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Fast food prices, obesity, and the minimum wage

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ABSTRACT

Recent proposals argue that a fast food tax may be an effective policy lever for reducing population weight. Although there is growing evidence for a negative association between fast food prices and weight among adolescents, less is known about adults. That any measured relationship to date is causal is unclear because there has been no attempt to separate variation in prices on the demand side from that on the supply side. We argue that the minimum wage is an exogenous source of variation in fast food prices, conditional on income and employment. In two-stage least-squares analyses, we find little evidence that fast food price changes affect adult BMI or obesity prevalence. Results are robust to including controls for area and time fixed effects, area time trends, demographic characteristics, substitute prices, numbers of establishments and employment in related industries, and other potentially related factors.

1. Introduction

The rise in obesity rates at the end of the twentieth and the beginning of the twenty-first centuries remains one of the major public health issues in the developed world today. Although the trend is a continuation of earlier increases in population weight (Komlos and Brabec, 2011), new challenges have emerged. Obesity is a concern to both health care professionals and policy makers because of its long term health consequences, including diabetes, cardiovascular disease, and cancer (National Task Force on the Prevention and Treatment of Obesity, 2000). Finkelstein et al. (2009) found that medical costs associated with obesity may have been $147 billion in 2008, or almost 10% of all U.S. medical expenditures. As a result of both the pecuniary and health costs associated with obesity, a large body of research has attempted to isolate the many contributing determinants of obesity in the United States.

Recently fast food consumption has received considerable attention in terms of its contribution to the rising prevalence of obesity. Cutler et al. (2003) characterize changes in weight as determined by changes in the balance of calorie intake and output. They conclude that caloric intake has increased by enough since the 1970s to explain the rising trend while caloric output has remained essentially unchanged, so consumption rather than output may be an appropriate policy target. For example, research conducted by Chou et al. (2004), Chen et al. (2009), Dunn et al. (2012), and Lhila (2011) suggest that *access* to fast food restaurants may have a small to moderate influence on the prevalence of obesity.[1] However, other evidence exploiting exogenous variation in the supply of restaurants

[1] Notably, Chen et al. (2009) and Dunn et al. (2012) both account for endogeneity when estimating the impact of fast-food *access* on obesity. The former utilizes a spatial autoregressive framework, while the latter employs an instrumental variables approach.

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(through the placement of Interstate Highways) casts doubt on this suggestion by finding either effects limited to a few population subgroups (Dunn, 2010) or no effects at all (Anderson and Matsa, 2011). In a similar vein, we focus attention on fast food prices as a potential policy lever. Specifically, several policy makers and some in the media have proposed taxes and restrictions on the fast food industry as a means of curbing fast food consumption (e.g. Lazarus, 2011; Lim, 2011). These articles suggest that the implementation of taxation on high calorie fast foods will reduce obesity rates by creating incentives to avoid fast foods.

Although it is intuitive that increasing the price of calorie dense foods would lead to healthier eating and weight outcomes, many factors contributing to overeating and obesity must be accounted for in order to isolate price effects. For example, Fig. 1 depicts annual deflated, normalized national food price indexes and the estimated median obesity rate among the states and Washington, DC between 1998 and 2010 (the restaurant series were first made available in 1998). The relationship between the Limited Service Restaurants and Snacks price index (LSRPI), which includes fast food restaurant prices, and obesity prevalence is unclear. Taking a close look, it appears that larger percent increases in the LSRPI tend to be accompanied by smaller percentage point increases in obesity prevalence. However, a broader perspective reveals that both the Food at Home and the Full Service Restaurants and Snacks series end up at a relatively *lower* level than the LSRPI across the time period, suggesting that an increase in the price of fast food relative to other foods – as would presumably be the case if fast food were taxed effectively – is associated with rising obesity. Neither of these interpretations tells a complete story, so careful research efforts are warranted to more clearly identify the relationship. In addition, theoretical models of consumption and weight accumulation do not yield definitive conclusions on the response of weight to price changes. For example, Anderson and Matsa (2011) show that a rational agent who consumes more restaurant calories may reduce other calorie consumption, and the inverse would also hold. Schroeter et al. (2008) develop a model demonstrating that, under certain conditions, a tax on high calorie foods can lead to an *increase* in body weight.

While some research on the effects of fast food prices on weight report a negative association between obesity and fast food prices among adolescents (e.g. Chou et al., 2005; Powell, 2009), other studies find that estimated fast food price coefficients are not statistically significant for other age groups (e.g. Chou et al., 2004, 2005). In light of research that leverages exogenous variation in the accessibility of restaurants to study its effect on obesity, we find it surprising that fast food *price* effects have not been studied in a similar way. Therefore, our goal is to contribute to the literature examining fast food prices and obesity by (1) exploiting an exogenous change in fast food prices such as what might arise from the taxation of fast food and (2) adding to existing knowledge about the relationship between fast food prices and obesity among adults.

In this study, we use two-stage least-squares (2SLS) methods that leverage variation in federal and state minimum wage mandates to address the possibility that ordinary least squares (OLS) estimates of fast food price effects yield biased estimates. Since fast food prices are equilibrium outcomes, it is likely that the observed relationship between prices and obesity prevalence across time and place does not simply reflect the response of weight to fast food price changes. Factors that mediate the

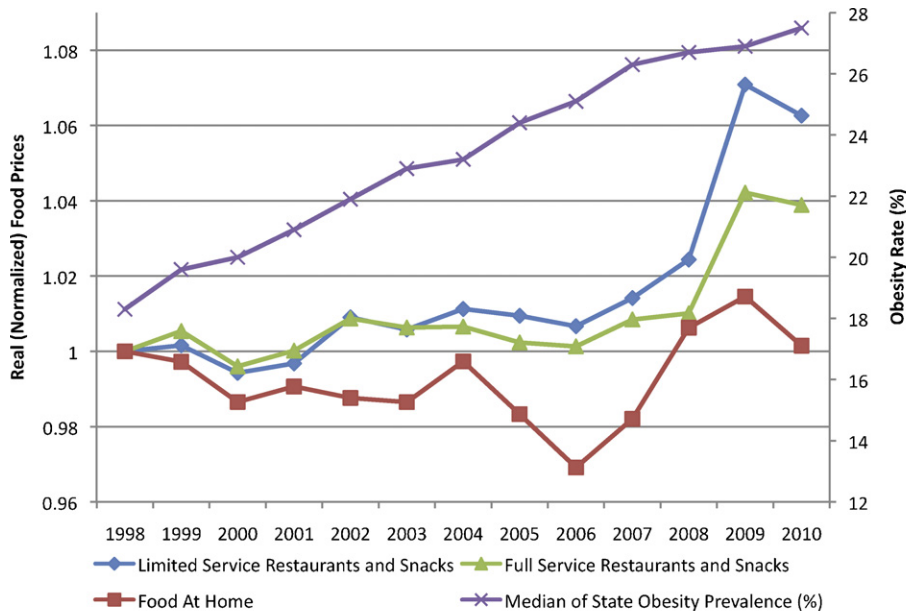

Fig. 1. Real (normalized) food prices and obesity prevalence in the United States, 1998–2010. *Notes*: The left axis measures three normalized food price indexes, deflated by the consumer price index. The price index series were obtained from the U.S. Department of Labor, Bureau of Labor Statistics. The right axis measures the median state-level prevalence of obesity estimated from the Behavioral Risk Factor Surveillance System, provided by the Office of Surveillance, Epidemiology, and Laboratory Services at the Centers for Disease Control and Prevention.

