates would be based on a small number of observations in each group.

money price of food amounts to a smaller percentage change in the full price and hence to a smaller percentage change in consumption the larger the time price component.

Another component of the full price of food consumption is the monetary value of the future health consequences of that consumption. This component is positive in the case of food served in fast-food restaurants and acts as a tax on its consumption. On the other hand, the component is negative in the case of fruits and vegetables and acts as a subsidy to its consumption. A one percent change in money price results in a larger percentage in full price the smaller are the future health costs or the larger are the future health benefits.

The time price of consuming food is likely to be smaller for low-income families who have lower wage rates and hence a lower value of time. This effect is reinforced if fast-food restaurants are more likely to locate in poor areas, but weakened if supermarkets and grocery stores are less likely to locate in these areas. While previous research has documented these locational patterns, Lee (2012) finds that both fast-food restaurants and large-scale grocery stores are more prevalent in poor neighborhoods.[13] These considerations suggest that minorities may be more sensitive to fast-food restaurant prices while leaving differences in the sensitivity of responses to fruits and vegetables prices an open issue.

Future costs should be less important to parents and youths in poorer, less educated families, and future benefits should be more important to parents and youths in richer, more-educated families. Future costs and benefits are smaller for poor parents who have low wage rates and plausibly expect their children to have low wage rates as adults. Another factor is that the poor and the less educated are likely to have lower time discount factors (higher rates of time preference for the present) than the rich and more educated (Becker and Mulligan 1997).

[13] Lee's data are at the census-track level, but it is not clear how many tracts are represented.

Variations in the evaluation of future costs and benefits imply larger fast-food price elasticities for the poor but smaller fruits and vegetables price elasticities.

Turning to the results in Tables 4 and 5, one sees that that the eight fruits and vegetables price coefficients are positive for whites and six of the eight coefficients are significant. In Table 4, for white females, the elasticity is approximately 2 for BMI, and PBF-Boileau and PBF-Chumlea, while it is lower at 0.42 for PBF-DXA. The range for white males is from 3.58 for PBF-Boileau to 0.74 for BMI. The pattern is very different for nonwhites in Table 5. Only one coefficient is positive and significant (female PBF-DXA), and all four male coefficients are negative. These results are consistent with an explanation that stresses higher rates of time preference for the present and lower expected future wage rates among nonwhite youths and their parents than among white youths.

The fast-food price results tell a somewhat similar story, although here the evidence is less conclusive. Focusing on a comparison of white and nonwhite males, one sees that the eight fast-food restaurant coefficients are negative and significant for the latter group, while only four of the eight coefficients are negative and significant for the former group. Moreover, two of the coefficients are positive (for BMI and PBF-DXA) for white males. The average elasticity range for non-whites (between -1.12 for BMI and -1.74 for PBF-Boileau) is much tighter than for whites (between -0.28 in the case of PBF-DXA and -2.20 in the case of PBF-Boileau). These findings are consistent with the lower future costs and benefits hypothesis, but they also are consistent with an explanation that stresses lower time costs for nonwhite youths or their parents.

For both white and nonwhite females, the price of fast food plays a much less important role in body composition outcomes than for white and nonwhite males. Only one of the nonwhite female coefficients is negative. Two of the eight coefficients are positive for white females, and only three of the six positive coefficients are significant. These results mirror the gender difference in the fast-food restaurant price obtained from non-race specific regressions. It is puzzling, however, that the price elasticity for nonwhite females never is negative.

Finally, male body composition outcomes continue to be more responsive to changes in the price of a calorie in food consumed at home than female outcomes when separate estimates are obtained for whites and nonwhites. While the elasticities are bigger for white males than for nonwhite males, the coefficients on which they are based are estimated more precisely for the latter group. For example, the PBF-DXA coefficient is not significant for white males. A similar conclusion emerges when the estimates for white and nonwhite females are compared. Except in the case of PBF-Boileau, the white female elasticity is bigger than the nonwhite female elasticity. But the calorie price coefficient never is significant for the former group, while three of the four coefficients are significant for the latter group.

