Unobserved Preference Heterogeneity in Demand Using Generalized Random Coefficients

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Abstract

We model unobserved preference heterogeneity in demand systems as random Barten scales in utility functions. These Barten scales appear as random coefficients multiplying prices in demand functions. Consumer demands are nonlinear in prices and may have unknown functional structure. We therefore prove identification of additive Generalized Random Coefficients models, defined as additive nonparametric regressions where each regressor is multiplied by an unobserved random coefficient having an unknown distribution. Using Canadian data, we estimate energy demand functions with and without random coefficient Barten scales. We find that not accounting for this unobserved preference heterogeneity substantially biases estimated consumer-surplus costs of an energy tax.

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1   Introduction

For discretely demanded goods, unobserved preference heterogeneity is typically modeled using random coefficients, as in Berry, Levinsohn, and Pakes (BLP 1995). Allowing for substantial unobserved random preference heterogeneity, as BLP does, has proven to be necessary for realistic evaluations of the impacts of price changes on demand. In this paper we propose an analogous way to introduce unobserved preference heterogeneity in continuous demand systems.

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The application we consider is energy demand by consumers. Energy is consumed in continuous quantities and displays substantial nonlinearities in income and price effects. Therefore, energy cannot be appropriately modeled using discrete demand methods like BLP, and instead requires the methodology of continuous demand systems.

We demonstrate the importance of accounting for random coefficient type unobserved preference heterogeneity in energy demand. In particular, we show that failure to do so results in a dramatic underestimate of the variance of impacts of energy price changes across consumers. Accounting for this variation is crucial for correctly assessing the true costs to society of energy policies such as a carbon tax. We show that measures of social welfare that ignore this unobserved preference heterogeneity yield substantially biased estimates of the full costs to society of an energy tax on consumers, by failing to fully account for the tax's distributional impacts.

One of the most commonly used methods for incorporating *observable* sources of preference heterogeneity (such as the impacts of age or family size) in continuous demand systems is via Barten (1964) scales. Barten scales deflate the prices faced by consumers, and so have a structure that is analogous to random coefficients on prices, in that they multiply each price in the demand system. This suggests that a natural way to introduce unobserved preference heterogeneity into continuous demand systems is to allow random variation in the Barten scales via random coefficients on prices.

However, randomly varying Barten scales introduces a substantial econometric difficulty because, unlike discrete demand models such as multinomial logit, realistic continuous demand models are highly nonlinear in prices, due to constraints such as homogeneity and Slutsky symmetry. We therefore require a general type of random coefficients that can be identified and estimated in nonlinear, or even nonparametrically specified, demand functions. We define "generalized random coefficients" to be random coefficients applied to variables in a general nonlinear or nonparametric model, in contrast to ordinary random coefficients that are applied in linear index models.

In this paper we first provide some identification theorems, showing that the joint distribution of random coefficients can be nonparametrically identified in nonlinear, and in additive nonparametric, regression models.

We then apply these results to identification of random Barten scales in demand systems. This application includes proving a new theorem that nonparametrically characterizes the preferences associated with demand functions having a certain additive structure. This result is relevant because it allows us to exploit the regularity conditions required for nonparametric identification of generalized random Barten coefficients in additive models.

Based on these identification theorems, we estimate energy demand functions for a set of Canadian consumers. To illustrate the importance of allowing for unobserved heterogeneity in Barten scales, we evaluate the (partial equilibrium) impacts of a hypothetical tax on energy goods, like a carbon tax. Among other results, we find that allowing for unobserved preference heterogeneity has a large impact on the estimated distribution of the relative costs (consumer surplus impacts) of the tax. For example, we find that this distribution across consumers has a standard deviation that is more than twice as large in our model compared to an analogous model that does not allow for such unobserved preference heterogeneity.

Consider first our proposed generalization of random coefficients models. Suppose an observed variable Y depends on a vector of observed regressors $X = (X_1, ..., X_K)$, and on a set

of unobserved errors $U_0, U_1, ..., U_K$ that are (possibly after conditioning on other covariates Z) independent of X. We propose a generalized random coefficients model given by

$$Y = G(X_1 U_1, ..., X_K U_K) \quad \text{or} \quad Y = G(X_1 U_1, ..., X_K U_K) + U_0 \tag{1}$$

for some function G. We focus mainly on results for the special case of equation (1) where G takes the additive model form

$$Y = \sum_{k=1}^{K} G_k(X_k U_k) + U_0 \tag{2}$$

and the functions $G_1, ..., G_K$ are unknown.

