Food Expenditures away from Home by Elderly Households

Steven T. Yen
Panagiotis Kasteridis
John B. Riley

Department of Agricultural and Resource Economics
The University of Tennessee
Knoxville, TN 37996-4518
http://economics.ag.utk.edu/

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Abstract

This study investigates the differentiated effects of economic and socio-demographic variables on food away from home (FAFH) expenditures by type of facility among elderly households in the United States. Using data from the 2008–2010 Consumer Expenditure Surveys, the systems of expenditures on full-service, fast food, and other restaurants are estimated with a multivariate sample selection estimator which also accommodates heteroscedasticity in the error distribution. Statistical significance of error correlations among equations justifies estimation of the sample selection systems. Income, employment statuses, race, education, geographic region, and household composition are important determinants of FAFH expenditures. Income contributes to full-service and fast-food expenditures by the elderly implying that the future of FAFH industry is tied to macroeconomic conditions. Better education is associated with greater probabilities and larger levels of expenditures at all facilities. Effects of the Supplemental Nutrition Assistance Program (SNAP) are found to be strong and negative, invalidating policy concerns for the general population that participation in the program might enhance consumption of less healthy FAFH.

Keywords Censoring · equivalence scale · elderly · food away from home · sample selection system

JEL D12, Q13, C31

Introduction

Americans spend nearly half of their food dollars on food away from home (FAFH). Total spending on FAFH by all families and individuals was $433.5 billion in 2010 and constituted 41.3 percent of total food expenditure, up from 32.0 percent in 1980 (USDA-ERS 2011). A variety of foodservice firms, including full-service restaurants, fast-food restaurants, drinking places, and retail stores, compete in the food industry. Full-service restaurants and fast-food companies, with total sales of $219.5 billion and $209.5 billion, captured 72.2% of the FAFH market in 2010 (USDA-ERS 2011). By 2009, there were over 221,513 full-service establishments and 268,459 fast-food establishments in the United States (U.S.) (U.S. Census Bureau, No date).

Owing to rapid income growth, urbanization, and globalization, FAFH consumption have also increased in many other countries such as Spain (Mutlu and Gracia 2006; Angulo et al. 2007), Greece (Mihalopoulos and Demoussis 2001), China (2006; Bai et al. 2010), Malaysia (Tan, 2010), Taiwan (Chang and Yen 2010), and many other Asian countries (Pingali 2007).

The literature has identified a list of economic and socio-demographic factors that contribute to FAFH consumption including income, household size and structure, regions, employment, age, education, race, and ethnicity (McCracken and Brandt 1987; Yen 1993; Byrne et al. 1996; Jensen and Yen 1996). Food assistance programs such as the Supplemental Nutrition Assistance Program (SNAP), have also had an impact on FAFH (Liu et al. 2012).

During the 2007-09 Great Recession, encompassing falling incomes, mounting unemployment, high food prices, and high participation rates in the federal food assistance programs, Americans ate out less due to tightened budget. Several of the demographic factors triggering FAFH have also changed remarkably in the 2000s (Cherlin 2010). For instance, increasing proportions of single parents, Hispanic and Asian immigrants, and the elderly in the population marked significant changes in family and ethnic composition. Continuation of rapid demographic changes may result in a new upward trend in FAFH consumption in the next few years. This possibility escalates the concerns about the public health implications of dining out. There is a large body of empirical literature pointing out that FAFH is less healthy than food at home (Mancino et al. 2009), and a number of policies have been implemented in order to raise the public awareness of the benefits of a healthy diet. Therefore, there exists a great interest in consumers' expenditures on FAFH among policy makers, the academia, and food marketers.

Type of facility has been the subject of investigation in consumer food purchases. Chung and Myers (1999) find store type more important in driving food price disparities than the geographical location of a household; such food price disparities can in turn affect consumer food purchases. On FAFH expenditures, previous studies have addressed differences between full-service and fast-food restaurants and found that the effects of economic and socio-demographic factors on FAFH generally differ by type of facility (McCracken and Brandt 1987; Byrne et al. 1996; Stewart and Yen 2004; Binkley 2006; Liu et al. 2012).

Much of the empirical literature on FAFH is dated except Liu et al. (2012). Further, this literature addresses the general population in a country. No study, to our knowledge, has investigated FAFH expenditures among the elderly. The elderly population may well have a distinct expenditure pattern on FAFH due to differences in lifestyle as well as economic conditions such as greater financial stability relative to younger households. This study fills this empirical void by focusing on FAFH by type of facility and for the elderly population in the U.S. To that end, FAFH expenditures for the younger population will also be investigated for comparison. Further, as in other microdata, the data used in this study contain censoring (zero observations) in the expenditure variables. Statistical procedures not accounting censoring produce inconsistent empirical estimates. To address this data feature, the sample selection system, an extension of the bivariate sample selection model (Heckman 1979) advanced by Yen (2005) and used in previous studies of FAFH (Stewart and Yen 2004; Liu et al. 2012), is further extended to accommodate heteroscedasticity in the error terms. Heteroscedasticity, if not accounted for, can cause statistical inconsistency in empirical models with limited dependent variables.