relationship, such as previously discussed fast food availability in terms of restaurant density and travel distance, will affect both weight and prices (e.g. Anderson and Matsa, 2011; Brennan and Carpenter, 2009; Currie et al., 2010; Dunn, 2010). In addition, unobserved heterogeneity in preferences for fast food, exercise, and overall health will shift the demand for fast and other foods as well as affect overall population weight. The validity of the estimation strategies implemented here rests on the claim that changes in the state and federal minimum wage act as exogenous shocks to county-level fast food prices and that, conditional on covariates capturing demographic, employment and income characteristics, minimum wage changes affect the obesity rate through fast food price changes. By controlling for characteristics through which the minimum wage affects the demand for fast food, notably income and employment status, we specifically attempt to isolate the supply-side effects of the minimum wage on fast food prices. We also test the soundness of this strategy through sensitivity analyses below.

The advantage of using the minimum wage as a source of exogenous variation is twofold: (1) there is an established literature that identifies the structural effects of minimum wage changes on fast food prices with consistent, strong results and (2) changes in the minimum wage are plausibly exogenous with respect to fast food prices, as discussed above, and can be implemented using established econometric techniques that address exogenous variation at the state and county levels. These key strengths of our empirical strategy allow us to incorporate a well established literature examining state level policies affecting fast food prices to estimate a price response that might occur in the context of a state level policy change with the goal of increasing fast food prices.

First, we find a marginally significant and small association between fast food prices and obesity in OLS estimation models. We are also able to demonstrate that (1) the minimum wage is a strong instrument for fast food prices conditional on the included covariates and (2) when the minimum wage is used as an instrument for fast food prices there is no statistical evidence to suggest that increasing fast food prices would play an important role in reducing BMI or obesity levels. Overall, the results suggest that public policies targeting fast food prices in isolation are not likely to be effective at significantly impacting the average weight of the adult population.

2. Background

2.1. Obesity

With regard to the principal issues at hand, the accumulated evidence exploring changes in obesity rates suggests that there is no single factor that wholly explains its rising prevalence. Cutler et al. (2003) focus attention on mechanisms that have increased calorie consumption over time, rather than decreased calorie expenditure. Lakdawalla and Philipson (2009) argue that technological change has both reduced food prices and made labor more sedentary. Chou et al. (2004, 2005) summarize a variety of associations, including income and demand responses such as relative food, cigarette, and alcohol prices that suggest a number of simultaneous trends that may have contributed to rising obesity rates.

Powell (2009) additionally controls for individual-level fixed effects using the 1997 National Longitudinal Survey of Youth (NLSY97) to focus on the association between fast food prices and obesity rates among adolescents, finding a negative relationship. This confirms a prior analysis by Chou et al. (2008) of the NLSY97 in the same context. With similar implications, Powell and Bao (2009) estimate a positive relationship between the price of fruits and vegetables and child BMI. Other studies examining fast food prices and weight find similar results for children or adolescents, in some cases in the upper quantiles of the BMI distribution, when considering data from Monitoring the Future (Powell et al., 2007; Auld and Powell, 2009) and the Medical Expenditure Panel Survey (Monheit et al., 2009). Although there is a growing body of evidence that higher fast food prices are associated with lower adolescent weight, there remains some question about the extent to which the relationship is causal because until now there has been no attempt to separate variation on the demand side from that on the supply side. In this paper we attempt to more closely identify such a relationship among a less studied group, namely persons age 18 and older.

In a closely related area of research, there are studies that consider the reduced form relationship between federal and state level minimum wage changes and obesity using data from the Behavioral Risk Factor Surveillance System (BRFSS). Meltzer and Chen (2009) find that a change in the real minimum wage is negatively associated with BMI and the probability of being obese among adults when controlling for demographic characteristics, notably excluding fast food prices. Jo and Lim (2009) find no association between the minimum wage and obesity prevalence among low income households when controlling for fast food prices, suggesting that minimum wages do not directly impact obesity. Taken together, these findings suggest that the minimum wage may affect weight directly through income for some low-wage workers and indirectly through fast food prices, but the magnitudes are unclear.[2] While we also present reduced form results for comparison with these previous studies, our goal here differs in that we attempt to isolate exogenous price effects using a 2SLS approach in order to inform a wider range of policies aimed at affecting fast food prices with the goal of reducing obesity prevalence.

2.2. Fast food prices and minimum wages

The traditional minimum wage literature tends to focus on the impact of minimum wage increases on employment, unemployment, and wages. Of course, the magnitude of minimum wage effects on unemployment and wages has been a long debated issue, and it is important to

[2] According to the U.S. Bureau of Labor Statistics, Current Population Survey (2010) workers with hourly wages at or below the Federal minimum wage made up 6% of all hourly-paid workers. Hourly-paid workers represent 58.8% of all wage and salary workers.

the surrounding policy discussion.[3] Yet, in recent years researchers have begun to investigate more of the indirect or second-order impacts of this labor-market mandate. Particular to this paper, we utilize changes in minimum wage laws, and the corresponding impact on fast food prices, to allow us to exogenously measure the impact of fast food prices on obesity.

There is a growing literature that analyzes the impact of minimum wage changes on retail prices with a particular focus on prices in the limited-service restaurant sector (which has the highest percentage of minimum wage workers of all four-digit industries listed in the North American Industry Classification System (NAICS)). In general, increases in the minimum wage lead to increased costs of operation in sectors that employ a high number of low-wage workers, and this may lead to a strong price pass-through effect. Card and Krueger (1995), Aaronson (2001), and MacDonald and Aaronson (2006) investigate this possibility. While Card and Krueger (1995) find mixed results, Aaronson (2001) and MacDonald and Aaronson (2006) find that minimum wage increases have an immediate and significant impact on prices in the limited-service restaurant sector. Specifically, using data on food prices from the American Chamber of Commerce Researchers Association (ACCRA), Aaronson (2001) finds that the prices of McDonald's hamburgers and KFC (or Church's) chicken rose by approximately 1.5% in response to a 10% increase in minimum wages. This outcome, combined with research papers that have identified a relationship between fast food prices and obesity, suggests that changes in minimum wages may indirectly play a part in determining obesity prevalence by impacting fast food prices.