In these models the vector $U = (U_1, ..., U_K)$ represents unobserved heterogeneity in the dependence of Y on X, while U_0, if present, represents measurement error or other independent variation in Y. We provide conditions under which the joint distribution of the vector U is nonparametrically identified. If present, U_0 is assumed independent of these other errors and has a marginal distribution that is also nonparametrically identified.

In our empirical application, Y will be a measure of energy demanded by a consumer, G will be a Marshallian demand function, each X_k will be the price of a good k divided by a consumer's total expenditures, and each U_k (other than U_0) will be a Barten scale. All previous empirical implementations of Barten scales have exactly these forms, but with every U_k other than U_0 specified as deterministic functions of observable characteristics that affect preferences, such as age or family size. In contrast, we allow the Barten scales to be random, and show their joint distribution can be nonparametrically identified, under low level regularity conditions.

One of our identification theorems shows that if G is known, then under mild conditions the joint distribution of the elements of U is nonparametrically identified. We also provide a theorem giving conditions under which, in equation (2), each function G_k can be nonparametrically identified (unlike our other identification result, this theorem employs an argument analogous to identification at infinity). Combining both theorems then allows us to simultaneously nonparametrically identify the joint distribution of U and nonparametrically identify each G_k function. Combining both theorems also provides some overidentification that we show can be exploited to generalize the model a bit (relaxing the additivity assumption by adding some interaction terms), and to some extent also relaxes our dependence on identification at infinity type arguments.

Imposing the additivity of equation (2) directly on Marshallian demand functions yields some implausible restrictions on preferences. However, we show that, when $K = 2$, these restrictions can be relaxed by suitably transforming Y. In particular, we prove a theorem showing that when $K = 2$, if Y is defined as a logit transformed budget share, then demands will take the additive form implied by equation (2) if and only if indirect utility has a correspondingly additive form. This theorem also provides closed form expressions for the indirect utility function corresponding to nonparametrically specified demand functions that are additive in this way. These closed form expressions greatly simplify our later consumer surplus and welfare calculations.

We first provide a literature review bearing on the econometric identification of models containing random coefficients and on the modeling of preference heterogeneity in continuous

demand systems. We then present our main identification theorems, followed by our theorem characterizing the nonparametric connection between preferences and logit transformed demands. We next provide our empirical implementation of the random Barten scales model, including consumer surplus calculations on the hypothetical impacts of large increase in the price of, or taxes on, energy goods. We then conclude, although we follow this with an appendix that includes proofs of Theorems and an extensive set of analyses verify the robustness of our empirical results to a wide variety of alternative model specifications, including adding complexity to the utility function specification, relaxing the parametric structure on preference and error distributions, and dealing with potential endogeneity of regressors.

2 Literature Review

We use generalized random coefficients to represent price scales in consumer demand models. There is a long history of using such scales to empirically model observed sources of preference heterogeneity. See, e.g., Rothbarth (1943), Prais and Houthakker (1955), Barten (1964), Pollak and Wales (1981) and Jorgenson, Lau, and Stoker (1982), and see Lewbel (1997) for a survey. Barten (1964) type price scales (hereafter: *Barten scales*) take the form of multiplying each price in a demand function by a preference heterogeneity parameter, as in equation (1). It is therefore a natural extension of this literature to include unobserved preference heterogeneity in Barten scales.

We apply estimated demand functions and estimated Barten scale distributions to do welfare analyses. In particular, we use a Barten scaled energy demand function to perform consumer surplus calculations for an energy price change (as in Hausman 1981). Our consumer surplus calculations can be interpreted as a variant of Hoderlein and Vanhems (2011), who introduce unobserved preference heterogeneity into the Hausman model. The first of these two papers introduced scalar preference heterogeneity into the model nonparametrically, while the latter incorporated heterogeneity in the form of ordinary linear random coefficients. As an alternative to modeling unobserved heterogeneity, Hausman and Newey (2014) provide bounds on average consumer surplus.

In contrast, our model follows the prior consumer demand literature by including preference heterogeneity in the form of Barten scales, differing from the prior demand literature in that our Barten scales include unobserved heterogeneity (a smaller additional difference is the way we also include an additive measurement error). We also apply our empirical results to estimate Atkinson (1970) type social welfare functions, and thereby analyze the extent to which allowing for unobserved preference heterogeneity affects estimated tradeoffs between mean impacts and inequality of impacts of a tax or price change in energy.