Theoretical framework

We motivate our empirical specification by the discrete random utility theory. A household maximizes the random utility function subject to a full-income constraint:

$$\max_{q,c,\ell}\{U(Dq,c,\ell;s)\mid p'q+c=w(T-\ell)+m\}, \tag{1}$$

where $q=[q_1,...,q_n]'$ is vector of quantities with positive prices $p=[p_1,...,p_n]'$, c is a composite commodity for other goods with price normalized at unity, s is a vector of demographic characteristics, m is non-wage earning; and for working members such as household head and spouse, ℓ is a vector of leisure activities with prices (wage rates) w, T is a vector of time endowments. $D=\text{diag}(d_1,...,d_n)$ is a diagonal matrix with each binary indicator d_i indicating a potential consumer of q_i. Assume the utility function $U(Dq,c,\ell;s)$ is strictly quasi-concave and increasing with respect to c, ℓ, and positive elements of Dq. Then, solving (1) yields the notional demand q^* for FAFH—a vector of optimal quantities demanded without non-negativity constraints, as function of prices, wage rates, and non-wage earning (Becker 1965). This constrained utility maximum framework motivates two alternative specifications for the demand functions. First, assume all individuals are potential consumers of q_i in which case $d_i=1$ for all i and censoring of each q_i corresponds to a corner solution governed by a Tobit mechanism.

Second, when an individual can otherwise be a potential non-consumer, either $d_i = 1$ and utility maximum occurs in the interior of the choice set (*viz*., $q_i > 0$), or $d_i = 0$ and $q_i = 0$ since price $p_i > 0$ by assumption. In this latter case, censoring in each q_i is governed by a sample selection mechanism. Express the notional demands as a system of equations for latent expenditures (y_i^*)

$$y_i^* = x'\beta_i + v_i,\ i = 1,\ldots,n, \tag{2}$$

where *x* is the vector of explanatory variables, β_i are parameter vectors, and v_i are random disturbances which reflect the unobservable (discussed below).

Data, samples and variables

Data are drawn from the 2008, 2009, and 2010 Consumer Expenditure Surveys (U.S. BLS 2009, 2010, 2011). Each survey provides consecutive two-week information on FAFH expenditures that are categorized by type of facility, and on economic and socio-demographic characteristics of the households. The final sample consists of 20,523 observations. We focus on FAFH expenditure by elderly households. For comparison, expenditures by younger households are also estimated. The sample is segmented into an elderly sample with household heads age ≥ 55 (n = 7,860 households) and a younger sample (n = 12,663). To accommodate economy of scale in consumption and the role of household composition, expenditures are expressed per equivalence scale according to the U.S. National Research Council definition (Citro and Michael 1995). By that definition, household equivalence scale is calculated as (number of adults + 0.7 × number of children)$^{0.65}$. The resulting equivalence scale is used to deflate both FAFH expenditures and pre-tax non-wage earning (henceforth, income).

Sample statistics of the expenditures are presented in the appendix (Table A1). The elderly spend more ($28.88) at full-service restaurants than younger households ($27.54), whereas the younger households spend more at fast-food restaurants ($25.50) than the elderly households ($16.08) per equivalence scale per two weeks. The elderly spend nearly twice as much at full-service restaurants as at fast-food restaurants, whereas the younger households spend about as much at the two types of facilities. Expenditures at other facilities represent a very small proportion of the FAFH budget—$1.50 among the elderly and $4.42 among the younger households. In terms of percentages, 54.61% of the elderly households consume at full-service restaurants, 63.22% at fast food restaurant, and 19.73% at other facilities during the two-week period; the percentages are higher among the younger households: at 57.04%, 76.93%, and 38.51%, respectively.

Drawing on the random utility theory above, non-wage earning (income) is used as the explanatory variable. Income is expected to increase FAFH consumption for both household groups. Demographic variables include age, employment statuses of household heads and spouses (if present), household composition, and dummy variables indicating geographical regions, seasons, home ownership, SNAP participation status, race, and education (Table A1).

Many state agencies of SNAP provide nutrition education to assist recipients in making healthy food and active lifestyles choices. SNAP participation may thus decrease consumption of FAFH as the participants opt for healthier diets. However, the program benefits frees up resources for FAFH as well as other goods. The net effect of SNAP on FAFH is thus unclear. We include a dummy variable to indicate SNAP participation status.

We draw on previous studies in the selection of additional explanatory variables (McCracken and Brandt 1987; Soberon-Ferrer and Dardis 1991; Yen 1993; Jensen and Yen 1996; Stewart and Yen 2004; Liu et al. 2012). Household composition variables are among the most commonly used in analysis of consumer demand, FAFH, and time-saving goods and services.

Homeowners may consume more food away from home because of greater financial stability on the one hand. On the other hand, they may have more or less cash flow depending on mortgage payments (versus rent) which can affect FAFH expenditure. The effects of home ownership on FAFH expenditures are therefore unclear. Tastes and eating habits may differ by race. Because food preferences and other unobserved characteristics may differ across geographic regions and seasons, dummy variables are also included to account for these differences. Due to the absence of prices in the single cross section used, these regional and seasonal dummy variables can accommodate regional and seasonal price variations as well, ameliorating biases due to omission of prices. Education, gender, and age are expected to influence FAFH expenditures to different extents by facility.

Because work competes with home production (meal-preparation) for the time endowment, and because the trade-off between time use and diet quality can play an important role in food consumption decisions (Becker 1965), households with working members are hypothesized to consume FAFH more often. Dummy variables indicating full-time and part-time employment by the household head and spouse are included in the selection equations as part of the parameter identification strategy (discussed below).[1] Finally, dummy variables are included to indicate the years 2009 and 2010. Definitions and descriptive statistics of all explanatory variables are provided in the appendix for both the elderly and younger households (Table A1). There are notable differences in the sample statistics between the two types of households. For instance, the

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[1] The theory of conditional demand suggests use of work hours, not wage rates, as the explanatory variable (Shaw and Feather 1999). Because our elderly sample contains a large proportion of non-working household heads (42.8%) and spouses (75.9%), dummy variables are used instead of work hours.

elderly households register an average income as high as \$17,540 per equivalent scale per year, compared to only \$4,030 for the younger households. Household age compositions are also very different: number of members age < 18 is only 0.10 among the elderly households, compared to 0.94 for the younger households.

