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The Distinct Impact of Food Stamps on Food Spending

Parke Wilde and Christine Ranney

The Southworth hypothesis predicts that inframarginal food stamp recipients should choose the same bundle of goods, whether they receive coupons or cash. Empirical research has contradicted this prediction. Here, we present a model that retains some attractive features of the Southworth hypothesis, while relaxing the key assumption that appears to be incorrect. In particular, we allow different forms of benefits to have distinct effects on desired, or unrestricted, food spending. Two categories of previously commonly used empirical models are evaluated as special cases of our more general model. We estimate this model using data from two cash-out experiments.

Key words: food consumption, food stamps, nonlinear regression, program evaluation

Introduction

Economists have remained uncertain of how much more effective are food stamps than cash benefits at increasing food expenditures. This uncertainty persists because the most commonly used theoretical model, the Southworth hypothesis, has been contradicted by a large body of empirical evidence. Ironically, applied models can be misspecified if they assume that the Southworth hypothesis is correct, and they can be equally suspect if they ignore this hypothesis altogether.

The Southworth hypothesis centers on the distinction between two categories of food stamp coupon recipients: (*a*) unconstrained or inframarginal recipients, whose food expenditures exceed the value of their coupon benefits, and (*b*) constrained or extramarginal recipients, whose food expenditures are less than or equal to the value of their coupon benefits. The hypothesis maintains that inframarginal recipients should choose the same bundle of goods whether they receive food stamps or cash.

Twenty years of empirical research have contradicted this prediction. In light of this research, it makes sense to propose a model that distinguishes between inframarginal and extramarginal food stamp recipients, in a manner consistent with theory but without assuming that inframarginal recipients are unaffected by the form of their benefits.

We present and test such a model using data from food stamp cash-out experiments in San Diego and Alabama. An appealing feature of this model is that two important categories of models used previously can be treated as special cases: one assumes the Southworth hypothesis is correct, and the other ignores the distinction between inframarginal and extramarginal recipients.

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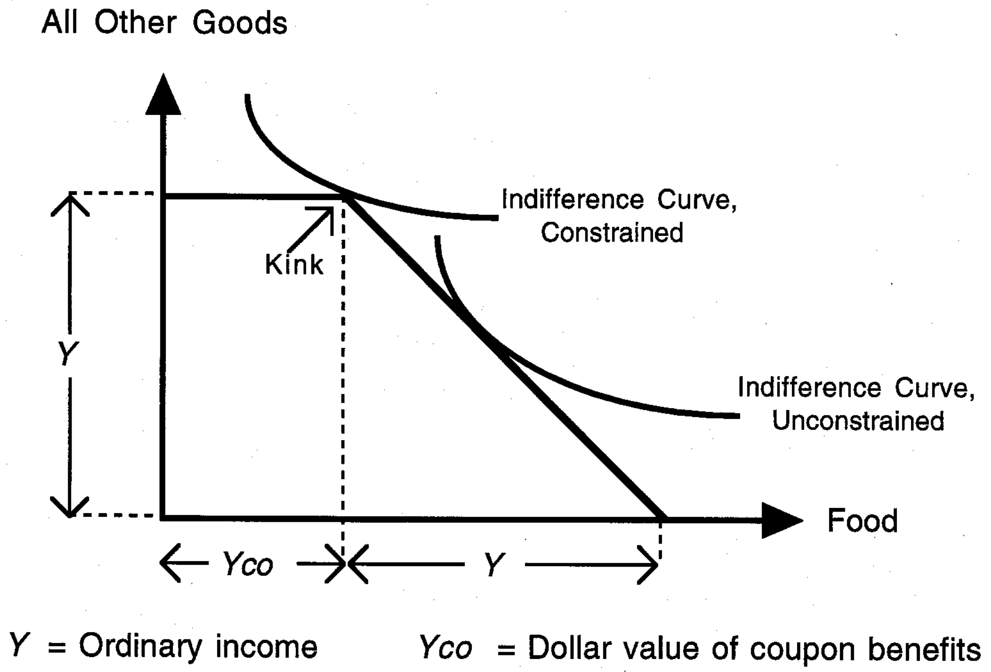

Figure 1. Choice under the Southworth hypothesis.