Other papers that introduce nonseparable unobserved preference heterogeneity in continuous demand systems include Brown and Walker (1989), Lewbel (2001), Beckert (2006) Matzkin (2007b), and Beckert and Blundell (2008). Lewbel and Pendakur (2009) propose a continuous demand system model in which the standard separable errors equal utility parameters summarizing preference heterogeneity, and do welfare calculations showing that accounting for this unobserved heterogeneity has a substantial impact on the results. Lewbel and De Nadai (2011) show how preference heterogeneity can be separately identified from measurement errors. A

related empirical model to ours is Comon and Calvet (2003), who use repeated cross sections and deconvolution to identify a distribution of unobserved heterogeneity in income effects.

Nonparametric identification and estimation of ordinary random coefficients models is considered by Beran and Hall (1992), Beran, Feuerverger, and Hall (1996) and Hoderlein, Klemelae, and Mammen (2010). Recent generalizations include random coefficient linear index models in binary choice, e.g., Ichimura and Thompson (1998), Gautier and Kitamura (2010), and semiparametric extensions of McFadden (1974) and Berry, Levinsohn, and Pakes (1995) type models, e.g., Berry and Haile (2009).

Ordinary random coefficients are the special case of the additive model in equation (2) in which each G_k is the identity function. Additive models are a common generalization of linear models; see, Hastie and Tibshirani (1990), Linton (2000), and Wood (2006), and in the particular applications of additivity to consumer demand systems include Gorman (1976) and Blackorby, Primont, and Russell (1978).

This paper also contributes to the literature on estimation of models with nonseparable errors, in particular where those errors arise from structural heterogeneity parameters such as random utility parameters. Older examples of such models include Heckman and Singer (1984) and Lewbel (2001). More recent work focusing on general identification and estimation results include Chesher (2003), Altonji and Matzkin (2005), Hoderlein, and Mammen (2007), Matzkin (2007a, 2008), and Imbens and Newey (2009).

Fox and Gandhi (2013) provide general conditions for identification of random utility parameters in multinomial choice problems, including linear index models with random coefficients, and models analogous to Berry and Haile (2009) that exploit Lewbel (2000) type special regressors. They note that the only general sufficient condition known for one of their identifying assumptions is utility functions that are real analytic functions.

A related result to ours is Hoderlein, Nesheim, and Simoni (2011), who provide a high level condition they call T-completeness that suffices for nonparametric identification of a vector of random parameters within a known function. They provide some examples where T-completeness can be shown to hold, such as when error distributions are in the exponential family, or are parameterizable by a single scalar. Our model when G is known is a special case of their general setup, and so our theorem proving identification for this model provides a new framework where T-completeness could be satisfied. More generally, one goal of our analysis is to provide relatively low level conditions that serve to identify our model, instead of high level, difficult to verify conditions as in Fox and Gandhi (2013), or like T-completeness.

Perhaps the result that comes closest to our identification theorem is Matzkin (2003), which in an appendix describes sufficient conditions for identification of a general class of additive models with unobserved heterogeneity. The biggest difference between our results and Matzkin (2003) is that we identify the joint distribution of U, while Matzkin assumes the elements of U are mutually independent. However, even our model when $K = 1$ (the case where there is no joint distribution to be identified) while employing a structure very similar to Matzkin's, does not satisfy her identification assumptions and so even in that case our Theorem is new and cannot be derived from her results.

3 Generalized Random Coefficient Model Identification

In this section we first consider additive models given by equation (2). Specifically, Theorem 1 below shows nonparametric identification of each function G_k and the marginal distribution of each random coefficient U_k. We then provide, in Theorem 2, separate assumptions under which the joint distribution of the random coefficients can be identified in the more general model of equation (1) when G is known. We then combine both theorems to nonparametrically identify both the joint distribution of random coefficients and the functions G_k. We also discuss conditions for identification of some more general models.

Theorem 1 employs an argument similar to identification at infinity (actually identification at zero), or what Khan and Tamer (2010) call "thin set identification." However, when we combine Theorems 1 and 2 we obtain very many overidentifying restrictions, and therefore our results do not depend solely on thin set identification.