5. Conclusions

The proportion of children who are obese has reached epidemic levels in the last three decades. The rising prevalence of childhood obesity is a source of concern among public health officials because of the well-documented health problems tied to obesity for both children and adults. There are a large number of policy efforts under way to stop or reverse this trend. For example, the Child Nutrition and Women Infants and Children (WIC) Reauthorization Act of 2004 required that all local education agencies participating in the National School Lunch Program create local wellness policies no later than July 2006. The Kids Walk-to-School Program developed by the Centers for Disease Control and Prevention (CDC) aims to increase opportunities for daily physical activity by encouraging children to walk to and from school in groups accompanied by adults. An increasing number of schools are limiting access to foods high in fats and sugars by banning soda machines and snack bars in cafeterias and school stores. The School Breakfast and the National School Lunch Programs are two federal entitlement programs that provide nutritionally balanced, low-cost or free breakfasts and lunches to millions of children each school day.

Regulating the prices of healthy and unhealthy foods is a promising option for influencing obesity outcomes among youths. Previous studies of the analysis of the effect of prices on obesity outcomes of youths have exclusively used BMI or BMI-based indicators of obesity. However, obesity is defined as excess of body fat, and it is body fat (and not fat-free mass) that is responsible for the detrimental health effects of obesity. Therefore, an increasing number of studies point to the limitation of BMI in distinguishing fat from fat-free mass. They caution against a sole reliance on BMI and point to a need for developing alternative measures of obesity.

In this study, we consider alternative measures of obesity based on body composition rather than a BMI-based measure to investigate the effects of food prices on obesity among children and to assess their performance relative to BMI. In particular, we estimate the effects of the prices of various types of food on percentage body fat outcomes derived from BIA and DXA using data from various waves of NHANES. We also examine whether the effects of food prices on these outcomes differ between nonwhite and white adolescents and motivate this analysis by the lower socioeconomic background that characterizes nonwhites.

Our findings suggest that increases in the real price of one calorie in food for home consumption and the real price of fast-food restaurant food lead to improvements in body composition outcomes among youths. We also find that an increase in the price of fruits and vegetables has negative consequences for these outcomes. The former effects are more important for males compared to females, for nonwhite males compared to white males and in the case of the calorie price for nonwhite females compared to white females. The "healthy food" price effect, reflected by the price of fruits and vegetables, is more important for whites compared to nonwhites.

There are two important implications of our study. One pertains to future research with such measures of body composition as the PBF and the other pertains to public policy. With regard to the first issue, many of our estimates suggest that the PBF is at least as sensitive to prices as BMI and in some instances more sensitive to prices than BMI. Given that, it would be useful to employ an obesity indicator defined by PBF as an additional outcome. We have placed this issue on an agenda for future research because obesity cutoffs based on PBF result in an implausibly large percentage of youths being classified as obese. For example, Boreham, Twisk, and Savage (1997) classify adolescent boys with PBF greater than 20 percent as obese and adolescent girls with PBF greater than 24 percent as obese. Grossman, Tekin, and Wada (2012) show that if these cutoffs are applied to their NLSY97 data, approximately 50 percent of males and 90 percent of females are classified as obese. Even if the cutoffs for adults recommended by the National Institute of Diabetes and Digestive and Kidney Diseases (2006) for adults of greater than 25 percent for males and greater than 30 percent for females are used, approximately 22 percent of males and 60 percent of females are identified as obese. The development of more reasonable cutoffs deserves high priority on an agenda for future research. Part of that undertaking should involve an examination of the characteristics of individuals classified as obese by one measure but not the other and vice versa. It also is valuable to obtain PBF cutoffs that result in the same percentage of obese youths as BMI cutoffs.

With regard to the policy implications of our research, “fat taxes” or taxes on foods with high caloric content have received a considerable amount of recent attention in the so-called “war on obesity.” One specific version is a tax on sugar-sweetened beverages (soda). Research summarized by Grossman and Mocan (2011) finds that soda taxes have very modest effects on calories consumed from soda on BMI. Indeed, Fletcher, Frisvold, and Tefft (2010) find the modest decline in soda-based calories is completely offset by increases in the consumption of other high-calorie drinks such as juice and milk. Our results do not directly speak to the potential impacts of a soda tax, but they do suggest that a tax on meals purchased in fast-food restaurants or a subsidy to the consumption of fruits and vegetables would lead to better obesity outcomes among adolescents. These findings take account of any adverse effects due to substitution towards or away from other food items in response to taxes and subsidies.

Of course, we also find that an increase in the price of a calorie regardless of its source would improve obesity outcomes. Clearly, a food tax could be imposed to increase the price of a calorie. If it took the form of a specific excise tax (fixed amount per calorie in a gram of each food type), it would have the desirable effect of raising the relative prices of foods that are cheap sources of calories. An ad valorem tax (fixed percentage of price) would not have that effect because it would not alter the relative price of dense food. But the latter tax would be much easier to impose and administer. Taken at face value, our results suggest that such a tax might be an effective tool in the war on obesity.