3. Data

3.1. Minimum wage

This study utilizes the effective minimum wage in each state from 1990 to 2008, where the minimum wage variable is calculated as the greater of the state minimum wage (if one exists) and the federal minimum wage. Information on state minimum wages was collected from the material on state labor-law changes presented annually in the January edition of the *Monthly Labor Review*, along with previously published information on state minimum wages at the start of our sample period (Addison and Blackburn, 1999). We deflate minimum wages and all other dollar values used in the analysis (including fast food and food at home prices as well as incomes) to 2008 dollars using the national Consumer Price Index of all urban consumers and all items provided by the United States Department of Labor's Bureau of Labor Statistics (BLS).

Six separate increases in the nominal federal minimum wage occurred over our sample time frame. The clustered nature of these increases (1990 and 1991, 1996 and 1997, 2007, and 2008) has created a great deal of variation in the real federal minimum wage, which is highlighted by large stretches of decline between nominal increases. Federal increases in the minimum wage throughout the 1990s caused the average annual real minimum wage (in year 2008 dollars) to rise between \$5.53/h in 1990 to nearly \$6.54/h in 1998. By 2006, though, inflation had eroded the national minimum to approximately \$5.41/h, before recent federal increases returned the real minimum to levels similar to those seen in the late 1990s. Of course, the effective minimum wage observed in many states has often been higher than the federal mandate. At the beginning of 1990, there were fifteen states with minimum-wage levels above the federal mandate of \$3.35. Over the following 19 years there were over 100 increases in state-level minimum wages in which the resulting minimum wage was above the federal standard for some period of time.

3.2. Obesity data

We use data from the 1990 to 2008 waves of BRFSS to study the relationship between the minimum wages, fast food prices, and weight outcomes. BRFSS is administered by U.S. state and territory health departments and the Centers for Disease Control and Prevention (CDC) in order to monitor ongoing health risks. The time frame over which (self-reported) height and weight are reported corresponds with many changes in state and federal minimum wages. In addition, BRFSS includes basic demographic characteristics and information about a respondent's employment status and income.

We calculate BMI and weight categories for each individual in BRFSS, excluding pregnant women.[4] In our empirical analysis we include relevant individual demographic characteristics and state level control variables in addition to assigning the minimum wage rate for the individual's reported state of residence (described above). The demographic information used as control variables in the regression analysis includes sex, age, race/ethnicity, income, employment status, and educational attainment.

3.3. Price data

The analysis incorporates price data from the American Chamber of Commerce Researchers Association (ACCRA) that are drawn from price surveys used to compile the *ACCRA Cost of Living Index*. The surveys span the years 1990–2008 and prices in the surveys are collected quarterly. Specifically, we use the inflation-adjusted prices of two fast food items to examine price effects of the

[3] See Neumark and Wascher (2006) for a detailed summary of employment focused minimum wage literature.

[4] Cawley (2000) offers a method for adjusting height and weight for self-reporting bias that is based on both self-reports and physical measurements in the National Health and Examination Nutrition Survey (NHANES) III. Meltzer and Chen (2011) explain that NHANES and BRFSS have different sampling designs, and notably, BRFSS is conducted over the phone. We therefore adopt the unadjusted BMI values for our analysis. As a sensitivity analysis we estimated the primary models using the adjusted BMI values and found no qualitative differences in the results. The results are available on request.

minimum wage. These include a McDonald's *Quarter-Pounder with cheese*, and a fried chicken drumstick and thigh at Kentucky Fried Chicken and/or Church's Fried Chicken (with or without extras, whichever price is lower).[5] We also construct a food-at-home price index using ACCRA prices by following Chou et al. (2004) and Powell (2009) in calculating the average of thirteen grocery food items, weighted by their BLS expenditure shares.

An advantage of ACCRA price data is that they cover approximately 480 Core Based Statistical Areas (CBSAs), which allows for a large panel of localized prices that can be used to calculate average county-level prices. We provide a detailed description of the calculation of county-level prices in Appendix A. A disadvantage of the ACCRA data is that the raw data may be "noisy". In particular, the ACCRA data are collected by local Chamber of Commerce staff and may suffer from variable sample sizes over time and uneven data quality across CBSAs.[6] Using a price smoothing method proposed by Aaronson (2001) we verified that our results are robust to this potential shortcoming (these results are available on request).[7] In addition, the ACCRA data are not drawn from a collection of CBSAs intended to represent a nationally representative sample of the U.S. population.[8] In part to address concerns that the ACCRA data represent urban, higher income CBSAs we later include an analysis that considers a low income subgroup. The results, as shown, are consistent with those for the full sample.

3.4. Covariates

We control for a collection of demographic and household characteristics drawn from the BRFSS survey, including sex, age, race/ethnicity, income, employment status, and educational attainment. To calculate income, we assigned the midpoint of the BRFSS income category (one of seven) identified by the respondent, deflated to 2008 dollars. Aside from age, the other BRFSS control variables are coded as dichotomous variables. We also control for potentially relevant county and state level factors. For selected relevant industries (grocery stores, fitness and recreational sports centers, full-service restaurants, and limited-service eating places), we include the number of establishments per 10,000 residents and the percent of the population employed in each county. The industry data were collected from the 1990–2008 Quarterly Census of Employment and Wages (QCEW) published by BLS, and the county population data were provided by the United States Bureau of the Census (Census). State by year estimates of the percent of the population using food stamps were drawn from the 1990–2008 Annual Social and Economic Study of the Current Population Survey, jointly produced by BLS and Census. Lastly, we control for food sales tax rates across states and years, gathered from a combination of sources including the "Book of the States" (The Council of State Governments, 1990–2007), LexisNexis Academic searches, and state departments of revenue web sites.

3.5. Combined data

Table 1 reports summary statistics for the key data variables used in the analysis. Before dropping any observations with missing values the combined repeated cross-section of BRFSS respondents over the 18 year time period includes nearly 4 million observations. We successively dropped observations with missing values, yielding the following sample size decreases: a 829,466 (21%) decrease when dropping those with missing BRFSS state or county identifiers; a 464,031 (15%) decrease when dropping those with missing demographic characteristics including sex, age, race/ethnicity, education, income, and employment status; a 111,115 (4%) decrease when dropping those with missing BMI or who report they are pregnant; a 783,022 (31%) decrease when dropping those with missing values for county measures of employment and establishments in related industries; and a 1,017,500 (59%) decrease when dropping those with missing ACCRA price information. The size of the remaining sample used in the analysis is 711,081.