Econometric procedure

Censoring in our expenditure variables need to be addressed to obtain consistent parameter estimates. Earlier studies of FAFH employed models such as the Tobit model (McCracken and Brandt 1987), single-hurdle model (Yen 1993), and double-hurdle model (Jensen and Yen 1996; Mutlu and Gracia 2006) to accommodate the censored data. The sample selection system estimator has also been used in Stewart and Yen (2004) and, more recently, Liu et al. (2012). Except Jensen and Yen (1996), these previous studies have not addressed heteroscedasticity in the error terms, which can cause statistical inconsistency in the empirical estimates. We extend the sample selection system procedure of Stewart and Yen (2004) and Liu et al. (2012), suggested in Yen (2005), by accommodating heteroscedasticity of the error terms.

Consider a three-good system where outcome in each FAFH expenditure (y_i) is governed by a binary sample-selection rule

$$\begin{aligned} \log y_i &= x'\beta_i + v_i && \text{if } z'\alpha_i + u_i > 0 \\ y_i &= 0 && \text{if } z'\alpha_i + u_i \leq 0, \end{aligned} \tag{3}$$

where i = 1,2,3 for full service, fast food, and other facilities, respectively, z and x are vectors of explanatory variables, α_i and β_i are conformable parameter vectors, and u_i and v_i are random errors. Assume the concatenated error vector $[u', v']' = [u_1, u_2, u_3, v_1, v_2, v_3]'$ is distributed as six-dimensional normal with zero means, standard deviations $[1,1,1,\sigma_1,\sigma_2,\sigma_3]'$, and correlation matrix

$R = [\rho_{ij}]$ such that correlation between the error terms in the corresponding selection and level equations, (u_i, v_i), is $\rho_{i+3,i}$. The sample selection system (3) is more flexible than the Tobit system in that the binary and level outcomes are governed by separate stochastic processes. Comparison of the sample selection system and Tobit system is addressed in the empirical section below.

Each dependent variable y_i in (3) is transformed by natural logarithm. Such transformation is common in estimation of endogenous selection and switching regression model and ameliorates potential nonnormality and heteroscedasticity of the error term (Yen 2005; Yen and Rosiński 2008). Importantly, heteroscedasticity in the error terms can cause inconsistency in parameter estimates in limited dependent variable models. To accommodate heteroscedasticity in the error terms, each of the standard deviations is parameterized as a function of explanatory variables *h* with parameter vector γ_i:

$$\sigma_i = \exp(h'\gamma_i),\ i = 1,2,3. \tag{4}$$

The sample likelihood function is identical to that presented in Yen (2005), except the additional parameterization in (4). Denote the *k*-dimensional probability density function (pdf) as ϕ_k, cumulative distribution function (cdf) as Φ_k, and define standard normal variates $t_i = (\log y_i - x'\beta_i)/\sigma_i$ for i = 1,2,3. Then, for an all-positive regime, the likelihood contribution is

$$L = (\prod_{i=1}^{3} y_i^{-1}\sigma_i^{-1})\,\phi_3(t_1,t_2,t_3;R_3)\int_{-z'\alpha_1}^{\infty}\int_{-z'\alpha_2}^{\infty}\int_{-z'\alpha_3}^{\infty} h(u_1,u_2,u_3 \mid v_1,v_2,v_3)\,du_3du_2du_1, \tag{5}$$

where R_3 is correlation matrix among error terms (v_1, v_2, v_3) which is the 3 × 3 lower-right sub-matrix of *R*, and $(\Pi_{i=1}^{3} y_i^{-1})$ is the Jacobian of the transformations from $[v_1, v_2, v_3]'$ to $[t_1, t_2, t_3]'$. In (5), the conditional density $h(u_1,u_2,u_3 \mid v_1,v_2,v_3)$ is trivariate normal (Kotz et al. 2000, pp. 111–

112) which can be integrated by usual means (Yen 2005).

Consider next a partially censored regime, in which y_1 is censored at zero and (y_2, y_3) are positive. The likelihood contribution is

$$L = (\prod_{i=2}^{3} y_i^{-1}\sigma_i^{-1})\,\phi_2(t_2, t_3; R_2) \int_{-\infty}^{-z'\alpha_1} \int_{-z'\alpha_2}^{\infty} \int_{-z'\alpha_3}^{\infty} h(u_1, u_2, u_3 \mid v_2, v_3)\, du_3 du_2 du_1, \qquad (6)$$

where R_2 is correlation matrix between error terms (v_2, v_3) which is the 2 × 2 lower-right sub-matrix of R. Finally, for a sample regime with (y_1, y_2) and $y_3 > 0$, the likelihood contribution is

$$L = y_3^{-1}\sigma_3^{-1}\,\phi_3(t_1) \int_{-\infty}^{-z'\alpha_1} \int_{-\infty}^{-z'\alpha_2} \int_{-z'\alpha_3}^{\infty} h(u_1, u_2, u_3 \mid v_3)\, du_3 du_2 du_1. \qquad (7)$$

The integrals in (6) and (7) can be simplified following a procedure similar to that in (5) because the conditional densities $h(u_1, u_2, u_3 \mid v_2, v_3)$ and $h(u_1, u_2, u_3 \mid v_3)$ are also trivariate normal. The likelihood contributions for other partially censored regimes are covered in (6) and (7) by permuting the error terms. It is clear from (5), (6), and (7) that estimation of the sample selection system requires evaluations of trivariate normal cumulative distribution functions (cdf's) for all sample observations. These probability integrals are evaluated with Gaussian quadratures. The parameters to estimate include those from the selection equations (α_i), level equations (β_i), heteroscedasticity specification (γ_i) for i = 1,2,3, and error correlations (ρ_{ij}) for $i > j$.

