Sociodemographic Differences in Fast Food Price Sensitivity

Katie A. Meyer, ScD; David K. Guilkey, PhD; Shu Wen Ng, PhD; Kiyah J. Duffey, PhD; Barry M. Popkin, PhD; Catarina I. Kiefe, MD; Lyn M. Steffen, PhD; James M. Shikany, DrPH; Penny Gordon-Larsen, PhD

**IMPORTANCE** Fiscal food policies (eg, taxation) are increasingly proposed to improve population-level health, but their impact on health disparities is unknown.

**OBJECTIVE** To estimate subgroup-specific effects of fast food price changes on fast food consumption and cardiometabolic outcomes.

**DESIGN, SETTING, AND PARTICIPANTS** Twenty-year follow-up (5 examinations) in a biracial US prospective cohort: Coronary Artery Risk Development in Young Adults (CARDIA) (1985/1986-2005/2006, baseline N = 5115). Participants were aged 18 to 30 years at baseline; design indicated equal recruitment by race (black vs white), educational attainment, age, and sex. Community-level price data from the Council for Community and Economic Research were temporally and geographically linked to study participants' home address at each examination.

**MAIN OUTCOMES AND MEASURES** Participant-reported number of fast food eating occasions per week, body mass index (BMI), and homeostasis model assessment insulin resistance (HOMA-IR) from fasting glucose and insulin concentrations. Covariates included individual-level and community-level social and demographic factors.

**RESULTS** In repeated measures regression analysis, multivariable-adjusted associations between fast food price and consumption were nonlinear (quadratic, \( P < .001 \)), with significant inverse estimated effects on consumption at higher prices; estimates varied according to race (interaction \( P = .04 \)), income (\( P = .07 \)), and education (\( P = .03 \)). At the 10th percentile of price ($1.25/serving), blacks and whites had mean fast food consumption frequency of 2.20 (95% CI, 2.07-2.33) and 1.55 (1.45-1.65) times/wk, respectively, whereas at the 90th percentile of price ($1.53/serving), respective mean consumption estimates were 1.86 (1.75-1.97) and 1.50 (1.41-1.59) times/wk. We observed differential price effects on HOMA-IR (inverse for lower educational status only [interaction \( P = .005 \)] and at middle income only [interaction \( P = .021 \)]) and BMI (inverse for blacks, less education, and middle income; positive for whites, more education, and high income [all interaction \( P < .001 \)]).

**CONCLUSIONS AND RELEVANCE** We found greater fast food price sensitivity on fast food consumption and insulin resistance among sociodemographic groups that have a disproportionate burden of chronic disease. Our findings have implications for fiscal policy, particularly with respect to possible effects of fast food taxes among populations with diet-related health disparities.

**Author Affiliations:** Author affiliations are listed at the end of this article.

**Corresponding Author:** Penny Gordon-Larsen, PhD, Department of Nutrition, Gillings School of Global Public Health, University of North Carolina, 123 W Franklin St (University Square), Campus Box 8120, Chapel Hill, NC 27516 (pglarsen@email.unc.edu).

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Fiscal policies, such as targeted taxes or subsidies to motivate healthy diet choices, have received increased consideration as obesity prevention strategies.\textsuperscript{1,2} Research shows a role for food price in dietary consumption\textsuperscript{3-7} and diet-related health outcomes, such as obesity.\textsuperscript{4,8} Economic theory suggests that lower-income individuals are more sensitive to price changes.\textsuperscript{9,10} Recent findings support greater response to food prices among sociodemographic subpopulations at higher risk for diet-related health outcomes, such as racial minority or low-income groups.\textsuperscript{11,12} However, most prior work on food price effects has been small experiments and cross-sectional studies and has yielded inconsistent results.\textsuperscript{3,8,13,14}

The fast food sector is frequently the target of policy discussion related to the literature linking fast food intake with excessive portion sizes, poor nutrient intake patterns, excessive caloric intake, and obesity.\textsuperscript{15-19} These studies have found distinctly different patterns of fast food access and intake among low-socioeconomic status (SES) subpopulations,\textsuperscript{20} leading cities (eg, Los Angeles, California) to ban the building of new fast food restaurants in selected low-SES areas\textsuperscript{21} or to suggest beverage portion size limitations (eg, New York, New York).