Note: The axes are scaled so that a unit of food or a unit of other goods each costs one dollar.

Theory and Literature Review

If food stamp coupons can only be used to purchase food, then a recipient family's budget constraint is piecewise linear or "kinked," when drawn in two dimensions (fig. 1). If utility is increasing in both arguments, then constrained families may consume less food if the Food Stamp Program is cashed out, but unconstrained families will be unaffected. Likewise, constrained families may prefer cash to coupons, but unconstrained families will be indifferent (Southworth; Mittelhammer and West).

These predictions are routinely assumed to be true in the current theoretical literature on cash and in-kind government benefits, even in articles that argue in favor of in-kind transfers (Gahvari). In current empirical work, these hypotheses also are frequently assumed to be true (see Murray, for example, and the censored-dependent-variable models in Moffitt).

Despite the logic and popularity of these hypotheses, empirical research has not supported them. The studies reported in table 1 are representative of a large body of research that estimated Engel functions for food stamp and income variables. Especially after the elimination of the purchase requirement for food stamps in 1979, it became clear that most recipients were unconstrained, so the observed differences were considered surprising.

To test the Southworth hypothesis more formally, Senauer and Young estimated a tobit model that distinguished between constrained and unconstrained families. They found that the proportion of total income received as food stamps affected food expenditure, even when the benefits were in theory unrestricted. Thus, the authors rejected the Southworth model as "incomplete." Our model below can be seen as a way of extending and

Table 1. Estimated Marginal Propensity to Consume Food at Home, Selected Studies

Study	Marginal Propensity to Consume Food: Food Stamps	Marginal Propensity to Consume Food: Money Income
West and Price	0.30	0.05
Chavas and Yeung	0.37	0.13
Smallwood and Blaylock	0.23	0.10
Senauer and Young		
1978 Michigan PSID data	0.33	0.05
1979 Michigan PSID data	0.26	0.07
Moffitt		
Linear form	0.16	0.13
Logarithmic form	0.11	0.12
Levedahl		
Linear form	0.48	0.08
Semi-log form	0.50	0.10
Double-log form	0.29	0.09
Translog form	0.69	0.19

Sources: Condensed from Fraker; Moffitt; Levedahl.

correcting the Southworth-based model rejected by Senauer and Young, in a manner consistent with their discussion.

In the early 1990s, the U.S. Department of Agriculture (USDA) authorized four cash-out demonstrations, in part to settle this controversy. Two of these studies, in San Diego County and parts of Alabama, were "pure" cash-out experiments: recipients were randomly assigned to receive either food stamp coupons or cashable checks for the equivalent value, and no other changes were made in program operation.

In spite of this scientific design, the results were difficult to interpret (table 2). In the San Diego study, the check cohort spent on average 6.78% less than the coupon cohort on food at home per week per equivalent nutrition unit (ENU, a measure of family size in adult male equivalent units adjusted for age, sex, and number of guests at meals) (Ohls et al.).[1] In Alabama, the difference in food spending per ENU between the check and coupon cohorts was near zero and not statistically significant. When measured at the household level, the check cohort even appeared to have slightly higher food spending than the coupon cohort; although, again, this difference was not significant (Fraker et al. 1992).

In the Alabama sample 26% of all coupon recipients could be classified as constrained; while in San Diego only 5% of coupon recipients could be classified as constrained. Thus, counterintuitively, there was a significant cash-out effect in San Diego with a small number of constrained families and no significant cash-out effect in Alabama with a large number of constrained families.

[1] Technically, the survey instrument measured the money value of purchased food used at home in a given week. Throughout, we use "food spending" as shorthand for this variable.

Table 2. Impact of Food Stamp Cash-out on Food Spending, San Diego and Alabama

	Check	Coupon	Difference (%)	*t*-Statistic
Weekly food spending per ENU:				
San Diego	\$33.28	\$35.70	−6.78	2.45
Alabama	\$33.43	\$33.66	−0.68	0.31
Monthly program benefits per FCU:				
San Diego	\$116.20	\$116.46	−0.22	0.07
Alabama	\$169.27	\$168.80	0.28	0.09
Sample size:				
San Diego	542	536		
Alabama	1209	1080		

Notes: ENU is equivalent nutrition unit, a measure of household size, adjusted for age and sex of household members and for the number of guest meals. FCU is food consumption unit.
Sources: Fraker et al. 1992, 1993; and Ohls et al.