The sample selection system nests two restricted specifications: (i) an independent model which corresponds to parametric restrictions $\rho_{ij} = 0$ for all $i \neq j$, viz., with all error correlations equal to zeros; (ii) a pairwise selection system which corresponds to $\rho_{ij} = 0$ for all $i \neq j$ except $\rho_{41} \neq 0$, $\rho_{52} \neq 0$, and $\rho_{63} \neq 0$. These restricted models can be estimated by imposing the above parametric restrictions. In addition, the independent model for each of the expenditures in the system, can be estimated by the probit model based on the binary (0/1) outcome related to y_i using the whole sample, and OLS for each $\log(y_i)$ using the truncated sample (conditional on $y_i > 0$). The pairwise selection system consists of three bivariate sample selection models (Heckman 1979) for the three expenditures which can be estimated as such separately. Tests of the sample selection system against the two nested models can be done with the Wald, Lagrange multiplier (LM), or likelihood-ratio (LR) test (Engle 1984).

Marginal effects of probabilities, conditional levels, and unconditional levels are calculated to explore the effects of explanatory variables. Specifically, for each good *i*, the probability, conditional mean and unconditional mean of y_i are (Yen and Rosiński 2008, p. 5)

$$\Pr(y_i > 0) = \Phi_1(z'\alpha_i), \tag{8}$$

$$E(y_i \mid y_i > 0) = \exp(x'\beta_i + \sigma_i^2/2)\Phi_1(z'\alpha_i + \rho_{i+3,i}\sigma_i)/\Phi_1(z'\alpha_i), \tag{9}$$

$$E(y_i) = \exp(x'\beta_i + \sigma_i^2/2)\Phi_1(z'\alpha_i + \rho_{i+3,i}\sigma_i). \tag{10}$$

Differentiating (differencing) (8), (9), and (10) with respect to variables *x*, *z*, and *h* gives the marginal effects of continuous (discrete) explanatory variables, which are extensions of those in Yen and Rosiński (2008, p. 5) due to the heteroscedasticity specification (4); analytic derivatives are available upon request. We evaluate all marginal/discrete effects for all sample observations and average them over the sample. Standard errors of these marginal effects are calculated by a mathematical approximation procedure known as the delta method (Spanos 1999, p. 493).

Specification tests and estimation results

Choice of variables (*h*) in the heteroscedasticity specification (4) is carried out empirically. Income is found to perform better, in terms of statistical significance, than the other variables and is retained. Next, we test for equality of parameters between the elderly and younger households.

Using a likelihood ratio (LR) test, similar to Chow test in linear regression models, the hypothesis of equal parameters is rejected (LR = 4652.90, df = 165, p-value < 0.0001), which justifies segmentation of the sample and our focus on the elderly households. The parameter estimates and, more importantly marginal effects (reported below) are found to differ notably between the elderly and younger households; these differences would have been masked by the use of a pooled sample and further justify separate analysis for the two groups of households.

We then carry out a series of specification tests, focusing on the elderly households. First, heteroscedasticity of error terms in the level equations is tested. Results of t-tests indicate statistical significance of income in the heteroscedasticity equations for full-service and other facilities, both at the 5% level of significance. Further, Wald test result suggests joint significance of income in the heteroscedasticity equations (Wald = 9.17, df = 3, p-value = 0.027), supporting the heteroscedasticity specification. The heteroscedastic sample selection system and is then compared to the heteroscedastic trivariate Tobit system, based on log-likelihood contributions at the maximum-likelihood estimates for both models. Result of Vuong's (1989) test, with standard normal statistic z = 37.60, suggests that the sample selection system performs better than the Tobit alternative in fitting the data. The sample selection system is then tested against the nested independent and pairwise selection (Heckman 1979) systems discussed above. Of the 15 error correlation coefficients, 14 are significant at the 1% level of significance and one at the 5% level by individual t-tests (Table A2). As expected, we reject both the pairwise selection system (Wald = 1926.39, LR = 1876.64, LM = 1618.81, df = 12) and independent system (Wald = 11861.66, LR = 2132.49, LM = 1851.58, df = 15) with p-values < 0.0001. In sum, results suggest separate analysis with the elderly sample, the sample selection system is preferred to the Tobit alternative, and the heteroscedastic specification is justified. Maximum-likelihood estimates for the elderly sample are presented in the appendix (Table A2), and estimates for younger households are available upon request.

Marginal effects

Marginal effects of variables on the probability, conditional level, and unconditional level for the elderly households are presented in Table 2. These effects differ, notably for many variables, between the elderly households and younger households. For instance, whereas income contributes to full-service and fast-food expenditures by both groups, the elderly are more responsive to income changes than their younger counterparts, in terms of probability and levels of expenditures at full-service restaurants. A different pattern is found for fast foods, with a greater effect on probability but a smaller effects on conditional level for the elderly households than the younger households. Income affects expenditure at other facilities among the elderly households but does not the younger households. The effects of employment statuses (full-time and part-time) are all positive for both household groups, although the effects are generally larger for the younger households than the elderly households. Differences are also found in many other variables. Marginal effects for younger households are available upon request.