We quantified the associations between community-level fast food prices and individual-level fast food consumption, insulin resistance, and body weight within sociodemographic subgroups over 20 years of follow-up in a biracial cohort of young and middle-aged adults. We hypothesized greater price effects among blacks, as compared with whites, and among participants with relatively lower educational attainment or income. These questions are relevant to fiscal policy considerations\textsuperscript{12} and particularly to pricing strategies for eliminating disparities in dietary behaviors and health outcomes.

Methods

Study Sample
The Coronary Artery Risk Development in Young Adults (CARDIA) study is a multicenter, longitudinal study of cardiometabolic risk factors in adulthood.\textsuperscript{22} Black and white adults (aged 18-30 years; \( N = 5115 \)) were recruited from 4 metropolitan areas (Birmingham, Alabama; Chicago, Illinois; Minneapolis, Minnesota; Oakland, California). Participants were recruited for balance with respect to age, sex, race, and education. Of the surviving cohort, 81% were examined at a year 7 follow-up, 79% at year 10, 74% at year 15, and 72% at year 20. The study protocol was approved by institutional review boards of each participating institution; each study participant provided informed written consent.

Study Data

Dietary Assessment
Frequency of fast food consumption was assessed with the question, “How many times in a week or month do you eat breakfast, lunch, or dinner at a restaurant or cafeteria (eat-in or take-out)?” We note that participant responses to these 2 questions may not reflect mutually exclusive venue choices.

Clinical Measures
At each examination, weight (to 0.1 kg) and height (to 0.5 cm) were measured by trained clinical staff; body mass index (BMI) was calculated as weight in kilograms divided by height in meters squared. Fasting insulin levels and serum glucose concentrations were obtained from venous blood samples. Glucose concentration was measured using hexokinase coupled to glucose-6-phosphate dehydrogenase. The homeostasis model assessment of insulin resistance (HOMA-IR) score was calculated as [fasting glucose (in milligrams per deciliter) \( \times \) fasting insulin (in microunits per milliliter)] \( \div \) 405. Higher scores indicated increased insulin resistance.\textsuperscript{23}

Food Prices
Consumer price data were from the Council for Community and Economic Research (C2ER).\textsuperscript{24} Each quarter, the C2ER ascertained prices of more than 60 consumer goods and services to estimate cost of living for approximately 300 US cities.

The C2ER collected price data for 3 fast food items: a 12-in (30-cm) thin-crust cheese pizza from Pizza Hut or Pizza Inn, a \( \frac{1}{4} \)-lb (113-g) burger with cheese from McDonald’s, and 2 pieces of fried chicken from Kentucky Fried Chicken or Church’s Fried Chicken. We rescaled the price of pizza by dividing by 4 for serving-size comparability (on a calorie basis) with the burger and fried chicken and created a fast food index as the equal-weighted mean of the 3 fast food items. Our rescaling ensured that the price index was not dominated by pizza (\$4.92 at baseline vs \$1.66 for burger and \$1.66 for chicken).

We adjusted prices for inflation by dividing prices by the Bureau of Labor Statistics’s consumer price index (CPI),\textsuperscript{25} which has an index base period of 1982 to 1984 (CPI = 1.0). This adjustment removed effects of national-level secular price changes from those unique to each food item.

We linked C2ER data to participants on the basis of the quarter and year of their examinations and their geocoded residence at the time of each examination. Each CARDIA 1-year examination period covered 4 quarters. In addition to quarterly variability in local prices, variability in participant residence locations contributed to the distribution of food prices. We assigned price data to participants at the smallest geographic unit for which a match was available. At baseline, 24% of participants were assigned county-level price data, 49% Core-Based Statistical Area (CBSA)\textsuperscript{26}-level data, and 27% population-weighted state-level data. We note that although participants were recruited from 4 metropolitan areas, there was appreciable C2ER variability within those areas.

Individual-Level Covariates
Participants provided demographic and socioeconomic information on standardized questionnaires. At baseline, participants reported their age, sex, and race; at all examinations, participants reported their current educational attainment (years); and from year 5 onward, their family income (within categories).
Community-Level Covariates
We linked US Census data to each participant at the tract level, on the basis of participant residence at the time of the examination (using 1980 US Census for year 0, 1990 for years 7 and 10, and 2000 for years 15 and 20). Tract-level data improved distinction of social differences that may influence dietary behaviors, as compared with county-level data, and were more statistically reliable than block group-level data. We derived an index of neighborhood socioeconomic deprivation from a principal components analysis of 4 indicators: (1) percentage of population with income less than 150% of the federal poverty level, (2) median family income, (3) percentage of population with less than a high school education, and (4) percentage of population with at least a college degree.27

We estimated population density within a 3-km Euclidian buffer around participant residence as a weighted mean population count for block groups within the buffer, weighted by the proportion of total buffer area covered by each block group. The cost-of-living index was from the C2ER.