As described above, the ACCRA data are not intended to be representative of the entire U.S. population. After dropping observations with missing values, including missing ACCRA price information, we compared key characteristics of this sample with the original BRFSS sample. Compared with the original sample, individuals in the constructed sample are on average 2.5 percentage points more likely to be male, 1.49 years younger, 0.6 percentage points more likely to be black, 0.8 percentage points less likely to be Hispanic, 1.7 percentage points less likely to be a high school graduate, 5.0 percentage points more likely to be a college graduate, earn $3000 more in annual household income, and 5.3 percentage points more likely to be employed for wages. All differences are significant at the 1% level when conducting *t* tests. Overall, these differences are expected given that ACCRA is intended to allow price comparisons for urban, higher income households (as discussed earlier). Although it is evident from the mean comparisons that no relevant subgroup is wholly excluded from the final sample, in order to address concerns that the results do not represent effects for low income households we later conduct an analysis restricting the sample to low-income respondents and find that the results are indeed robust.

4. Empirical methodology

4.1. Primary model

The motivation for the empirical methodology is to take an alternative approach to the existing literature

[5] These are the two fast food prices measures that are found to be most sensitive to changes in minimum wages by Aaronson (2001).

[6] The 2SLS approach we utilize also helps mitigate possible attenuation bias that may be caused by measurement error in the ACCRA data.

[7] The price smoothing technique consists of replacing price changes greater than 5% that persist for less than three periods by linearly interpolated values.

[8] An overview of the ACCRA data includes the following description: "The ACCRA Cost of Living Index is designed to provide the best possible means to compare cost of living differences among urban areas based on the price of consumer goods and services appropriate for professional and managerial households in the top income quintile." Available at http://www.coli.org/Method.asp (last accessed 20.11.11).

Table 1
Summary statistics.

	Mean	Std dev
Body mass index (BMI)	26.357	5.187
Obese	0.196	–
Fast food price index (2008 dollars)	$2.92	0.217
Minimum wage (2008 dollars)	$6.46	0.608
Food-at-home price index (2008 dollars)	$2.50	0.222
Male	0.510	–
Age	44.107	16.793
Black	0.112	–
Hispanic	0.110	–
Graduated from high school	0.577	–
Graduated from college	0.320	–
Household income (in 1000s) (2008 dollars)	$54.97	30.872
Employed for wages	0.584	–
Self-employed	0.082	–
Out of work (more than 1 year)	0.016	–
Out of work (less than 1 year)	0.028	–
Homemaker	0.068	–
Student	0.043	–
Retired	0.146	–
Unable to work	0.033	–
Grocery stores: employment (% of county pop)	0.915	–
Fitness centers: employment (% of county pop)	0.169	–
Traditional restaurants: employment (% of county pop)	1.541	–
Fast food restaurants: employment (% of county pop)	1.448	–
Grocery stores: establishments (per 10,000 people)	2.505	0.881
Fitness centers: establishments (per 10,000 people)	0.847	0.333
Traditional restaurants: establishments (per 10,000 people)	6.101	1.824
Fast food restaurants: establishments (per 10,000 people)	7.875	1.748
Proportion of state population using food stamps	0.088	–
State sales tax rate on food	1.092	1.951

Note: N = 711,081. Individual level weight and demographic data are drawn from the 1990–2008 waves of BRFSS, and the means are estimated using the BRFSS survey weights. The price indexes are calculated using ACCRA price data. Establishment counts and employment are drawn from the Quarterly Census of Employment and Wages. The proportions using food stamps are estimated using data from the CPS March Supplement. The minimum wages and state food sales tax rates were compiled in independent efforts, as described in the text. All dollar values are deflated using the CPI from BLS.

in the form of a 2SLS analysis of the effect of fast food price changes on weight. In determining the most appropriate empirical approach, we recognize that an individual-level analysis, such as that estimated by Chou et al. (2004) is most common in the obesity literature. We also conducted a county-level analysis that implemented a similar methodology to that of Aaronson (2001). Unlike Aaronson (2001), which only looked at the effects of minimum wages on prices, our parallel analysis also consisted of estimating the response of changes in average county BMI to changes in county-level prices. Although the first-stage of our analysis reproduced the major findings in Aaronson (2001) and yielded estimates that were qualitatively similar to those presented here, we do not report them because of concerns that the BRFSS sampling design does not readily allow for accurate county-level aggregate estimates (Gregg et al., 2009).[9]

Generally, we conduct an OLS analysis relating BMI to fast food prices followed by 2SLS analyses where fast food prices are instrumented for using the minimum wage. Since BRFSS consists of repeated cross section data on multiple observations that can be geographically separated, and because real fast food prices vary within counties over time, a county and time (year and quarter) fixed effects approach was selected.[10] The general econometric model is as follows:

$$Y_{icqt} = FF_{cqt} * \beta + X_{icqt} * \gamma + Z_{cqt} * \psi + W_{sqt} * \theta + a_c + \tau_t + (\alpha_c * t) + \delta_q + \varepsilon_{ict} \quad (1)$$

where Y_{icqt} is the weight outcome (BMI or obesity status) of individual i in county c during year t and quarter q, FF is the county-level real fast food price index (in 2008 dollars), X is a vector of individual level characteristics (e.g. race/ethnicity, education, etc.), Z is a vector of county level characteristics (e.g. food-at-home price index, the number of grocery stores and fast food restaurants, etc.), W is a vector of state level characteristics (percent of the population using food stamps, food sales tax rates, etc.), α is a vector of county fixed effects, τ is a vector of year fixed effects, δ is a vector of quarter fixed effects ($\alpha_c{}^*t$) accounts for the county-specific time trend, and ε is the idiosyncratic error term.[11] The county fixed

[9] The results of this analysis are available on request.

[10] We also verified that our results are qualitatively robust to the inclusion of period (year-quarter) fixed effects, rather than controlling for year and quarter separately.

[11] We recognize that the use of an OLS specification with limited dependent variables (e.g. obesity status) can suffer from bias and may provide unreasonable estimates for certain values of the independent variable. All regressions utilizing a limited dependent variable were also estimated using a probit specification in order to verify the validity of the linear probability estimates, but excluded due to a preference for simplicity and ease of interpretation. Probit estimates are available from the authors upon request.

effects control for any persistent aspects of a county that might affect an individual's weight, including constant differences in institutional, economic, cultural, or demographic characteristics that might influence their diet, exercise, or other behaviors that could impact a local population's BMI. The year fixed effects control for any national-level macroeconomic effects that consistently influence all locations, while the quarter effects control for seasonal aspects. The net result of the county and time fixed effects is to focus attention on changes in weight outcomes over time within a particular location.[12]

Although Eq. (1) captures any fixed factors that may cause BMI to vary across people in different counties, there will be changes over time within counties in other important determinants of BMI that must be captured using additional controls. Most of these variables come from the BRFSS data and have previously been identified as having a significant influence on weight outcomes. These include an individual's age, income, and dummy variables for race, ethnicity, education, employment status. In addition, county-level data on the number of grocery stores, fitness centers, fast food restaurants, traditional restaurants per 10,000 residents and the corresponding employment as a percent of the population have also been incorporated.[13] Next, the proportion of a state that is eligible for food stamps and each state's food sales tax rate are incorporated. Lastly, an index of food-at-home prices, discussed earlier, is also included.