Focusing on the elderly households (Table 2), the marginal effects for most variables differ substantially, in signs and magnitudes, across facilities. Income has positive effects on expenditures at full-service and fast-food restaurants but negative effects on expenditures at other facilities. The magnitudes differ among facilities. Overall, the marginal effects of income on FAFH are relatively small, with a $10,000 increase in income per equivalence scale increasing expenditure at full-service (fast-food) restaurants by less than $3 ($1) per equivalence scale for two weeks. These positive but small effects of income are in agreement with findings from previous studies for the general population in the U.S. (e.g., McCracken and Brandt 1987; Jensen and Yen

1996; Liu et al. 2012). Larger effects of income are reported for China (Bai et al. 2010) and Spain (Angulo et al. 2007), also on the general population. Income has negative but barely noticeable effects on the levels of expenditures at other facilities.

The roles of household composition are mixed. Numbers of adults age 18–64 and age > 64 have positive and significant effects on the probabilities of consuming at all three types of facilities but, conditional on consumption, the levels of expenditures are significant, negative and large. An additional member age 18–64 (age > 64) decreases expenditure at full-service restaurants by $7.02 ($4.75) per equivalent scale per two weeks. These conflicting roles of variables on probabilities and levels of expenditures can be masked by the use of the more restrictive Tobit system and highlight an important advantage of the sample selection parameterization. Not surprisingly, number of younger household members (age < 18) has significant and positive effects on probability (3.72%), conditional level ($0.97), and unconditional level ($0.49) of expenditure at other facilities.

Age has negative and significant effects on probabilities and levels of expenditures at both full-service and fast-food restaurants, while its negative effect on other facilities is seen only through probability. The negative effect of age on probability of fast food is particularly notable, at –6.05%. Thus, expenditures at all types of facilities decrease with age, in terms of probabilities and, in the case of full-service and fast-food, by way of expenditure levels as well.

Despite the small proportions of working household heads (57.3%) and spouses (24.1%), employment statuses have by far the most notable effects on FAFH expenditures. Supporting our hypothesis on the role of work in FAFH, full-time and part-time employment by both household head and the spouse have positive, significant, and relatively large impacts on FAFH. Full-time work by the household head (spouse), for instance, increases the probability of consuming at full-service restaurants by 9.12% (9.72%), conditional level by $9.08 ($9.88), and unconditional level by $9.95 ($10.97) per equivalence scale per two weeks. The effects of part-time work by both household head and spouse are also positive and very notable in magnitudes. Positive effects of employment on FAFH are found in previous studies, despite the different methodology and variables are used. Stewart and Yen (2004), also using the sample selection system approach, find positive impact of work hour by household managers on the probabilities of consuming FAFH. Yen (1993), based on much earlier Consumer Expenditure Survey data, find that wife's work hour has a positive effect on lunch and dinner away from home. More recently, Liu et al. (2012) find employment has a positive effect on fast food and full-service expenditures for husband-wife households with and without children. Positive effects of employment statuses have also been reported for other countries, such as China (Bai et al. 2010), Spain (Mutlu and Gracia 2006), and Taiwan (Chang and Yen 2010).

Race plays a role in FAFH expenditures. Compared to Blacks, White households are 14.26% more likely to consume at full-service restaurants and spend $18.77 ($15.91) more per equivalent scale per two weeks, conditional (unconditional) on consumption. Households of other races are 9.12% more likely to consume and also consume $37.89 ($29.49) more per equivalent scale per two weeks at full-service restaurants, conditional (unconditional) on consumption; these households also consume more at fast-food and other facilities.

Education has positive effects on FAFH. Compared to households headed by an individual with high school education, households less than high school education are less likely to consume and also consume less at full-service restaurants, while households with college and post-graduate education are more likely to consume and also consume more at full-service restaurants. Households with post-graduate education are 10.11% more likely to consume at full-service restaurants and spend $17.55 ($16.21) more per equivalence scale per two weeks than households with only high-school education. College and post-graduate education also increase expenditures at fast-food and at other facilities.

Unlike studies for the general population in which the roles of SNAP are less clear (Liu et al. 2012), we find consistent effects of SNAP for the elderly. Participation in SNAP decreases the probabilities (except fast food), conditional levels, and unconditional levels of expenditures at all three facilities. The effects on full service are particularly notable, with SNAP participation decreasing the probability of consuming at full-service restaurants by 5.83%, conditional level by $19.33, and unconditional level by $13.37 per equivalence scale per two weeks. Fast food is a substitute for SNAP-participating elderly in terms of probability (positive), although its effects on conditional level ($10.48) and unconditional level ($7.43) are also both negative.

Regional differences are also present. Compared to households in the West, households from the Northeast are 2.51% more likely to consume at other facilities; they also spend $2.55 ($1.82) more on fast food, conditional (unconditional) on consumption. Households in the Midwest spend less at full-service restaurants; they are also more likely to spend at other facilities. Finally, seasonal effects are scant and barely noticeable, suggesting more consumption of fast food during summer months and more at other facilities during the fall.

Concluding remarks

Much of the empirical literature addresses FAFH by the general population in the U.S. and other countries. In this study, we investigate FAFH by elderly households in the U.S., focusing on expenditures by type of facilities. We find differentiated effects of income, education, employment statuses, and other socio-demographic characteristics on FAFH expenditures across facilities by the elderly households. These effects differ from those among younger households. The elderly households do have different consuming patterns. Although not directly comparable due to the use of different econometric approaches, measures, and variables, our results generally echo findings for other countries that economic and socio-demographic characteristics play key roles in FAFH expenditures.