Statistical Analysis
We used repeated measures negative binomial regression to model longitudinal associations between community-level fast food price and weekly frequency of individual-level fast food consumption, and linear regression to model BMI and HOMA-IR. In secondary analysis, we quantified negative binomial associations between fast food price and frequency of eating at restaurants or cafeterias. Food price and outcome data were concurrently assessed at each examination year.

The negative binomial is a generalization of the Poisson model for count-dependent variables. The negative binomial introduces heterogeneity into the model that allows the mean and variance of the outcome to differ, thereby loosening the equidispersion assumption of the Poisson model. Dispersion parameters were significant in our models, indicating that the negative binomial was appropriate. The estimation method was maximum likelihood using Stata’s XTNBREG (negative binomial) and XTREG (linear) commands. We tested for heterogeneity in the relationship between fast food price and outcomes over study periods by including an interaction term for time. We log-transformed (ln) HOMA-IR scores and BMI to improve normality of the data and reduce the influence of right-skewed observations.

Because we considered price and all control variables exogenous, we used random effects to account for within-individual correlation of serial measures (ie, random-effects model). The random-effects specification allows for a 2-component error structure with time-invariant and time-varying individual specific errors. We adjusted for participant sex (male vs female), race (black vs white), baseline study center (4 cities), maximum reported income over follow-up (units of $10 000 from the midpoint of categorical responses), and maximum educational attainment (years) (both continuous); and time-varying age (6 categories: 18-24, 25-29, 30-34, 35-39, 40-44, 45-50 years), year of examination (5 categories: year 0, 7, 10, 15, 20), and geographic variables (all continuous, assigned to the individual), including neighborhood deprivation, cost-of-living index, and population density.

In a second modeling strategy, designed to account for possible correlation between the observed covariates such as price and time-invariant unobservables, we decomposed within-person and between-person variability. For fast food price and all time-varying covariates, this model (“within-person”) included both (1) the within-person mean across all 5 waves as an estimate of between-person effects and (2) the difference from the mean for each wave (ie, the time-varying part) as an estimate of within-person effects.28,29 We also adjusted for examination year and time-invariant covariates.

We tested for differences among sample subpopulations in their sensitivity to the price of fast food by including cross-product terms for fast food price and race, education, or income in the regression model. We ran separate interaction models for the following individual-level characteristics: race (2 levels: black vs white), highest attained education (2 levels: <16, ≥16 years), and highest reported income (3 levels: <$40 000, $40 000-$75 000, >$75 000), adjusting for the other 2 variables in the regression model (eg, in the price by race interaction model, we adjusted for education and income).

We used the MARGINS postestimation command in Stata to estimate multivariable-adjusted means and mean marginal effects (slopes) of price on study outcomes. Where there was evidence for nonlinear price effects, we estimated effects at decile cut points along the fast food price distribution (10th through 90th percentiles). To ensure that our estimates reflected the range of observed prices, we transformed fast food price to reflect a 2-SD unit increase in price.

We excluded observations (each participant had up to 5 observations) from analysis if the participant was pregnant at the time of examination (n = 120) or had missing data (n = 1446 fast food consumption, 10 fast food price, 12 neighborhood deprivation, 3 population density, 242 BMI, and 659 HOMA-IR). Two participants were missing data on sex, and 407 never reported their income. Analytic data sets include 18 300 observations for fast food consumption (from 4690 in year 0 to 2642 in year 20), 19 580 for BMI, and 19 094 for HOMA-IR.

Results
Over the 20-year follow-up, the prevalence and frequency of fast food consumption declined, as did the fast food price index; BMI and HOMA-IR increased (Table 1). Migration of the study sample and heterogeneity of food prices was such that 19% of the sample faced at least 1 between-survey price increase of at least $0.10 (1 SD of overall mean). The number of unique fast food prices varied by year, ranging from n = 60 at year 0 (mean [SD] price, $1.48 [$0.07]) to 483 at year 20 (mean [SD], $1.35 [$0.10]) [calculated at the community level; tabular data not shown].

For comparability with published literature, we quantified the price elasticity on consumption (where elasticity is the percent change in outcome per 1% change in price) at the combined-year mean fast food price ($1.39/serving). The multivariable-adjusted estimate from this analysis was −0.30 (95%
CI, –0.56 to –0.04), indicating a 0.30% decline in frequency of fast food consumption (times per week) for every 1% increase in fast food price.