The authors of the original report on the Alabama pure cash-out demonstration offered several plausible explanations for the surprisingly small cash-out effect in that study. In short, they attributed the lack of cash-out effect in Alabama to peculiarities in the research design and sample population (Fraker et al. 1992). The authors of the San Diego report mentioned that their modest cash-out effect appeared consistent with the small number of constrained families, as would be predicted by the Southworth hypothesis (Ohls et al.). Neither report described its results as inconsistent with the Southworth hypothesis, as we find below.

Procedure

This section begins with our full nonlinear maximum-likelihood model for estimating Engel functions for food expenditure, while taking explicit account of the kinked budget constraint generated by targeted food stamp benefits. The full model accepts Southworth's account of the behavior of constrained recipients but relaxes the assumption that unconstrained recipients necessarily treat food stamps and checks alike. This section then discusses two models that can be considered special cases of the full model. Finally, it describes specifications tests and data collection.

The full model postulates that for each family i, desired (or unrestricted) food spending (F_i^*) is a function of effective full income (FY_i), a vector of economic and demographic factors (X_i), and an element of random variation (e_i) due to heterogeneity of food preferences that is not captured by the independent variables:

$$F_i^* = G(FY_i, X_i) + e_i. \tag{1}$$

If the Southworth hypothesis is true, then FY_i is simply the sum of food stamp benefits and other income. In the more general case, where these different resources can have different effects on unrestricted food spending, we assume the following:

$$FY_i = Y_i + \gamma_1 Ych_i + \gamma_2 Yco_i. \tag{2}$$

In this equation, γ_1 (the "check factor") represents the number of dollars of ordinary income (Y_i) needed to have the same effect on food spending as one dollar of check benefits (Ych_i). Likewise, γ_2 (the "coupon factor") represents the number of dollars of ordinary income needed to have the same effect on food spending as one dollar of coupon benefits (Yco_i).

Due to the fragility of food demand estimates with different functional forms, reported by Levedahl, four forms for desired food expenditure were used, namely:

(3a) Linear $$G(FY_i, X_i) = \alpha_0 + \alpha_1 FY_i + \beta' X_i,$$

(3b) Semi-log $$G(FY_i, X_i) = \alpha_0 + \alpha_1 \log[FY_i] + \beta' X_i,$$

(3c) Double-log $$\log[G(FY_i, X_i)] = \alpha_0 + \alpha_1 \log[FY_i] + \beta' X_i,$$

and

(3d) Share $$\frac{G(FY_i, X_i)}{FY_i} = \alpha_0 + \alpha_1 \log[FY_i] + \beta' X_i,$$

where α_0 and α_1 are scalar parameters, and β is a k-dimensional vector of parameters corresponding to k demographic variables. These functional forms were chosen for the range of Engel relationships they are capable of modeling and for their frequent appearance in the literature.

The linear form imposes the assumption that the marginal propensity to consume food is constant across income levels. The semi-log and double-log forms display concave curvature with respect to income, potentially describing low-income consumers' propensity to spend a higher share of marginal income on food. The share form has been discussed by Working; Leser; and Deaton and Muellbauer, and it was one of the forms estimated in Moffitt's study. Its merits include satisfying some requirements of the economic theory of consumer choice ("adding-up") and permitting the easy imposition of others ("homogeneity" and "symmetry"). However, the presence of the income variable on both sides of equation (3d) may increase the susceptibility of the share functional form to measurement error in that variable (Moffitt).