One concern is that the identification strategy outlined to this point is predicated on the assumption that, after the inclusion of fixed effects and time-varying controls, individuals in different counties that are impacted by changes in fast-food prices (the treatment group) are comparable to individuals in counties that are not impacted (the control group). Yet, there is the potential that unobserved characteristics in different counties that vary over time are related to fast-food prices and BMI. To address this limitation, we have also included controls for county-specific time trends in Eq. (1).

Lastly, the inclusion of county fixed effects, time fixed effects, and county-time trends does not prevent correlation in Eq. (1) error term, ε, within counties over time, or across counties at a point in time (for example, see Arellano, 1987). Ignoring this correlation can lead to biases in the standard errors, especially when the level of aggregation varies across the variables included in the model (see Bertrand et al., 2004). As a result, all standard errors have been clustered to allow for any type of correlation structure among the error terms for a given county.

4.2. Model including lagged independent variables

We extend our analysis by implementing an empirical strategy that utilizes lagged as well as contemporaneous fast food prices to capture the effects of fast food prices on obesity over time. Since weight gain is frequently modeled as a cumulative process (see Cutler et al., 2003; Schroeter et al., 2008 as examples) both lagged and contemporaneous prices may be appropriate to include when modeling the impact of fast food prices on BMI. Aaronson (2001) also demonstrates that there may be lagged effects of minimum wage changes on fast food prices in accordance with theories that firm output prices may not instantaneously respond to increased input costs. We therefore also include both lagged and contemporaneous minimum wages.

Estimation of the extended model is challenged by a tradeoff between a richer set of lagged independent variables and a shorter sample period, particularly when estimating the 2SLS models. Whenever incorporating lagged fast food prices we add three semi-annual (i.e. occurring every two quarters) lags which allow us to detect lagged price effects over the span of 1.5 years. We chose not to include quarterly lags over this time span because in the 2SLS models each endogenous fast food price variable requires an instrument, and this would have required us to include six minimum wage lags. This would shorten the sample period by six years since minimum wage is measured annually. Instead, by incorporating semi-annual price lags we need only three annual minimum wage lags to identify the model, thus mitigating the sample period reduction. For the 2SLS estimation this framework allows us to account for 1.5 years of lagged price effects and at least 1.5 years, but up to 3 years, of lagged minimum wage effects (the span of the lagged minimum wage effects varies by endogenous variable).[14]

5. Results

5.1. OLS analysis

In Table 2 we report OLS results for the primary and lagged models, which describe the relationship between BMI or obesity and the relevant covariates discussed above (including fast food prices). Panel A presents estimates

[12] If federal changes in the minimum wage were binding across all states, then this change would be completely absorbed by the time-fixed effect. As this is not the case, changes in the federal minimum wage mandate are still relevant to our analysis.

[13] All establishment data is incorporated in to the regression analysis as the number of establishments per 10,000 people in a county. All employment data is incorporated as the percentage of the population that works in each industry respectively.

[14] We also tested for the optimal number of lags by regressing the fast food price index on the contemporaneous minimum wage and a varying number of lags and by regressing BMI on the contemporaneous fast food price index and a varying number of lags (while maintaining the same sample across regressions). Both Akaike's and Schwarz's Bayesian information criteria (AIC and BIC) were considered in selecting the optimal number of lags. A smaller AIC or BIC value indicates a better fit (StataCorp, 2009). We were unable to use this method to test the 2SLS specifications directly because 2SLS is not estimated using maximum likelihood methods. The results indicated that a larger number of minimum wage lags improved the fit of the first model while a smaller number of fast food price index improved the fit of the second. Therefore, in addition to estimating models that include both lagged fast food prices and lagged minimum wages as reported in the text and tables, we also conducted the 2SLS estimation using the contemporaneous fast food price only with contemporaneous and lagged minimum wages. The results are qualitatively similar to models not including any lags, and they are available on request.

Table 2
OLS analysis: the impact of fast food prices on weight.

	A. Contemporaneous analysis		B. Lagged analysis	
	BMI	Obese	BMI	Obese
Fast food price index (2008 dollars)	−0.080*	−0.006*	0.044	0.002
	(0.043)	(0.003)	(0.074)	(0.006)
Fast food price index (2008 dollars) (2-quarter lag)	...	...	−0.079	−0.011
			(0.087)	(0.007)
Fast food price index (2008 dollars) (4-quarter lag)	...	...	−0.120	−0.006
			(0.091)	(0.007)
Fast food price index (2008 dollars) (6-quarter lag)	...	...	−0.036	0.004
			(0.094)	(0.007)
Food-at-home price index (2008 dollars)	−0.021	−0.001	0.013	0.004
	(0.063)	(0.005)	(0.103)	(0.008)
Male	1.109***	0.012***	1.115***	0.012***
	(0.029)	(0.002)	(0.034)	(0.002)
Age	0.308***	0.016***	0.306***	0.016***
	(0.003)	(0.000)	(0.003)	(0.000)
Age^2	−0.003***	−0.0002***	−0.003***	−0.0002***
	(0.000)	(0.0000)	(0.000)	(0.0000)
Black	1.906***	0.106***	1.902***	0.104***
	(0.040)	(0.003)	(0.050)	(0.003)
Hispanic	0.689***	0.029***	0.663***	0.025***
	(0.049)	(0.003)	(0.058)	(0.004)
Household income (in 1000's of 2008 $)	−0.009***	−0.001***	−0.009***	−0.001***
	(0.000)	(0.0000)	(0.000)	(0.0000)
Educational status:				
Less than high school	#	#	#	#
Graduated from high school	−0.349***	−0.029***	−0.367***	−0.031***
	(0.027)	(0.002)	(0.038)	(0.003)
Graduated from college	−1.122***	−0.078***	−1.137***	−0.080***
	(0.033)	(0.002)	(0.041)	(0.003)
Employment status:				
Employed for wages	−0.118**	−0.011***	−0.070	−0.007
	(0.050)	(0.004)	(0.069)	(0.005)
Self-employed	−0.585***	−0.040***	−0.509***	−0.037***
	(0.055)	(0.004)	(0.080)	(0.006)
Out of work (more than 1 year)	0.225***	0.017***	0.322***	0.023***
	(0.075)	(0.006)	(0.101)	(0.008)
Out of work (less than 1 year)	#	#	#	#
Homemaker	−0.558***	−0.033***	−0.522***	−0.031***
	(0.056)	(0.004)	(0.074)	(0.005)
Student	−0.807***	−0.038***	−0.804***	−0.037***
	(0.068)	(0.005)	(0.087)	(0.006)
Retired	−0.090	−0.008*	−0.038	−0.005
	(0.056)	(0.004)	(0.078)	(0.006)
Unable to work	1.249***	0.085***	1.292***	0.085***
	(0.066)	(0.005)	(0.098)	(0.007)
Grocery stores: employment % of county population	0.017	0.005	0.153	0.019
	(0.100)	(0.007)	(0.129)	(0.011)
Fitness centers: employment % of county population	−0.052	−0.017	−0.106	−0.019
	(0.210)	(0.017)	(0.314)	(0.024)
Traditional restaurants: employment % of county population	0.010	−0.001	0.016	−0.001
	(0.055)	(0.005)	(0.086)	(0.007)
Fast food: employment % of county population	−0.020	0.000	0.140	0.005
	(0.062)	(0.005)	(0.119)	(0.010)
Grocery stores: establishments per 10,000 people	−0.047	−0.004	−0.034	−0.003
	(0.033)	(0.003)	(0.053)	(0.004)
Fitness centers: establishments per 10,000 people	0.123*	0.005	0.116	0.007
	(0.065)	(0.005)	(0.111)	(0.009)
Traditional restaurants: establishments per 10,000 people	−0.019	−0.000	−0.057*	−0.003
	(0.023)	(0.002)	(0.034)	(0.003)
Fast food: establishments per 10,000 people	−0.002	0.002	−0.012	0.001
	(0.020)	(0.002)	(0.029)	(0.002)
Proportion of state population using food stamps	−0.171	−0.023	0.036	−0.001
	(0.540)	(0.044)	(0.904)	(0.066)
State food sales tax rate	0.012	−0.001	0.037	0.002
	(0.012)	(0.001)	(0.023)	(0.002)