Because fast food price was nonlinearly associated with consumption (P for quadratic price term <.001), we estimated effects at each decile cut point (10th through 90th percentiles) along the 20-year fast food price distribution. In age-adjusted and multivariable-adjusted models, mean frequency of fast food consumption was higher among blacks, compared with whites, and was also generally higher among participants who had lower educational attainment (<16 vs ≥16 years) (Table 2).

In general, results from random effects and within-person effects models did not differ substantively. Thus, we present estimates from the random-effects models for ease of interpretability. Within-person and between-person esti-
According to Decile Cut Points on the Fast Food Price Distribution

According to the study, the fast food price was inversely associated with HOMA-IR, with stronger estimated effects among those with lower income. The study also found that fast food price was inversely associated with BMI among sociodemographic groups, with inverse association in blacks but positive in whites. Furthermore, the study explored the 3-way interactions among fast food price and joint sociodemographic characteristics and found that the differences in fast food price sensitivity varied across different sociodemographic categories.

### Table 2. Age-Adjusted and Multivariable-Adjusted Subgroup-Specific Mean Fast Food Consumption Frequency According to Decile Cut Points on the Fast Food Price Distribution

<table>
<thead>
<tr>
<th>Characteristic</th>
<th>Fast Food Consumption Frequency, Mean (95% CI), Times/Wk</th>
</tr>
</thead>
<tbody>
<tr>
<td>Race</td>
<td></td>
</tr>
<tr>
<td>Age adjusted</td>
<td></td>
</tr>
<tr>
<td>Black</td>
<td>2.49 (2.34-2.63)</td>
</tr>
<tr>
<td>White</td>
<td>1.76 (1.65-1.86)</td>
</tr>
<tr>
<td>Multivariable adjusted</td>
<td></td>
</tr>
<tr>
<td>Black</td>
<td>2.20 (2.07-2.33)</td>
</tr>
<tr>
<td>White</td>
<td>1.55 (1.45-1.65)</td>
</tr>
<tr>
<td>Income&lt;sup&gt;a&lt;/sup&gt;</td>
<td></td>
</tr>
<tr>
<td>Age adjusted</td>
<td></td>
</tr>
<tr>
<td>&lt;$40 000</td>
<td>2.19 (2.00-2.37)</td>
</tr>
<tr>
<td>$40 000-$75 000</td>
<td>2.29 (2.14-2.44)</td>
</tr>
<tr>
<td>&gt;$75 000</td>
<td>1.95 (1.82-2.08)</td>
</tr>
<tr>
<td>Multivariable adjusted</td>
<td></td>
</tr>
<tr>
<td>&lt;$40 000</td>
<td>1.66 (1.52-1.80)</td>
</tr>
<tr>
<td>$40 000-$75 000</td>
<td>1.97 (1.84-2.09)</td>
</tr>
<tr>
<td>&gt;$75 000</td>
<td>1.92 (1.79-2.04)</td>
</tr>
<tr>
<td>Educational attainment&lt;sup&gt;b&lt;/sup&gt;</td>
<td></td>
</tr>
<tr>
<td>Age adjusted</td>
<td></td>
</tr>
<tr>
<td>&lt;16 y</td>
<td>2.35 (2.22-2.49)</td>
</tr>
<tr>
<td>≥16 y</td>
<td>1.83 (1.71-1.94)</td>
</tr>
<tr>
<td>Multivariable adjusted</td>
<td></td>
</tr>
<tr>
<td>&lt;16 y</td>
<td>2.05 (1.94-2.17)</td>
</tr>
<tr>
<td>≥16 y</td>
<td>1.67 (1.56-1.77)</td>
</tr>
</tbody>
</table>

<sup>a</sup> Age-adjusted models also include examination year. Multivariable-adjusted models include age, study center, examination year, sex, race, maximum educational attainment, highest reported income, population density, neighborhood deprivation score, and cost of living.

<sup>b</sup> Real (deflated to 1982-1984 prices) decile cut points for the 10th through 90th percentiles of fast food price were $1.25, $1.30, $1.32, $1.36, $1.38, $1.40, $1.47, $1.49, and $1.53. The numbers of observations within each decile were 1777, 1879, 1894, 1540, 1731, 1853, 1482, 1860, 2151, and 2203.

* Maximum reported during study.