For unconstrained families (identified by the dummy parameter value D_i=0), observed food spending (F_i) equals desired food spending. For constrained families (D_i=1), observed food spending equals the value of food stamp coupons (Yco_i). Thus,

$$F_i = D_i Yco_i + (1 - D_i)[G(FY_i, X_i) + e_i], \tag{4}$$

where

$$D_i = 1 \qquad \text{if } Yco_i \geq G(FY_i, X_i) + e_i,$$

$$D_i = 0 \qquad \text{otherwise.}$$

In his article on the Puerto Rico cash-out, Moffitt appended a second error term to his version of (4) to account for measurement error in the dependent variable. If there is measurement error, the main hazard of the "one-error" model used here is that some observations could be misclassified as constrained or unconstrained. The one-error model was considered adequate for this application after a range of cutoff points for distinguishing constrained and unconstrained families was found to yield similar results.

Several assumptions have been required for econometric models of this type (Senauer and Young; Moffitt). Food is taken to be a good, so families have no reason to leave food stamp coupons unspent (the cross-sectional data did not permit considering possible savings behavior or illegal resale of food stamps). The error term is taken to be independently and identically normally distributed.

The log-likelihood function for our full model is therefore:

$$(5)\quad \log[L(\alpha_0, \alpha_1, \beta_1, \ldots, \beta_k, \gamma_1, \gamma_2, \sigma_e | F, Y, Y_{ch}, Y_{co}, X)]$$
$$= \sum_{i=1}^{n} \left[D_i \log\left[\Phi\left(\frac{Yco_i - G(FY_i, X_i)}{\sigma_e}\right)\right] + (1 - D_i)\log\left[\frac{1}{\sigma_e}\phi\left(\frac{F_i - G(FY_i, X_i)}{\sigma_e}\right)\right]\right],$$

where Φ is the standard normal cumulative distribution function, and ϕ is the standard normal density function. We maximized this log-likelihood function with respect to all parameters using the Limdep statistical package.

We also estimated two models that can be considered special cases of our full model. The first special case is a Southworth-consistent specification, which assumes that coupons, checks, and income have the same effect on desired food spending (as in the censored-dependent-variable models used by Moffitt and by Murray). To estimate this model, we restricted the "check factor" and the "coupon factor" to equal unity and maximized the resulting log-likelihood function. This restricted model is equivalent to the tobit model that would be appropriate if the Southworth hypothesis were assumed to be correct.

The second special case is a model that ignores the kinked budget constraint (as did the models used by West and Price and by Chavas and Yeung). Ignoring the kinked budget constraint is equivalent to assuming D_i equals zero in (4) for all individuals, so observed food spending equals desired food spending. We estimated this model using ordinary least squares.

An advantage of using experimental data in this study is that for the first time specification tests for desired food spending could be estimated using the check cohort alone, greatly simplifying the tests (necessarily, this procedure requires omitting the coupon variable). We used the check cohort to test for heteroskedasticity, which can lead to inconsistent estimates in nonlinear maximum-likelihood models. Also, using the full sample, we checked two alternative definitions of the dummy variable (D_i), to confirm that our results were not highly sensitive to the cutoff point used to distinguish constrained and unconstrained families.[2]

The data collection is described in the cash-out demonstration reports (Fraker et al. 1992; Ohls et al.). In-person interviews were conducted with 1,143 recipients in San Diego and 2,386 recipients in Alabama, with a response rate of 78% at each location. The food expenditure data were calculated from detailed questions on food use during the seven days preceding the interview. Only "house-keeping" households (not home-

[2] In addition, we attempted rigorous tests of the four functional forms. Box-Cox transformations can be used to compare the first three forms, but this method proved cumbersome due to the multiple parameters within FY_i. Furthermore, neither a nonnested J-test nor a test using the Box-Cox transformation were powerful enough to rule out either the linear or semi-log form in favor of the other. Formal tests were not pursued for the remaining forms.

Table 3. Mean Values for Variables Used to Estimate Food Demand, San Diego and Alabama

Variable	San Diego	Alabama
Money value of food used at home per AME	30.36	29.46
Food stamp benefits per AME ($)	12.17	18.19
Check benefits per AME ($) (0 for coupon recipients)	6.29	9.53
Coupon benefits per AME ($) (0 for check recipients)	5.88	8.66
Ordinary income per AME ($)	91.98	58.68
Full income per AME ($)	104.15	76.87
Family size in AME	2.44	2.14
Other income in household per AME ($)	20.61	4.38
Urban residence	—	0.48
Person sampled		
Is Asian	0.12	—
Is Black	0.20	0.68
Is Hispanic	0.33	0.00
Didn't complete 8th grade	0.16	0.28
Completed high school	0.57	0.41
Is <30 years old	0.44	0.24
Children present in unit	0.94	0.60
Elderly present in unit	0.02	0.25
Female head present in unit	0.85	0.87

Note: AME is adult male equivalent, a measure of household size, adjusted for age and sex of household members.

less) were included. Mean values for the variables are reported in table 3. Approximately half the recipients in each sample received coupons, and half received checks.