A good deal of caution is required here. Taxes are blunt instruments that impose significant welfare costs on individuals who consume food in moderation. Moreover, in the case of adolescents, an additional issue is that parents may more easily and immediately affect the choices made by their children than the government. Indeed, some of our results point to higher rates of time preference and lower expected future wage rates among nonwhite parents and youths as explanations of why minorities are more sensitive to fast-food prices and less sensitive to fruits and vegetables prices than whites. These interpretations add to the wide range of benefits to early childhood intervention programs emphasized by Heckman and colleagues (for example, Conti and Heckman 2012). These programs aim to improve the cognitive and non-cognitive skills of minorities and perhaps to give them more of a future orientation as well. Hence, we view our contribution as an input into the policy debate concerning the most effective ways to reverse the upward trend in obesity. We have shown that selective taxes or subsidies may be able to accomplish part of this goal. We also have shown that uniform increases or decreases in the price of food do have the expected impacts on body weight. An integrated approach to nutrition could be more focused and effective than the current approach used by the federal and states governments in taxing and subsidizing nutritional goods and allocating food stamps, which currently involves little or no consideration of the impact of these decisions on youth obesity. We leave it to others to evaluate the external costs and benefits of policies to combat obesity.

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Table 1
Descriptive Statistics

Variable	All		Whites		Nonwhites	
	Males	**Females**	**Males**	**Females**	**Males**	**Females**
PBF-Boileau[1]	19.847	29.401	19.306	28.371	20.916	31.355
	(9.359)	(8.575)	(9.262)	(8.574)	(9.467)	(8.238)
PBF-Chumlea[1]	20.322	31.024	19.815	30.055	21.329	32.862
	(7.777)	(8.410)	(7.686)	(8.385)	(7.865)	(8.154)
PBF-DXA[1]	24.158	33.742	24.344	33.522	23.793	34.188
	(7.872)	(7.082)	(7.642)	(6.976)	(8.298)	(7.279)
BMI	23.171	23.465	23.021	22.935	23.457	24.465
	(5.474)	(5.692)	(5.259)	(5.260)	(5.866)	(6.307)
Price of a calorie of food	1.056	1.057	1.041	1.044	1.085	1.082
for home consumption	(0.136)	(0.139)	(0.135)	(0.142)	(0.134)	(0.129)
Price of fast-food	1.010	1.013	1.008	1.011	1.015	1.017
	(0.074)	(0.074)	(0.0717)	(0.0723)	(0.0779)	(0.0775)
Price of fruits and vegetables	1.029	1.033	1.013	1.017	1.061	1.062
	(0.153)	(0.158)	(0.156)	(0.164)	(0.143)	(0.142)
White	0.663	0.652				
	(0.473)	(0.477)				
Black	0.153	0.158			0.454	0.454
	(0.360)	(0.365)			(0.498)	(0.498)
Hispanic	0.184	0.190			0.546	0.546
	(0.388)	(0.393)			(0.498)	(0.498)
Age	15.836	15.829	15.87	15.86	15.77	15.76
	(2.302)	(2.268)	(2.309)	(2.256)	(2.287)	(2.293)
$0.00 \leq$ PIR* < 1.85	0.433	0.450	0.333	0.356	0.630	0.627
	(0.480)	(0.480)	(0.460)	(0.462)	(0.459)	(0.463)
$1.85 \leq$ PIR ≤ 3.00	0.180	0.184	0.180	0.192	0.179	0.170
	(0.371)	(0.374)	(0.375)	(0.381)	(0.365)	(0.359)
$3.00 <$ PIR ≤ 5.00	0.387	0.365	0.487	0.452	0.191	0.203
	(0.476)	(0.469)	(0.492)	(0.486)	(0.372)	(0.384)
PIR is missing	0.062	0.070	0.0463	0.0647	0.0927	0.0797
	(0.241)	(0.255)	(0.210)	(0.246)	(0.290)	(0.271)
Household size	4.328	4.306	4.167	4.165	4.646	4.570
	(1.387)	(1.354)	(1.295)	(1.264)	(1.503)	(1.472)
Household head married	0.636	0.609	0.695	0.663	0.520	0.507
	(0.456)	(0.461)	(0.436)	(0.445)	(0.471)	(0.473)
Household head female	0.418	0.491	0.368	0.452	0.516	0.564
	(0.493)	(0.500)	(0.482)	(0.498)	(0.500)	(0.496)
Household head male	0.582	0.509	0.632	0.548	0.484	0.436
	(0.493)	(0.500)	(0.482)	(0.498)	(0.500)	(0.496)
Household head	0.099	0.107	0.0916	0.106	0.115	0.109