Table 2 (*Continued*)

	A. Contemporaneous analysis		B. Lagged analysis	
	BMI	Obese	BMI	Obese
Sample size:	711,081		366,114	
Counties:	625		523	
Sum of fast food price coefficients:	. . .	. . .	−0.191	−0.011
P-value: test of significance			0.1476	0.2794

Notes: Each column represents a separate regression. All regressions include indicator variables for county, year, and quarter; controls for county time-trends; and all covariates described in Table 1.
symbol denotes that this row represents the omitted value of the categorical variable when the coefficients were estimated.
*** Statistical significance at the 0.01 level.
** Statistical significance at the 0.05 level.
* Statistical significance at the 0.10 level.

showing the contemporaneous impact of changes in fast food prices and other covariates on BMI and obesity, while Panel B presents the analogous estimates for the lagged model. Results indicate that demographic factors such as age, race/ethnicity, education, gender, employment status, and income have effects on BMI and obesity that are consistent with previous findings (e.g. Chou et al., 2004) and similar across models. Moreover, estimates suggest that the number of fitness centers per capita (controlling for the percent of the population employed in the fitness center industry) is potentially related to BMI and obesity. We also find some evidence that local fast food prices may have an effect on BMI and obesity since a one dollar increase in the fast food price index is associated with a decrease of 0.08 BMI points and a 0.6 percentage point decrease in the likelihood of being obese. However, these estimates are significant only at the 10% level and are small in magnitude.

To further explore the relationship between fast food prices and obesity, in Table 3 we conduct an analysis of the effects of fast food price changes on consumption of food items that one might expect to be directly or indirectly impacted. Specifically, we look at how price changes impact the self-reported consumption of hamburgers, fried chicken, fruits and vegetables, and overall grams of fat. In both the contemporaneous and lagged models we find almost no statistical evidence that consumption of these items is impacted by changes in fast-food prices. The one exception is a large and significant reduction in the total number of grams of fat per day (more than 70) in response to a sustained increase in the fast food price index. This finding is less reliable, however, in light of the fact that the contemporaneous model (where one would expect to see a consumption response to a price change) does not yield a significant coefficient estimate. Because BRFSS includes questions from the Dietary Fat module for a limited duration (between 1990 and 1994) and only for some geographic areas in our sample time frame, the sample sizes are quite small. We are therefore cautious in placing too much emphasis on these estimates. Nevertheless, if we take the estimates at face value they suggest that consumers are relatively price inelastic (at least at these low price levels) with respect to changes in fast food prices.

5.2. 2SLS approach

The OLS results presented in Table 2 provide inconclusive evidence for the role that fast food prices play in

Table 3
OLS analysis: the impact of fast food prices on consumption.

	Fat consumed per day (g)	Hamburgers	Fried chicken	Fruits and vegetables
A. Contemporaneous analysis				
Fast food price index (2008 dollars)	1.609	2.117	0.139	1.783
	(3.992)	(1.371)	(1.163)	(1.384)
Sample size:	6,286	6039	4110	388,968
Counties:	56	56	56	545
B. Contemporaneous and lagged analysis				
Fast food price index (2008 dollars)	−28.630*	4.877	7.658**	1.905
	(16.273)	(5.059)	(3.188)	(2.075)
Fast food price index (2008 dollars)	−41.569**	−7.666	−3.348	0.256
(2-quarter lag)	(19.624)	(7.282)	(5.264)	(0.925)
Fast food price index (2008 dollars)	−6.240	6.169*	14.023**	0.804
(4-quarter lag)	(11.807)	(3.597)	(5.291)	(1.965)
Fast food price index (2008 dollars)	4.898	0.023	1.673	−2.305*
(6-quarter lag)	(7.154)	(1.813)	(2.204)	(1.325)
Sample size:	2864	2757	1870	206,760
Counties:	31	31	31	457
Sum of fast food price coefficients:	−71.541**	3.403	20.006	0.66
P-value: test of significance	0.0144	0.7660	0.3397	0.8443

Notes: Each column of each panel represents a separate regression. All regressions include indicator variables for county, year, and quarter; controls for county time-trends; and all covariates described in Table 1.
** Statistical significance at the 0.05 level.
* Statistical significance at the 0.10 level.

Table 4
Reduced form analysis: impact of changes in minimum wages on BMI and obesity.