The study found that fast food price was inversely associated with BMI at the highest income level and inversely associated with BMI at the middle income level (P < .001). Our findings did not materially change when we restricted the analysis to participants with county-level or CBSA-level fast food price data. We did not find statistical evidence of 3-way interactions among fast food price and joint sociode-
Differences in Fast Food Price Sensitivity

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Figure 1. Slopes for the Estimated Effect of Fast Food Price on Fast Food Consumption Frequency According to Sociodemographic Subgroups: Coronary Artery Risk Development in Young Adults (CARDIA) 1985/1986 to 2005/2006

A, Race; B, income; C, educational attainment. Slopes (marginal effects) from repeated-measures negative binomial regression models of fast food consumption on fast food price, adjusted for age, race, sex, examination year, study center, maximum educational attainment, highest reported income, population density, cost of living, and neighborhood deprivation. Effect estimates reflect change in fast food consumption per 2-SD unit ($0.20) change in fast food price. Random effects were used to account for within-person clustering over examination periods. Fast food prices were deflated to 1982 to 1984 values. Prices are shown at decile cut points (10th to 90th percentiles: $1.25, $1.30, $1.32, $1.36, $1.38, $1.40, $1.42, $1.49, $1.53). The mean (SD) frequency of fast food consumption was 1.80 (2.34) times per week. Interaction term P values: race, P = .04; income, P = .07; educational attainment, P = .03. P values for trend (by group): P < .001 (blacks), P = A (whites). P < .001 (income, $<40 000), P = .04 (income, $40 000-$75 000), P = .09 (income, $>75 000), P = .03 (education, <16 years), and P = .03 (education, ≥16 years).

Discussion

Over 20 years of follow-up in the CARDIA cohort, fast food price was inversely associated with frequency of fast food consumption, with greater price sensitivity among blacks, as compared with whites, and among those with lower educational attainment. We observed stronger inverse associations between fast food price and insulin resistance among those with lower educational attainment and among those at the middle income level. Fast food price was inversely associated with BMI in blacks and among those with lower educational attainment or at the middle income level but was positively associated with BMI among whites and among participants with higher educational attainment or at the highest income level. These findings have implications for fiscal policy considerations, particularly as they relate to sociodemographic disparities in fast food consumption and subsequent health outcomes. These findings are particularly important because blacks and lower-SES populations obtain greater proportions of their energy from fast food than do other race-ethnic and SES groups.30

In prior CARDIA analysis, Duffey et al4 documented significant inverse associations between soda and pizza prices and diet-related outcomes, including total energy consumption, BMI, and insulin resistance but did not study sociodemographic differences in price sensitivity. Understanding differential price sensitivity is essential to anticipating the impact of fiscal policies on population subgroups. According to economic theory, population subgroups with less disposable income would be expected to be more sensitive to price.9,10 Recent econometric research supports this expectation.8,12 A global ecologic meta-analysis of country-level food prices and consumption documented greater sensitivity among lower-income countries and households.10 Similarly, findings from New Zealand indicate that food pricing policies would have greater impact on the diets of low-income and racial minority (Maori) groups, with the potential to decrease health disparities in these groups.12 However, documentation of differential price effects on food consumption and biologic outcomes has been a gap in the epidemiologic literature noted by several researchers5-31-33 that we address in this analysis.

Corroborating our findings, in a study of C2ER food prices and fast food consumption over 5 years of follow-up in the National Longitudinal Study of Adolescent Health, Gordon-Larsen and colleagues2 found that blacks were more sensitive to the price of soda, whereas whites were more sensitive to the price of burgers. In contrast, a 4-year study of children found that fast food consumption among whites, but not blacks, was sensitive to the price of fast food.34 Study results also vary with respect to the influence of income on the effect of food price on dietary behavior, with some showing stronger inverse associations among low-income participants6,8,35 but others showing no difference across income categories.34,36,37 Similar to some of our findings, which showed the strongest price effects among those in the middle-income group, Finkelstein et al30 found that individuals in a middle (rather than lowest) income group were most sensitive to the price of sugar-sweetened beverages.

The positive associations between fast food price and BMI among whites and those with higher educational attainment and income were unexpected. These findings may reflect chance, although such unanticipated findings are not uncommon in the food price literature. For example, in an experi-
mental study, subsidies on fruits and vegetables led to increased caloric consumption, from both fruits and vegetables and less healthy options.\textsuperscript{39} Similarly, simulation studies, using published estimates of elasticities and other inputs, have found that taxes on food away from home\textsuperscript{40} or saturated fats\textsuperscript{41} may adversely affect dietary behaviors and weight-related outcomes. These studies point to complicated food behavior dynamics and underscore the importance of increased specificity in evaluations of proposed fiscal policies.