living situation missing	(0.299)	(0.309)	(0.289)	(0.308)	(0.319)	(0.312)
HH education<8th grade	0.083	0.086	0.0394	0.0405	0.170	0.171
	(0.267)	(0.272)	(0.184)	(0.185)	(0.365)	(0.369)
HH education: 9-11 grade	0.150	0.151	0.0966	0.101	0.253	0.246
	(0.348)	(0.351)	(0.288)	(0.293)	(0.424)	(0.424)
HH education:	0.270	0.280	0.276	0.295	0.257	0.251
high school/GED	(0.435)	(0.442)	(0.441)	(0.449)	(0.424)	(0.425)
HH education: some	0.300	0.301	0.333	0.325	0.234	0.255
college	(0.451)	(0.452)	(0.467)	(0.463)	(0.411)	(0.429)
HH Education: college	0.198	0.182	0.255	0.239	0.0857	0.0769
or higher	(0.393)	(0.381)	(0.433)	(0.422)	(0.268)	(0.259)
Year: 1999	0.111	0.110	0.0974	0.0986	0.138	0.130
	(0.314)	(0.312)	(0.297)	(0.298)	(0.345)	(0.337)
Year: 2000	0.185	0.182	0.184	0.166	0.187	0.213
	(0.388)	(0.386)	(0.388)	(0.372)	(0.390)	(0.410)
Year: 2001	0.193	0.194	0.208	0.212	0.165	0.160
	(0.395)	(0.395)	(0.406)	(0.409)	(0.371)	(0.366)
Year: 2002	0.163	0.169	0.157	0.164	0.173	0.177
	(0.369)	(0.374)	(0.364)	(0.371)	(0.378)	(0.381)
Year: 2003	0.168	0.173	0.158	0.167	0.189	0.184
	(0.374)	(0.378)	(0.365)	(0.373)	(0.391)	(0.387)
Year: 2004	0.180	0.173	0.196	0.192	0.148	0.136
	(0.384)	(0.378)	(0.397)	(0.394)	(0.355)	(0.343)
Observations[1]	3,348	3,084	898	827	2,450	2,257

Notes: Standard deviations are in parentheses. Means and standard deviations are weighted using the NHANES sampling weights. All the prices are in real terms. * PIR stands for poverty-income ratios. [1] Because observation sizes differ slightly by the outcome, the mean and the observations sizes are reported for BMI, which had the maximum observation size. Observations sizes for PBF-Boileau, PBF-Chumlea, PBF-DXA are slightly less. They are available from authors upon request.

Table 2
Pairwise Correlations among Body Fat and BMI Measures

All Males				
	PBF-Boileau	PBF-Chumlea	BMI	PBF-DXA
PBF-Boileau	1.000			
PBF-Chumlea	0.987	1.000		
BMI	0.797	0.787	1.000	
PBF-DXA	0.827	0.857	0.681	1.000

All Females				
	PBF-Boileau	PBF-Chumlea	BMI	PBF-DXA
PBF-Boileau	1.000			
PBF-Chumlea	0.983	1.000		
BMI	0.826	0.890	1.000	
PBF-DXA	0.887	0.910	0.847	1.000

White Males				
	PBF-Boileau	PBF-Chumlea	BMI	PBF-DXA
PBF-Boileau	1.000			
PBF-Chumlea	0.986	1.000		
BMI	0.802	0.786	1.000	
PBF-DXA	0.829	0.863	0.671	1.000

White Females				
	PBF-Boileau	PBF-Chumlea	BMI	PBF-DXA
PBF-Boileau	1.000			
PBF-Chumlea	0.984	1.000		
BMI	0.835	0.897	1.000	
PBF-DXA	0.898	0.921	0.861	1.000

Nonwhite males				
	PBF-Boileau	PBF-Chumlea	BMI	PBF-DXA
PBF-Boileau	1.000			
PBF-Chumlea	0.988	1.000		
BMI	0.790	0.792	1.000	
PBF-DXA	0.841	0.867	0.702	1.000

Nonwhite females				
	PBF-Boileau	PBF-Chumlea	BMI	PBF-DXA
PBF-Boileau	1.000			
PBF-Chumlea	0.981	1.000		
BMI	0.811	0.882	1.000	
PBF-DXA	0.881	0.902	0.828	1.000

Notes: Sample sizes for pairwise correlations range between 2865 and 3023 for males, 1949 and 2759 for females, 772 to 898 for white males, 577 and 827 for white females, 2098 and 2453 for nonwhite males, and 1372 and 2258 for nonwhite males. Correlations are weighted using the NHANES sampling weights.