	A. Contemporaneous analysis		B. Lagged analysis	
	BMI	Obese	BMI	Obese
Minimum wage (2008 dollars)	−0.015	−0.0002	−0.012	−0.001
	(0.028)	(0.0018)	(0.029)	(0.002)
Minimum wage (2008 dollars) (1-year lag)	...	...	−0.042	0.002
			(0.060)	(0.003)
Minimum wage (2008 dollars) (2-year lag)	...	...	0.168**	0.006
			(0.068)	(0.005)
Minimum wage (2008 dollars) (3-year lag)	...	...	−0.098	−0.005
			(0.077)	(0.005)
Sample size	711,081		697,511	
Number of counties	625		621	
Sum of minimum wage coefficients:				
P-value: test of significance	...	...	0.016	0.002
			0.6095	0.7385

Notes: Each column represents a separate regression. All regressions include indicator variables for county, year, and quarter; controls for county time-trends; and all covariates described in Table 1.

determining BMI or obesity because an interpretation that these estimates represent a causal relationship would require the assumption that fast food price is exogenously determined. As discussed earlier, local fast food prices and BMI or obesity may be endogenous. Given these concerns, we turn to estimating the impact of fast food price changes on BMI and obesity using a 2SLS approach. This method requires an instrument that determines fast food prices, but that does not directly determine BMI. Importantly, given that we are using a fixed effects framework that includes linear trends, we must also use an instrument that varies non-linearly over time. We argue that, after controlling for demographic characteristics and other covariates including household income and employment status, the inflation-adjusted minimum wage satisfies these criteria.[15]

As we have discussed, previous research has identified a consistent link between minimum wage increases and increases in fast food prices (e.g. Aaronson, 2001; MacDonald and Aaronson, 2006). Moreover, we subject the minimum wage to weak instrument tests, which are discussed below. Verifying that minimum wage changes do not directly impact BMI is more challenging. In recent work, Meltzer and Chen (2009) found that increases in minimum wages have been associated with reduction in BMI, but Jo and Lim (2009) found that this outcome disappears once fast food prices are controlled for in their analysis, indicating that minimum wages may operate through fast food prices. In the event that the OLS estimates of fast food price effects on weight are biased, this could additionally bias the estimation of other correlated covariates (namely the minimum wage). Therefore, Jo and Lim's (2009) finding may only be suggestive. In order to determine if there is any identifiable direct relationship between minimum wage changes and obesity levels, we conducted a reduced form analysis where we simply substitute the state minimum wage for the fast food price index in our model presented in Eq. (1). Results are presented in Table 4, and it is clear that after the inclusion of demographic controls, county and state level covariates, fixed effects and time trends, there is no measurable impact of minimum wage changes on BMI or obesity in either the contemporaneous or lagged models. As a result, we have no strong reason to suspect that the minimum wage directly impacts BMI or obesity.

Table 5 presents results from regressions that demonstrate the impact of minimum wages on fast food prices.[16] Overall, the highly significant results suggest that the minimum wage is not a weak instrument (we formally test this below). Column A reports the contemporaneous effect of minimum wages on fast food prices, while Column B reports analogous contemporaneous and lagged effects. Similar to previous literature discussed earlier, the findings overall show a positive and highly significant relationship between the minimum wage and fast food prices. Column A shows that a one dollar increase in the minimum wage (in 2008 dollars) results in a seven or eight cent increase in the fast food price index. For Column B we report the sum of the minimum wage effects and corresponding *p*-values for the fast food price index in order to test the total effect of an extended change in the minimum wage regime. These results similarly suggest that a one dollar increase in the minimum wage would result in an increase in the fast

[15] This is the greater of the inflation-adjusted state or federally mandated minimum wage level in any location.

[16] One concern that would cast doubt on the validity of the minimum wage as an instrument is the possibility that minimum wage changes are directly affected by prices through increased political awareness during periods of higher inflation. Aaronson (2001) tests this by examining inflation during periods prior to minimum wage legislation enactments and finds no evidence of policy endogeneity. We confirm this finding by regressing the deflated minimum wage on two yearly lags of the deflated fast food price index in addition to state and year fixed effects and state-specific time trends, with standard errors clustered by state. The coefficients on the two lags are both positive, but neither is statistically significant.

Table 5
Impact of changes in minimum wages on fast food prices.

Instruments	A. Contemporaneous analysis Fast food price index	B. Lagged analysis Fast food price index
Minimum wage (2008 dollars)	0.075***	0.069***
	(0.016)	(0.016)
Minimum wage (2008 dollars)	...	0.021
(1-year lag)		(0.016)
Minimum wage (2008 dollars)	...	−0.009
(2-year lag)		(0.026)
Minimum wage (2008 dollars)	...	0.042
(3-year lag)		(0.035)
Sample size	711,081	697,511
Number of counties	625	621
Sum of minimum wage coefficients:	...	0.123***
P-value: test of significance		0.0013

Notes: Each column represents a separate regression. All regressions include indicator variables for county, year, and quarter; controls for county time-trends; and all covariates described in Table 1.
*** Statistical significance at the 0.01 level.

Table 6
Second-stage 2SLS results: the impact of fast food prices on body mass index.

	A. Contemporaneous analysis		B. Lagged analysis	
	BMI	Obese	BMI	Obese
Fast food price index (2008 dollars)	0.165	0.002	0.096	−0.003
	(0.144)	(0.011)	(0.372)	(0.030)
Fast food price index (2008 dollars)	...	...	0.902*	0.126***
(2-quarter lag)			(0.535)	(0.045)
Fast food price index (2008 dollars)	...	...	−0.973	−0.162***
(4-quarter lag)			(0.749)	(0.057)
Fast food price index (2008 dollars)	...	...	0.124	0.054
(6-quarter lag)			(0.491)	(0.036)
Sample size	711,081	360,649		
Number of counties	625	521		
Sum of fast food price coefficients:			0.149	0.015
P-value: test of significance	...	...	0.6378	0.5592
Kleibergen–Paap *F* statistic (weak instrument test):	26.231		8.87	

Notes: Each column represents a separate regression. All regressions include indicator variables for county, year, and quarter; controls for county time-trends; and all covariates described in Table 1. The Kleibergen–Paap *F* statistic reported for the lagged analysis is that associated with the contemporaneous first stage model (see text for more details).
*** Statistical significance at the 0.01 level.
* Statistical significance at the 0.10 level.

food prices of approximately twelve cents, with over half of the estimated increase occurring immediately.

The coefficients of interest from the second stage of the 2SLS estimation as well as results from additional tests are presented in Table 6. First, the heteroskedasticity-robust Kleibergen–Paap *F* statistics suggest that the minimum wage and respective lagged values pass weak instrument tests according to critical values computed by Stock and Yogo (2005).[17] Turning to the main results, we find that the fast food price index coefficients in Table 6 provide no evidence that fast-food prices negatively impact BMI or obesity.[18] Moreover, in all but one case, the estimated coefficient on price is actually positive, although never close to statistically significant. Overall, there is no strong or consistent evidence here to suggest that fast food prices impact average BMI or obesity prevalence in a meaningful way.

As a means for comparing the reasonableness of our OLS and subsequent 2SLS results, we can turn to estimates in

[17] The Kleibergen-Paap *F* statistic reported for the lagged models is that for the first-stage estimation of the contemporaneous fast food price. However, this statistic may be overstated if there are multiple endogenous variables (StataCorp, 2009). A more appropriate weak instrument test statistic for multiple endogenous variables is the Cragg and Donald (1993) minimum eigenvalue statistic, but only when estimating models that do not include heteroskedasticity-robust standard errors (or clustered standard errors, by implication).