Results

Qualitatively, the results from the full model appear similar to the pattern observed in the simple comparisons of mean food spending between the coupon and check cohorts in the cash-out experiments. Table 4 shows a sizeable difference between the "check factor" and "coupon factor" in San Diego (where few coupon families are constrained). It shows a smaller difference between the "check factor" and "coupon factor" in Alabama (where over one-quarter of coupon families are constrained).

For conciseness, table 4 contains results with the linear and semi-log forms, which had the lowest estimated root mean squared residuals.[3] Predicted values for this goodness-of-fit measure were estimated with the full model (check cohort alone), and with the first special-case model (both cohorts). The predicted values were then transformed appropriately, in the case of the double-log and share forms, before the root mean squared

[3] Results with the double-log and share functional forms are available from the authors.

Table 4. Piecewise Linear Constraint (PLC) Models of Food Spending, San Diego and Alabama

	San Diego		Alabama	
Variable	Linear	Semi-Log	Linear	Semi-Log
Constant	29.661	13.235	26.306	−7.544
	(2.966)	(9.603)	(1.955)	(6.415)
FY or log(*FY*)	0.033**	4.272*	0.065**	8.833**
	(0.012)	(1.678)	(0.010)	(1.217)
Check factor	2.985	3.824	6.527**	6.665**
	(2.492)	(3.273)	(0.970)	(1.290)
Coupon factor	6.704*	8.491*	5.484**	5.231**
	(3.406)	(5.003)	(0.856)	(0.993)
Size in AME	−3.426**	−3.484**	−3.293**	−3.455**
	(0.458)	(0.475)	(0.388)	(0.385)
Other income	−0.001	−0.001	0.028*	0.029**
	(0.007)	(0.007)	(0.011)	(0.011)
Asian	3.584*	3.498*	—	—
	(1.804)	(1.793)	—	—
Urban	—	—	−0.410	−0.277
	—	—	(0.718)	(0.722)
Black	0.670	0.637	2.438**	2.523**
	(1.216)	(1.213)	(0.801)	(0.803)
Hispanic	1.145	1.057	5.546	5.275
	(1.195)	(1.191)	(14.480)	(14.630)
Low education	0.033	0.225	−1.364	−1.352
	(1.786)	(1.790)	(0.913)	(0.921)
High education	0.459	0.541	−1.209	−1.146
	(1.084)	(1.081)	(0.822)	(0.826)
Head < 30	0.068	0.082	−1.660	−1.189
	(1.093)	(1.096)	(1.016)	(1.003)
Kids	1.284	1.017	−1.556	−1.546
	(2.129)	(2.129)	(1.194)	(1.177)
Elderly	5.921*	6.008*	−0.259	−0.478
	(2.808)	(2.808)	(0.957)	(0.969)
Female head	1.758	1.925	−0.557	−0.441
	(1.569)	(1.574)	(0.964)	(0.973)

Notes: Asymptotic standard errors are in parentheses. One asterisk indicates significantly different from zero at 90% confidence level. Two asterisks indicate significantly different from zero at 99% confidence level.

residuals were calculated. In both San Diego and Alabama, the two models ranked the functional forms in the same order, from best to worst: linear, semi-log, double-log, share.

What table 4 reveals, which the comparisons of mean food spending between experimental cohorts do not, is that even in San Diego the observed cash-out effect is not due to the behavior of constrained recipients. Rather, food stamp coupons, food stamp checks, and ordinary income have distinct effects on desired food spending, as well as on observed food spending.