It is possible that differential substitution patterns help explain our BMI findings, which could be considered a type of confounding bias. An example would be if those with greater financial resources substitute other energy-dense foods for fast food meals. In our data, increases in fast food price were associated with decreases in visits to fast food establishments among blacks and those with lower educational attainment, which may yield overall greater improvements in insulin resistance within those population subgroups. We lack price data for eating at non–fast food restaurants and thus could not further evaluate these pricing dynamics. Accounting for cross-price elasticities (the impact of a specific food price on the consumption of other foods) is a pervasive challenge in food price-demand studies.\textsuperscript{42} Limitations of our data did not allow us to better delineate price-consumption dynamics.

Several other limitations of our study merit discussion. The C2ER provides the lowest-level geographic price data available, but even with the variability captured in our study (range of 60 unique price data points at year 0 to 483 at year 20), it is possible that price estimates may miss important regional variability. Price data may be measured with error, or the C2ER, which attempts to reflect the consumption patterns of a professional household, may not represent the purchases of the CARDIA sample. Furthermore, our models assume a temporal relation whereby price influences consumption, but prices may themselves be confounded by demand or other individual-level or community-level characteristics, although we controlled for many participant and contextual variables, including the neighborhood deprivation and regional cost-of-living index.

The CARDIA diet history asked about frequency of fast food consumption but did not assess what people ate at fast food establishments, and the C2ER price data may not align well with participants' consumption of specific foods. We were also limited in our ability to control for residential selectivity. We adjusted for baseline study center and for several neighborhood attributes of each participant's current residence, but, to the extent that people's choices about where to live correlate differentially with both fast food price and consumption, residual bias may exist. Another limitation, not limited to our study, was the lack of variability in fast food prices over the study period, which hindered our ability to estimate outcome changes across a broad range of fast food prices.

Despite these limitations, our study also has important strengths. As compared with previous studies, CARDIA has substantial power for detecting subgroup associations because of the large sample size, the racial and socioeconomic diversity of the sample, and temporal variability in food prices over 20 years.
years of follow-up. The longitudinal and observational nature of our study is important in the field of food price studies, many of which have been either cross-sectional or small experimental studies. Furthermore, CARDIA observed people through important life cycle stages and provided temporal and geographic variability in dietary habits. The lack of effects observed in some cross-sectional studies may reflect limited price variability. In contrast, experimental studies have been able to show effects by differentiating on price but may lack external variability, and it is often unclear whether interventions will play out as expected in observed populations.

Conclusions
The success of fiscal food policies requires understanding of how different populations will respond to changes in food prices and whether policies will have the desired impact on those in most need. Whereas many low-income and minority populations fight price increases because they note that they unduly affect the poor, it is equally important for us to remember that these reduced unhealthy behaviors can ameliorate health disparities and to call out for strong support for the inclusion of subgroup-specific elasticity estimates in model-based evaluations of policy approaches, as proposed by others. The major results of this study, the impact of higher fast food prices on reducing consumption of fast food and risk of type 2 diabetes mellitus among blacks and lower-SES subpopulations, speak to this issue directly. Counterintuitive findings in this study and others lead us to seek additional studies on these and related potentially unhealthy eating behaviors so that comprehensive food pricing policies will best achieve desired outcomes.

REFERENCES
Why the Cost of Fast Food Matters

Mitchell H. Katz, MD

Readers may justifiably wonder why the Editors chose this article for publication in a clinical journal. After all, as clinicians we are hardly in a position to influence the price of fast food.

However, obesity is a serious problem among our patients, and we have few good clinical interventions. We want our patients to make good food choices. None of us want to nag or to live in a “nanny” state where food choices are legislated. Rather, as we say in public health (and at home), we want the healthful choice to be the easy choice.

Here is the economic problem. Fast food has been getting cheaper. As the authors show, when fast food is cheap, people choose it more often. Blacks and persons with less education were more sensitive to fast food price, indicating that low-cost fast food may be contributing to disparities in obesity rates.

What can anyone do about this? Changing food subsidies and taxation can be used to influence price. Instead of subsidizing corn syrup, these benefits should accrue to fruits and vegetables. In this way, the government is not telling people what to eat; it is just making it easier to make the right choice.

That is a policy that could have tremendous clinical benefit for our patients.