The findings of this study can inform policy deliberations by federal and state governments. Our estimates identify segments of the elderly population that should be targeted for nutrition education. The educated spend more on FAFH than the less educated, as do the employed than the unemployed/retired. Therefore, nutrition educators could focus their efforts on employed elderly with busy working schedules. Such efforts include warnings about the relatively higher levels of sodium, cholesterol, and saturated fats in FAFH meals, recommendations about healthy FAFH choices such as fruits, vegetables, milk, and oils, and educational messages about moderating consumption of fats, added sugars, and alcohol. Our estimate of the effect of SNAP participation on FAFH is also informative for policymakers whose goal is more nutritious and adequate food. Our finding suggests that the concern about SNAP promoting consumption of less healthy FAFH is groundless, for the elderly population, since we find the opposite.

The results of the study can also assist marketing strategies by foodservice firms. The strong impact of household composition changes on FAFH spending provides valuable information for the foodservice industry. For instance, we find household composition variables generally have negative effects on FAFH. Thus, promotional campaigns offering quantity discounts can be an effective tool for restaurants to attract large families. Senior citizen discounts could also be used as age reduces the probability and levels of fast food. Such discounts are currently available in many restaurants in the U.S.

Some conclusions about future trends of FAFH consumption can also be drawn. Income contributes to expenditures at both full-service and fast-food facilities. This suggests the restaurant industry will suffer in the case of a recession but will prosper when the economy recovers.

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Table 1 Sample statistics of FAFH expenditures per equivalent scale by facility

Expenditure	%	Full Sample		Consuming Sample	
($ / 2 weeks)	Consuming	Mean	SD	Mean	SD
Elderly households (*n* = 7,860)					
Full-service restaurants	54.61	28.88	53.86	52.89	63.58
Fast-food restaurants	63.22	16.08	25.75	25.44	28.48
Other facilities	19.73	1.50	5.85	7.62	11.27
Younger (*n* = 12,663)					
Full-service restaurants	57.04	27.54	52.31	48.27	61.61
Fast-food restaurants	76.93	25.50	31.14	33.15	31.73
Other facilities	38.51	4.42	13.80	11.48	13.80

Table 2 Marginal effects of explanatory variables on probabilities, conditional levels, and unconditional levels among elderly households

	Full-service Restaurants			Fast-food Restaurants			Other Facilities		
Variable	Probability	Level (C)	Level (U)	Probability	Level (C)	Level (U)	Probability	Level (C)	Level (U)
Continuous explanatory variables									
Income / 10	2.923***	2.722***	2.960***	1.564***	0.553**	0.740***	0.137	–0.422***	–0.081**
	(0.263)	(0.354)	(0.262)	(0.265)	(0.223)	(0.165)	(0.231)	(0.138)	(0.033)
Members < 18	–0.923	–6.047***	–3.968***	3.331**	0.326	1.024	3.721***	0.968**	0.485***
	(1.293)	(2.080)	(1.327)	(1.393)	(0.864)	(0.698)	(0.969)	(0.448)	(0.127)
Members 18–64	3.229***	–7.026***	–2.559***	4.676***	–1.039*	0.464	2.743***	–1.278***	–0.059
	(0.808)	(1.164)	(0.814)	(0.816)	(0.557)	(0.442)	(0.656)	(0.342)	(0.085)
Members > 64	5.999***	–4.745***	0.092	5.441***	–2.522***	–0.315	1.727*	–2.023***	–0.292**
	(1.029)	(1.471)	(1.012)	(1.015)	(0.731)	(0.551)	(0.894)	(0.527)	(0.126)
Age / 10	–2.736***	–1.908	–2.417***	–6.051***	–2.496***	–3.100***	–2.096***	–0.076	–0.175
	(0.821)	(1.220)	(0.829)	(0.790)	(0.599)	(0.442)	(0.763)	(0.453)	(0.110)
Binary explanatory variables									
Urban	0.107	11.276***	6.617***	0.167	2.815**	1.869*	–1.550	0.798	0.061
	(2.160)	(2.551)	(1.769)	(2.129)	(1.434)	(1.059)	(1.850)	(0.837)	(0.229)
Homeowner	8.471***	4.812**	6.621***	5.058***	–0.524	0.913	0.721	–1.767**	–0.310
	(1.440)	(2.015)	(1.262)	(1.413)	(1.033)	(0.741)	(1.231)	(0.791)	(0.193)
SNAP	–13.586***	–21.121***	–16.194***	–1.603	–6.809***	–4.721***	–1.183	–2.825***	–0.647***
	(2.739)	(2.747)	(1.501)	(2.590)	(1.406)	(1.025)	(2.324)	(0.954)	(0.226)
White	14.258***	18.770***	15.912***	1.931	0.587	0.846	1.677	1.595**	0.445**
	(1.842)	(2.078)	(1.248)	(1.756)	(1.210)	(0.886)	(1.496)	(0.800)	(0.192)
Other race	9.117***	37.889***	29.485***	1.980	6.823***	5.039*	5.214*	6.572***	2.082***
	(3.048)	(8.107)	(5.835)	(3.051)	(2.592)	(2.032)	(2.839)	(2.501)	(0.753)
< High school	–8.736***	–3.286	–5.913***	–3.941**	1.107	–0.273	–1.315	0.632	0.026
	(1.716)	(2.506)	(1.530)	(1.634)	(1.231)	(0.878)	(1.421)	(0.909)	(0.217)
College	5.472***	9.325***	8.089***	1.923	2.109**	1.846***	2.248**	0.514	0.280*
	(1.289)	(1.964)	(1.355)	(1.275)	(0.925)	(0.699)	(1.095)	(0.593)	(0.153)
Grad school	10.105***	17.553***	16.214***	2.278	2.270*	2.068**	4.342***	1.400	0.665**
	(1.776)	(3.156)	(2.351)	(1.813)	(1.336)	(1.037)	(1.615)	(0.879)	(0.259)
Northeast	0.443	3.146	2.053	0.610	2.549**	1.817*	2.505*	0.531	0.309
	(1.648)	(2.463)	(1.697)	(1.661)	(1.241)	(0.947)	(1.483)	(0.827)	(0.224)
Midwest	–2.393	–7.843***	–5.624***	–1.255	–0.330	–0.519	3.780***	0.722	0.452**
	(1.619)	(2.140)	(1.459)	(1.620)	(1.136)	(0.854)	(1.442)	(0.795)	(0.217)
South	–1.841	–2.154	–2.123	0.049	–0.156	–0.090	2.650**	–0.235	0.150
	(1.512)	(2.184)	(1.478)	(1.510)	(1.066)	(0.813)	(1.301)	(0.690)	(0.181)
Spring	0.330	2.606	1.682	1.903	1.315	1.331*	1.191	0.791	0.261
	(1.496)	(2.241)	(1.527)	(1.449)	(1.058)	(0.805)	(1.260)	(0.721)	(0.187)
Summer	–0.007	–0.311	–0.185	1.000	2.335**	1.775**	0.460	0.144	0.065
	(1.513)	(2.146)	(1.479)	(1.469)	(1.121)	(0.839)	(1.283)	(0.704)	(0.180)
Fall	–0.323	–0.687	–0.554	1.415	0.729	0.824	2.807**	0.533	0.332*
	(1.493)	(2.144)	(1.461)	(1.475)	(1.051)	(0.799)	(1.292)	(0.682)	(0.185)
Year 2009	–0.526	–0.378	–0.471	–1.124	0.612	0.122	–1.515	0.528	–0.007
	(1.299)	(1.889)	(1.296)	(1.275)	(0.902)	(0.677)	(1.047)	(0.589)	(0.146)
Year 2010	–1.924	–2.736	–2.497**	–2.379*	0.574	–0.211	–3.299***	1.565**	0.059
	(1.303)	(1.860)	(1.261)	(1.291)	(0.931)	(0.692)	(1.052)	(0.648)	(0.155)
Work FT (H)	9.122***	9.079***	9.947***	10.828***	4.188***	5.555***	9.810***	1.547***	1.097***
	(1.215)	(1.288)	(1.400)	(1.234)	(0.531)	(0.673)	(1.282)	(0.523)	(0.176)
Work PT (H)	3.609***	3.568***	3.884**	6.176***	2.393***	3.156***	6.170***	0.985**	0.716***
	(1.367)	(1.383)	(1.519)	(1.379)	(0.566)	(0.742)	(1.413)	(0.389)	(0.189)
Work FT (S)	9.716***	9.878***	10.968***	5.428***	2.089***	2.757***	5.326***	0.850**	0.607***
	(1.429)	(1.577)	(1.771)	(1.527)	(0.615)	(0.812)	(1.330)	(0.341)	(0.167)
Work PT (S)	6.804***	6.903***	7.630***	6.125***	2.391***	3.166***	4.337**	0.694*	0.499**
	(1.906)	(2.041)	(2.304)	(1.926)	(0.794)	(1.058)	(1.833)	(0.368)	(0.228)