Conveniently, the first special case (the model with the "check factor" and "coupon factor" constrained to equal unity) can be viewed simulteneously as a test of the statistical significance of these differences. In San Diego, a likelihood ratio test rejected the restricted model at the 95% confidence level with the double-log functional form and at the 90% confidence level with the linear and semi-log functional forms. These results replicate Senauer and Young's rejection of the Southworth-consistent specification and confirm the need for a less restrictive model.

In Alabama, the restricted model was rejected at the 99% confidence level with all four functional forms, but this rejection does not indicate the same type of cash-out effect that was found in San Diego. The restricted model was rejected primarily because both coupons and checks had different effects from ordinary income (a phenomenon that is discussed below). A more specific test of the equality of the "check factor" and "coupon factor" alone yielded lower chi-squared statistics for the likelihood ratio test, although surprisingly, these statistics were still significant at the 95% confidence level for all but the double-log functional form.

Taken at face value, these latter tests would indicate that in Alabama the small positive difference in food spending between the check cohort and the coupon cohort is statistically significant, unless the double-log functional form is the correct specification. While this pattern seems implausible at first, such a pattern could arise if the illegal resale of food stamps were common before the experiment began. In that case, one effect of cash-out would be to save some check recipients from the expense of paying a "penalty" to food stamp traffickers, thereby increasing the net income these recipients have for spending on food and other goods.

The second special case (the model that ignores the kink) yielded estimated parameters for the "check factor" and "coupon factor" in both experiments that generally were within 17% of those found with the full model (table 5). The qualitative implications of the full model and this second special case were effectively identical: the estimated cash-out effect was sizeable in San Diego but small or nonexistent in Alabama.

For the test of heteroskedasticity, we used the linear functional form and the check cohort alone to estimate a model where the variance of the error term was presumed to have the form $\sigma_e^2 FY_i^\gamma$. In both San Diego and Alabama, a Lagrange multiplier test failed to reject the null hypothesis of homoskedasticity (γ=0) at the 0.05 significance level. While this test did reject the null at the 0.10 significance level for the San Diego sample, heteroskedasticity-corrected estimates were almost exactly the same as the estimates from OLS (the difference was less than 3% for the income and benefit parameters).

To confirm that the results of the full model were not highly variable according to how the kink was defined, we also used food spending levels of 1.1 times *Yco* and 1.5 times *Yco,* instead of *Yco* itself, as cutoff points for distinguishing constrained and unconstrained families. These definitions can be seen as a way of accounting for the possibility that families may buy some food with cash even before their food stamp coupons are used up, due to convenience or budgeting difficulty. Slope parameters from these alternate models differed little from the parameters for the full model we report in table 4.

Discussion

Our results confirm that the Southworth-consistent model (the first special case) is not a reliable basis for empirical specifications, at least within the range of models studied here. The key assumption that appears to be incorrect is that food stamp coupons have the same effect as other income on desired food spending. By contrast, our results suggest

Table 5. Models of Food Spending Ignoring the "Kink," San Diego and Alabama

	San Diego		Alabama	
Variable	Linear	Semi-Log	Linear	Semi-Log
Constant	29.510	12.054	26.098	−10.026
	(2.934)	(9.465)	(1.858)	(6.004)
FY or log(*FY*)	0.033**	4.535**	0.062**	9.182**
	(0.012)	(1.653)	(0.010)	(1.137)
Check factor	3.472	4.076	6.622**	6.716**
	(2.490)	(3.175)	(0.947)	(1.211)
Coupon factor	7.862*	9.515*	6.376**	6.494**
	(3.586)	(5.149)	(0.916)	(1.141)
Size in AME	−3.392**	−3.460**	−3.052**	−3.188**
	(0.452)	(0.468)	(0.389)	(0.383)
Other income	−0.001	0.000	0.024*	0.025*
	(0.007)	(0.007)	(0.010)	(0.010)
Asian	3.538*	3.464*	—	—
	(1.787)	(1.775)	—	—
Urban	—	—	−0.521	−0.408
	—	—	(0.674)	(0.677)
Black	0.637	0.622	2.442**	2.526**
	(1.197)	(1.193)	(0.748)	(0.750)
Hispanic	1.884	1.102	3.169	3.303
	(1.183)	(1.178)	(11.540)	(11.520)
Low education	0.167	0.359	−1.443*	−1.403
	(1.778)	(1.784)	(0.852)	(0.859)
High education	0.481	0.571	−1.045	−0.955
	(1.069)	(1.066)	(0.776)	(0.779)
Head < 30	0.002	0.020	−1.659*	−1.195
	(1.077)	(1.081)	(0.977)	(0.961)
Kids	1.133	0.745	−1.615	−1.648
	(2.108)	(2.102)	(1.140)	(1.121)
Elderly	5.841*	5.845*	−0.114	−0.292
	(2.753)	(2.755)	(0.893)	(0.902)
Female head	1.665	1.855	−0.484	−0.368
	(1.562)	(1.566)	(0.898)	(0.905)