[18] We also verified that our results are qualitatively robust to the removal of county- and state-level covariates. For example, fast food availability has been previously identified as endogenous (Dunn, 2010; Anderson and Matsa, 2011). Both the first-stage and IV results are nearly identical when both including and excluding covariates that measure the quantity and employment proportions of restaurants and other related establishments.

the existing literature to estimate a rough ''back-of-the-envelope'' prediction of the implied impact of fast-food price changes on weight. Specifically, a one dollar increase in the minimum wage is predicted to increase fast-food prices by \$0.08. Given that the average fast-food index item price in our data is \$2.92; this would suggest a 2.75 price increase. Park et al. (1996) estimate an own-price demand elasticity of −1.0 for restaurant food, which would suggest that a 2.7% increase in price would result in a 2.7% decrease in consumption. Anderson and Matsa (2011) find that the marginal caloric difference in eating a meal away from home is 35 calories. A 2.7% reduction in consumption would correspondingly reduce caloric intake on average by 1 calorie per meal. According to Cutler et al. (2003), a reduction of daily caloric intake of one calorie per day, in perpetuity, would reduce an individual's steady-state weight by approximately 0.08 pounds. Hence, this rough calculation is consistent with our results, as it would not predict a meaningful impact of fast-food price changes on BMI or obesity arising from a change in the minimum wage and it subsequent impact on fast food prices.

Next, we are concerned that the results from a sample of the entire adult population may be masking important effects within certain subgroups. In particular, individuals who have a lower ability to pay will be more price-sensitive and these effects are covered up by the larger sample. To address this concern we investigated the effects of fast-food prices on BMI and obesity of individuals under 25 years old without a college degree and whose annual income is less than \$25,000 separately from older or higher income populations. The results, which are presented in Table 7, show no evidence to suggest that individuals who likely have a lower willingness to pay (listed as low income in Table 7) are impacted differently from the entire population.

Lastly, while these results do not address the estimated impacts of fast-food prices on adolescent weight that is observed in the literature (e.g. Chou et al., 2008; Powell and Bao, 2009), there remains some question about the extent to which the relationship is causal. This paper identifies that weak negative estimates of fast-food prices (shown in Table 2) disappear when fast-food prices are instrumented for effectively. Moreover, it may also be possible that parents are simply more responsive to changes in fast-food prices when purchasing for a family or when considering the health of their children, relative to healthier food options that are more price competitive. Nevertheless, this particular issue is outside the scope or focus of this project but is a potentially important question for future research.

Table 7
Second-stage 2SLS results: the impact of fast food prices on body mass index: low income sample.

	Low income sample	
Instruments	BMI	Obese
A. Contemporaneous effects		
Fast food price index, deflated	2.729	0.151
	(1.787)	(0.117)
Sample size	16,383	
Number of counties	521	
Kleibergen–Paap *F* statistic (weak instrument test):	23.3229	
B. Contemporaneous and lagged effects		
Fast food price index, deflated	0.873	−0.042
	(2.119)	(0.161)
Fast food price index (2008 dollars) (2-quarter lag)	−3.900	−0.280
	(2.854)	(0.265)
Fast food price index (2008 dollars) (4-quarter lag)	6.195**	0.688
	(2.859)	(0.271)
Fast food price index (2008 dollars) (6-quarter lag)	−5.682***	−0.518***
	(1.630)	(0.137)
Sample size	8382	
Number of counties	416	
Sum of fast food price coefficients:	−2.514	−0.152
P-value: test of significance	0.2868	0.4375
Kleibergen–Paap *F* statistic (weak instrument test):	9.7136	

Notes: Each column represents a separate regression. All regressions include indicator variables for county, year, and quarter; controls for county time-trends; and all covariates described in Table 1.
*** Statistical significance at the 0.01 level.
** Statistical significance at the 0.05 level.

6. Conclusion

The problems arising from obesity have been heavily discussed in the U.S. over the last decade. Accordingly, researchers have attempted to understand the underlying determinants of obesity in order to better inform policy makers about how to reduce obesity prevalence. Recently there has been increasing discussion by lawmakers about implementing taxes on foods and beverages that likely contribute to obesity (e.g. soda and fast food). This paper utilizes changes in minimum wages to investigate whether exogenous changes in fast food prices impact BMI at the individual level. Results suggest that, once potential endogeneity is accounted for, changes in fast food prices do not have a meaningful impact on BMI among adults.

We have shown that this result is robust to the inclusion of controls for area and county fixed effects, county time-trends, demographic factors, the price of substitutes (e.g. home food prices), variation in state sales taxes rate on food, variation in food stamp rates, as well as changes in establishments and employment in related industries (e.g. grocery, traditional restaurants, etc.). These estimates are also robust across both contemporaneous and lagged analyses and among a subgroup of lower income individuals. Furthermore, these results are consistent with recent work by Anderson and Matsa (2011), who find that banning restaurants entirely would have an economically insignificant effect on BMI.

Recent legislative activity indicates that more debate on the taxation of fast food as a means of combating obesity is expected. Unfortunately, our results suggest that such taxes may do little to fight obesity in the United States. The reason for the lack of a measurable price effect is unclear. It is possible that fast food prices are so low that it would require a large tax to detect a change in behavior, or that substitutes (e.g. groceries) are not readily available in many areas. Alternatively, it is possible that changes in fast food prices are met with the expected substitution to other products, but these substitutes have equal caloric content or otherwise positively impact weight. When observing changes in fast food prices, no corresponding statistically significant change in BMI or obesity can be observed. It seems that the full set of issues accompanying obesity, specifically nutritional education and the availability of substitutes, likely need to be addressed in concert with fast food taxation before it can be effective in combating obesity.

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Appendix A. Calculation of county-level food prices

As discussed earlier, the ACCRA price data cover approximately 480 Core Based Statistical Areas (CBSAs). A CBSA is simply a general term that refers to either a Metropolitan or Micropolitan Statistical Area as defined by the U.S. Office of Management and Budget. These refer to large (greater than 50,000) or small (between 10,000 and 50,000) core urban areas, respectively. They include all of the counties in the core area itself in addition to any adjacent counties considered be highly integrated with the core, both socially and economically.[19]

To calculate county-level ACCRA prices we obtained the November 2007 Core Based Statistical Areas (CBSAs) and Combined Statistical Areas (CSAs) file from the U.S. Census Bureau, Population Division, which includes a crosswalk between CBSAs and county FIPS codes. For some CBSAs ACCRA recorded multiple sets of prices. For example, ACCRA collected separate sets of prices for Appleton and New London, Wisconsin, both of which are in Outagamie County and assigned to the Appleton, Wisconsin Metropolitan Statistical Area. Since the available crosswalk simply assigns each county to a CBSA, we first calculated the unweighted average set of ACCRA prices for each year, quarter, and CBSA. Next, we assigned the corresponding average CBSA price set to each county using the crosswalk. Our calculated prices are therefore averages of available prices calculated for each CBSA, and each county within a given CBSA is assigned the same price set.

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