Asymptotic Standard errors in parentheses. Asterisks indicate levels of significance: *** = 1%, ** = 5%, * = 10%.

Appendix

Table A1 Definitions and sample statistics of explanatory variables

Variable	Definition	Elderly	Younger
Continuous explanatory variables			
Income	Pre-tax non-wage income per equivalent scale in past 12 months in $1,000	17.54 (21.47)	4.03 (11.01)
Members < 18	Number of children age < 18	0.10 (0.44)	0.94 (1.19)
Members 18–64	Number of adults age 18–64	0.98 (1.05)	1.90 (0.83)
Members > 64	Number of adults age > 64	0.81 (0.78)	0.03 (0.19)
Age	Age of household head	67.52 (9.53)	38.49 (9.82)
Binary explanatory variables (yes = 1, no = 0)			
Urban	Resided in urban area	0.93	0.95
Homeowner	Owned a home	0.80	0.59
SNAP	Any member received food stamps last year	0.05	0.09
White	Race is white	0.86	0.81
Other race	Race is of other race	0.04	0.07
Black	Race is black (reference)	0.10	0.12
< High school	Has less than high school	0.16	0.11
High school	High school graduate (reference)	0.29	0.25
College	Has a bachelor's or some college	0.42	0.53
Graduate	Has a graduate degree	0.13	0.11
Northeast	Resided in the Northeast	0.20	0.18
Midwest	Resided in the Midwest	0.24	0.26
South	Resided in the South	0.35	0.34
West	Resided in the West (reference)	0.21	0.22
Spring	Survey occurred during spring	0.26	0.27
Summer	Survey occurred during summer	0.24	0.24
Fall	Survey occurred during fall	0.26	0.25
Winter	Survey occurred during winter (reference)	0.24	0.24
Year 2008	Data came from year 2008 (reference)	0.32	0.33
Year 2009	Data came from year 2009	0.35	0.33
Year 2010	Data came from year 2010	0.33	0.34
Work FT (H)	Household head worked full-time	0.28	0.63
Work PT (H)	Household head worked part-time	0.15	0.19
Not working (H)	Household head not working (reference)	0.57	0.18
Work FT (S)	Spouse present × spouse worked full-time	0.27	0.35
Work PT (S)	Spouse present × spouse worked part-time	0.07	0.08
Not working (S)	Spouse absent or not working (reference)	0.66	0.57
Sample size		7,860	12,663

Standard deviations in parentheses. Household head is defined as the husband for a married household, and as the reference person for a single-person household. All households do not have a spouse present and, therefore, spouse' working statuses reflect their interactions with a "spouse" present dummy indicator.