Notes: Asymptotic standard errors are in parentheses. One asterisk indicates significantly different from zero at 90% confidence level. Two asterisks indicate significantly different from zero at 99% confidence level.

that a model ignoring the kinked budget constraint (the second special case) would have made little difference in key parameter estimates, at least for the data studied here. This conclusion to some extent corroborates the simple empirical models used in the 1970s and early 1980s, before it became common to incorporate the Southworth hypothesis explicitly. The best argument for preferring our full model over these simpler models is that, as long as nonlinear regression techniques are available, it makes sense to let the data decide whether the behavior of constrained recipients is significant enough to affect the empirical estimates.

In supporting our full model, it is also necessary to consider what decision-making process might lead recipients to behave as the model describes. One reason that Senauer and Young's article has not convinced the theoretical and empirical literatures to forego relying on the Southworth propositions is that these propositions are a straightforward consequence of rational choice subject to the traditional kinked budget constraint. The challenge for alternatives to the traditional theory is to answer the following question: why do coupon recipients who can afford the bundle of goods preferred by check recipients choose a different bundle?

Two approaches to answering this question are raised here, but both are still underdeveloped. First, game-theoretic models have been used to address intrahousehold control questions (Montalto). If different types of benefits are systematically allocated to different members of the household, the distinct impact of coupons and checks can be explained.

Though anecdotal, comments from focus group participants can illuminate these intrahousehold issues. According to one male check recipient in Alabama:

> It [checks] is a hassle in my family. I don't really care. My wife is the one who cares about it. . . . She's afraid I'm gonna end up in some local beer joint drunker than hell from now on. . . . We spend more in food than what the check amounts to anyway. She'd rather have food stamps, you know, than to get the check (Mazur and Ciemnecki, p. 15).

Second, models of "quasi-rational" choice (Thaler and Shefrin) suggest early steps toward dealing with difficult issues of self-restraint and wise decision making that seem to be on the minds of recipients and policymakers alike, but which have proven difficult to address with traditional economic models. If food stamp coupons help recipients to organize their monthly budget in a fashion that favors food spending, the distinct effect can again be explained. Corroborating an earlier conjecture by Fraker that some recipients might actually prefer coupons to checks for budgeting reasons, sizeable proportions of the San Diego sample agreed or strongly agreed with statements that stamps are better than checks for budgeting (52%) or stamps give more control over food expenditure (64%). As a matter of epistemology, the people being modeled by an economic theory need not hold opinions consistent with that theory—they need only make spending decisions as if they do—but these responses still appear inconsistent with the predictions of the Southworth model.

Taken literally, models of intrahousehold behavior and models of quasi-rational choice would lead to more complex specifications than have been attempted here. Indeed, for each alternative model the ideal specification would require data that are not commonly available. As an approximation, our full-model shares some of the essential characteristics of these alternative models: coupon families must spend at least their food stamp allotment on food, but even if the family overall is inframarginal, family members' utility-maximizing choices may still differ from those they would have made had they received checks.

Finally, the Alabama cash-out experiment confirmed the possibility, which had been discussed earlier (Fraker), that even check benefits might have a greater effect on food spending than ordinary income, if recipients are informed that the check benefits come from the Food Stamp Program. In San Diego, where food stamp check benefits and Aid to Families with Dependent Children were combined in a single monthly check, check benefits were treated more like ordinary income. The behavior of the Alabama recipients is consistent with, for example, Senauer and Young's suggestion that recipients might feel morally obligated to use food stamp benefits to increase their food spending.