Table 3
Estimates of the Effect of Food Prices on Body Composition by Gender

	All Males		All Females	
PBF - Boileau				
Price of a calorie	-16.436 (-1.397)		-21.063* (-1.770)	
Elasticity	[-0.874]		[-0.755]	
Price of fast food	-38.742** (-2.596)	-29.383 (-1.614)	-29.441* (-1.824)	-14.818 (-0.985)
Elasticity	[-1.970]	[-1.490]	[-1.010]	[-0.509]
Price of fruits and vegetables		13.839 (0.883)		25.079* (1.960)
Elasticity		[0.717]		[0.878]
R-squared	0.114	0.113	0.213	0.213
Observations	2,993	2,993	2,759	2,759
PBF - Chumlea				
Price of a calorie	-16.515 (-1.645)		-24.636* (-1.950)	
Elasticity	[-0.857]		[-0.837]	
Price of fast food	-28.825*** (-2.717)	-20.886 (-1.567)	-23.396 (-1.387)	-7.575 (-0.490)
Elasticity	[-1.430]	[-1.030]	[-0.762]	[-0.246]
Price of fruits and vegetables		9.687 (0.754)		25.720* (1.946)
Elasticity		[0.490]		[0.854]
R-squared	0.110	0.110	0.1989	0.198
Observations	3,002	3,002	2,759	2,759
PBF - DXA				
Price of a calorie	-15.650* (-1.909)		-22.520** (-2.031)	
Elasticity	[-0.684]		[-0.707]	
Price of fast food	-21.770** (-2.411)	-14.600 (-1.304)	16.182 (0.918)	34.553* (1.884)
Elasticity	[-0.910]	[-0.610]	[0.484]	[1.032]
Price of fruits and vegetables		7.403 (0.827)		26.623** (2.254)
Elasticity		[0.315]		[0.819]
R-squared	0.125	0.125	0.150	0.160
Observations	3,368	3,368	2,257	2,257
BMI				
Price of a calorie	-22.020*** (-4.843)		-11.618 (-1.389)	
Elasticity	[-1.000]		[-0.523]	
Price of fast food	-22.109*** (-4.973)	-17.378** (-2.329)	-4.744 (-0.553)	4.585 (0.674)
Elasticity	[-0.963]	[-0.757]	[-0.204]	[0.198]
Price of fruits and vegetables		-4.978 (0.793)		17.862** (2.609)
Elasticity		[-0.221]		[0.7861]
R-squared	0.138	0.137	0.152	0.153
Observations	3,351	3,351	3,085	3,085

Notes: All regressions are weighted using the NHANES sampling weights. t-statistics reported in parentheses are based on standard errors that are clustered at the county level. All the prices are in real terms. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. Elasticities reported in brackets are computed at sample means.

Table 4
Estimates of the Effect of Food Prices on Body Composition, Whites by Gender