Table A2 ML estimation of sample selection system with heteroscedastic errors: FAFH expenditures by elderly households

	Selection Equations			Level Equations		
Variable	Full-serv.	Fast-food	Other	Full-serv.	Fast-food	Other
Constant	–0.396**	0.844***	–0.860***	4.069***	3.388***	2.249***
	(0.183)	(0.183)	(0.226)	(0.229)	(0.203)	(0.500)
Income / 10	0.080***	0.044***	0.005	0.000	–0.001	–0.049***
	(0.007)	(0.008)	(0.009)	(0.008)	(0.008)	(0.016)
Members < 18	–0.025	0.094**	0.142***	–0.104**	–0.036	0.049
	(0.035)	(0.039)	(0.037)	(0.050)	(0.037)	(0.065)
Members 16–64	0.088***	0.132***	0.105***	–0.204***	–0.110***	–0.234***
	(0.022)	(0.023)	(0.025)	(0.027)	(0.024)	(0.052)
Members > 64	0.164***	0.153***	0.066*	–0.212***	–0.180***	–0.312***
	(0.028)	(0.029)	(0.034)	(0.035)	(0.032)	(0.072)
Age / 10	–0.075***	–0.170***	–0.080***	0.015	–0.009	0.036
	(0.023)	(0.023)	(0.029)	(0.029)	(0.027)	(0.069)
Urban	0.003	0.005	–0.058	0.250***	0.114*	0.147
	(0.059)	(0.060)	(0.068)	(0.073)	(0.069)	(0.131)
Homeowner	0.228***	0.140***	0.028	–0.066	–0.095**	–0.240**
	(0.039)	(0.039)	(0.048)	(0.052)	(0.045)	(0.099)
SNAP	–0.369***	–0.045	–0.046	–0.267**	–0.284***	–0.444**
	(0.075)	(0.072)	(0.092)	(0.110)	(0.081)	(0.201)
White	0.386***	0.054	0.065	0.155**	–0.005	0.195
	(0.051)	(0.049)	(0.059)	(0.068)	(0.055)	(0.130)
Other race	0.255***	0.056	0.187*	0.401***	0.211**	0.545***
	(0.088)	(0.087)	(0.096)	(0.114)	(0.090)	(0.196)
< High school	–0.235***	–0.109**	–0.051	0.104*	0.101*	0.113
	(0.046)	(0.045)	(0.056)	(0.063)	(0.053)	(0.121)
College	0.149***	0.054	0.085**	0.079*	0.054	0.020
	(0.035)	(0.036)	(0.041)	(0.044)	(0.040)	(0.084)
Grad school	0.279***	0.065	0.158***	0.120**	0.053	0.086
	(0.050)	(0.052)	(0.056)	(0.060)	(0.054)	(0.114)
Northeast	0.012	0.017	0.094*	0.053	0.089*	0.016
	(0.045)	(0.047)	(0.054)	(0.056)	(0.051)	(0.113)
Midwest	–0.065	–0.035	0.140***	–0.118**	0.006	0.015
	(0.044)	(0.045)	(0.052)	(0.055)	(0.050)	(0.109)
South	–0.050	0.001	0.100**	–0.008	–0.007	–0.090
	(0.041)	(0.043)	(0.049)	(0.052)	(0.046)	(0.100)
Spring	0.009	0.054	0.045	0.045	0.023	0.078
	(0.041)	(0.041)	(0.047)	(0.052)	(0.045)	(0.096)
Summer	0.000	0.028	0.018	–0.006	0.075	0.009

	(0.041)	(0.042)	(0.049)	(0.051)	(0.047)	(0.098)
Fall	–0.009	0.040	0.105**	–0.008	0.008	0.010
	(0.041)	(0.042)	(0.047)	(0.051)	(0.045)	(0.095)
Year 2009	–0.014	–0.032	–0.058	0.003	0.041	0.105
	(0.036)	(0.036)	(0.041)	(0.044)	(0.039)	(0.081)
Year 2010	–0.053	–0.067*	–0.128***	–0.018	0.058	0.280***
	(0.036)	(0.036)	(0.042)	(0.045)	(0.040)	(0.085)
Work FT (H)	0.250***	0.308***	0.350***			
	(0.034)	(0.036)	(0.043)			
Work PT (H)	0.099***	0.177***	0.222***			
	(0.038)	(0.041)	(0.048)			
Work FT (W)	0.267***	0.155***	0.193***			
	(0.040)	(0.044)	(0.046)			
Work PT (W)	0.188***	0.177***	0.157**			
	(0.053)	(0.057)	(0.063)			
Het. specification						
Constant				0.276***	0.181***	0.361***
				(0.019)	(0.019)	(0.069)
Income / 10				–0.010**	–0.002	–0.021**
				(0.004)	(0.005)	(0.011)
Error correlations (ρ_{ij})						
Fast-food (selection)	0.529***					
	(0.015)					
Other (selection)	0.389***	0.490***				
	(0.020)	(0.021)				
Full-serv. (level)	–0.806***	–0.245***	–0.201***			
	(0.016)	(0.019)	(0.019)			
Fast-food (level)	–0.107***	–0.699***	–0.082***	0.195***		
	(0.020)	(0.025)	(0.021)	(0.016)		
Other (level)	–0.121**	–0.227***	–0.510***	0.181***	0.128***	
	(0.054)	(0.066)	(0.131)	(0.030)	(0.028)	
Log likelihood	–58808.644					

Asterisks indicate level of significance: *** = 1%, ** = 5%, * = 10%.