Because recipients may treat food stamp checks differently under different cash-out program designs, the best estimates of the effect of replacing food stamp coupons with food stamp checks must still come from actual experiments. By contrast, adequate comparisons of food stamp coupons and ordinary income can be acquired using less expensive nonexperimental models. However, these models must allow for the distinct effects of food stamps and cash resources on desired food spending.

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References

Chavas, J.-P., and M. L. Yeung. "Effects of the Food Stamp Program on Food Consumption in the Southern United States." *S. J. Agr. Econ.* 14(1982):131–39.

Deaton, A., and J. Muellbauer. *Economics and Consumer Behavior.* Cambridge: Cambridge University Press, 1980.

Fraker, T. M. *The Effects of Food Stamps on Food Consumption: A Review of the Literature.* Washington DC: Mathematica Policy Research, Inc., 1990.

Fraker, T. M., A. P. Martini, J. C. Ohls, M. Ponza, and E. A. Quinn. *The Evaluation of the Alabama Food Stamp Cash-Out Demonstration.* Washington DC: Mathematica Policy Research, Inc., 1992.

———. "Food Stamp Cash-out Demonstration Findings: San Diego and Alabama." In *New Directions in Food Stamp Policy Research,* eds., N. Fasciano, D. Hall, and H. Beebout, pp. 27–47. Washington DC: U.S. Department of Agriculture, 1993.

Gahvari, F. "In-Kind versus Cash Transfers in the Presence of Distortionary Taxes." *Econ. Inquiry.* 33(1995): 45–53.

Leser, C. "Forms of Engel Functions." *Econometrica* 31(1963):694–703.

Levedahl, J. W. "The Role of Functional Form in Estimating the Effect of a Cash-Only Food Stamp Program." *J. Agr. Econ. Res.* 43(1991):11–19.

Mazur, M., and A. Ciemnecki. *Recipient Attitudes toward and Experiences with Food Stamp Checks and Coupons. Evidence from the Alabama Food Stamp Cash-Out Focus Groups.* Princeton NJ: Mathematica Policy Research, Inc., 1991.

Mittelhammer, R., and D. A. West. "Food Stamp Participation among Low-Income Households: Theoretical Considerations of the Impact on the Demand for Food." *S. J. Agr. Econ.* 7(1975):223–31.

Moffitt, R. "Estimating the Value of an In-Kind Transfer: The Case of Food Stamps." *Econometrica* 57(1989): 385–409.

Montalto, C. P. "Relative Bargaining Power of Spouses: Empirical Results Using an Economic Model of Cooperative Bargaining." Unpub. Ph.D. diss., Cornell University, 1992.

Murray, M. P. "How Inefficient Are Multiple In-Kind Transfers?" *Econ. Inquiry* 32(1994):209–27.

Ohls, J. C., T. M. Fraker, A. P. Martini, and M. Ponza. *The Effects of Cash-Out on Food Use by Food Stamp Program Participants in San Diego.* Princeton NJ: Mathematica Policy Research, Inc., 1992.

Senauer, B., and N. Young. "The Impact of Food Stamps on Food Expenditures: Rejection of the Traditional Model." *Amer. J. Agr. Econ.* 66(1986):37–43.

Smallwood, D. M., and J. R. Blaylock. "Analysis of Food Stamp Program Participation and Food Expenditures." *West. J. Agr. Econ.* 10(1985):41–54.

Southworth, H. M. "The Economics of Public Measures to Subsidize Food Consumption." *J. Farm Econ.* 27(1945):38–66.

Thaler, R., and H. M. Shefrin. "An Economic Theory of Self-Control." In *Quasi-Rational Economics,* ed., R. Thaler, pp. 77–90. New York: Russell Sage Foundation, 1991.

West, D. A., and D. W. Price. "The Effects of Income, Assets, Food Programs, and Household Size on Food Consumption." *Amer. J. Agr. Econ.* 58(1976):725–30.

Working, H. "Statistical Laws of Family Expenditure." *J. Amer. Statis. Assoc.* 38(1943):43–56.