	White Males		White Females	
PBF – Boileau				
Price of a calorie	-58.178** (-2.322)		-21.036 (-0.370)	
Elasticity	[-3.130]		[-0.771]	
Price of fast food	-77.68*** (-3.947)	-6.773 (-0.199)	-53.001** (-2.160)	-5.299 (-0.141)
Elasticity	[-4.050]	[-0.353]	[-1.880]	[-0.188]
Price of fruits and vegetables		68.24** (2.601)		58.551* (1.683)
Elasticity		[3.576]		[2.090]
R-squared	0.171	0.173	0.261	0.263
Observations	807	807	754	754
PBF - Chumlea				
Price of a calorie	-58.494*** (-3.584)		-35.780 (-0.631)	
Elasticity	[-3.070]		[-1.230]	
Price of fast food	-64.407*** (-3.872)	-3.167 (-0.119)	-50.057* (-1.967)	6.183 (0.168)
Elasticity	[-3.270]	[-0.160]	[-1.670]	[0.207]
Price of fruits and vegetables		49.720** (2.232)		59.697* (1.773)
Elasticity		[2.537]		[2.012]
R-squared	0.162	0.163	0.246	0.247
Observations	811	811	754	754
PBF – DXA				
Price of a calorie	-37.198 (-1.622)		-41.502 (-0.678)	
Elasticity	[-1.590]		[-1.290]	
Price of fast food	-31.046* (-1.770)	17.33 (0.928)	-81.770** (-2.613)	-41.472 (-0.565)
Elasticity	[-1.280]	[0.717]	[-2.450]	[-1.240]
Price of fruits and vegetables		48.29** (2.489)		13.896 (0.374)
Elasticity		[2.001]		[0.422]
R-squared	0.168	0.170	0.221	0.221
Observations	902	902	652	652
BMI				
Price of a calorie	-31.888*** (-3.548)		-31.492 (-0.893)	
Elasticity	[-1.440]		[-1.430]	
Price of fast food	-21.792** (-2.255)	6.224 (0.506)	-18.869 (-1.113)	27.027 (1.570)
Elasticity	[-0.953]	[0.272]	[-0.831]	[1.190]
Price of fruits and vegetables		16.892 (1.382)		45.164*** (3.246)
Elasticity		[0.744]		[2.002]
R-squared	0.221	0.221	0.206	0.209
Observations	898	898	827	827

Notes: All regressions are weighted using the NHANES sampling weights. t-statistics reported in parentheses are based on standard errors that are clustered at the county level. All the prices are in real terms. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. Elasticities reported in brackets are computed at sample means.

Table 5
Estimates of the Effect of Food Prices on Body Composition, Nonwhites by Gender

	Nonwhite Males		Nonwhite Females	
PBF – Boileau				
Price of a calorie	-21.106*** (-3.669)		-33.477*** (-4.087)	
Elasticity	[-1.090]		[-1.154]	
Price of fast food	-35.546*** (-4.747)	-36.131*** (-3.436)	0.981 (0.088)	9.622 (0.885)
Elasticity	[-1.720]	[-1.750]	[0.032]	[0.312]
Price of fruits and vegetables		-15.579* (-1.913)		12.151 (1.387)
Elasticity		[-0.790]		[0.412]
R-squared	0.114	0.114	0.154	0.151
Observations	2,186	2,186	2,005	2,005
PBF – Chumlea				
Price of a calorie	-17.675** (-2.965)		-37.907*** (-4.881)	
Elasticity	[-0.899]		[-1.240]	
Price of fast food	-27.530*** (-5.192)	-27.668*** (-3.451)	3.193 (0.260)	12.306 (1.083)
Elasticity	[-1.310]	[-1.310]	[0.099]	[0.380]
Price of fruits and vegetables		-11.761* (-1.824)		11.297 (1.165)
Elasticity		[-0.585]		[0.365]
R-squared	0.113	0.112	0.145	0.141
Observations	2,191	2,191	2,005	2,005
PBF – DXA				
Price of a calorie	-17.82*** (-3.942)		-11.04 (-1.263)	
Elasticity	[-0.812]		[-0.351]	
Price of fast food	-30.14*** (-5.326)	-29.81*** (-3.940)	29.21* (1.797)	
Elasticity	[-1.280]	[-1.270]	[0.866]	
Price of fruits and vegetables		-11.29 (-1.564)		27.79*** (3.076)
Elasticity		[-0.503]		[0.875]
R-squared	0.153	0.152	0.166	0.167
Observations	2,466	2,466	1,605	1,605
BMI				
Price of a calorie	-22.381*** (-5.479)		-14.206** (-2.382)	
Elasticity	[-1.030]		[-0.628]	
Price of fast food	-26.633*** (-6.328)	-25.542*** (-3.262)	-0.626 (-0.0853)	3.997 (0.703)
Elasticity	[-1.150]	[-1.100]	[-0.026]	[0.166]
Price of fruits and vegetables		-11.754* (-1.790)		9.968 (1.492)
Elasticity		[-0.531]		[-0.433]
R-squared	0.109	0.107	0.122	0.122
Observations	2,453	2,453	2,258	2,258

Notes: All regressions are weighted using the NHANES sampling weights. t-statistics reported in parentheses are based on standard errors that are clustered at the county level. All the prices are in real terms. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. Elasticities reported in brackets are computed at sample means.