Later sections provide the connections between these theorems and our Barten scales model of demand. However, we note upfront that in our empirical application X is positive (though not bounded away from zero), so it is relevant that our identification theorems allow for zero being on the boundary of the closure of the support of X.

For any random vectors A and B let $F_{A|B}(a \mid b)$ and $f_{A|B}(a \mid b)$ denote the conditional cumulative distribution function and conditional probability density function, respectively, of A given B. Let e_k be the K vector containing a one in position k and zeros everywhere else. Let $X_{(k)}$ denote the $K-1$ vector that contains all the elements of X except for X_k.

3.1 Additive Model and Marginal Distribution Identification

We first consider identification of the G_k functions and of the separate distributions of each U_k in equation (2). Here Y is the dependent variable, X is a set of covariates having random coefficients, and Z is a set of additional covariates that may affect the distribution of U.

ASSUMPTION A1: The conditional distribution $F_{Y|X,Z}(y \mid x, z)$ and the marginal distribution $F_Z(z)$ are identified. $(U_0, U_1, ..., U_K) \perp X \mid Z$ and $(U_1, ..., U_K) \perp U_0 \mid Z$. Either U_0 has a nonvanishing characteristic function (conditional on Z)[1] or U_0 is identically zero. $supp(U_0) \subseteq supp(Y)$. X has rectangular support and $\{0, e_1, ..., e_K\}$ is a subset of the closure of $supp(X)$.

ASSUMPTION A2: For $k \in \{1, ..., K\}$, $(U_k, X_k) \mid Z$ is continuously distributed, and for every $r \in supp(X_k U_k)$ there exists an $x_k \in supp(X_k)$ such that $f_{U_k}\left(x_k^{-1} r\right) \neq 0$.

ASSUMPTION A3: For $k \in \{1, ..., K\}$, G_k is a strictly monotonically increasing function. The free location and scale normalizations $G_k(0) = 0$ and $G_k(1) = y_0$ for some known $y_0 \in supp(Y)$ are imposed.

Assumption A1 first assumes identification of $F_{Y|X,Z}(y \mid x, z)$ and $F_Z(z)$, which would in general follow from a sample of observations of Y, X, Z with sample size going to infinity.

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[1] Formally, the condition on U_0 regarding a nonvanishing characteristic function required for the deconvolution step of the proof is only that the set of $t \in \mathbb{R}$ for which $E\left(e^{itU_0}\right) \neq 0$ is dense in $\mathbb{R}$. See, e.g., Meister (2005).

Identification of $F_{Y|X,Z}(y \mid x, z)$ is actually stronger than necessary for Theorem 1, since only certain features of this distribution are used in the proof. For example, it would suffice to only identify $F_{Y|X,Z}(y \mid x_k e_k, z)$ for $k = 1, ..., K$. However, more information regarding $F_{Y|X,Z}$ is used in Theorem 2 and other extensions.

Assumption A1 imposes conditional independence and support requirements on U, X and Z. The role of Z is to permit the error U_0 and random coefficients U_k to be correlated with X, thereby allowing elements of X to be endogenous. See, e.g., the correlated random coefficients model of Heckman and Vytlacil (1998). This allows for Heckman and Robb (1986) control function type endogeneity, with Z being control function residuals as in Blundell and Powell (2003, 2004). In particular, if $X_k = h_k(X_{(k)}, Q) + Z_k$ for some observed instrument vector Q and some identified function h_k (typically h_k would be $E(X_k \mid X_{(k)}, Q)$), then the conditional independence assumptions in A1 correspond to standard control function assumptions. Note that Z can be empty, so all the results given below will hold if there is no Z, in which case U is independent of X and so the regressors X are exogenous. The assumptions also permit Z to be discrete, and place no restriction on the dimension of Z, although control function residuals would generally be continuous and have dimension equal to the number of endogenous elements of X.

Assumption A2 assumes that the regressors and random coefficients are continuously distributed. Assumption A2 also calls for a mild relative support assumption on X_k and U_k. Later Theorem 2 will require a stronger support restriction.

The normalizations in Assumption A3 are free normalizations, because first if $G_k(0) \neq 0$ then we can redefine $G_k(r)$ as $G_k(r) - G_k(0)$ and redefine U_0 as $U_0 + G_k(0)$, thereby making $G_k(0) = 0$. Next, given a nonzero $y_0 \in supp(Y)$, there must exist a nonzero r_0 such that $G_k(r_0) = y_0$. We can then redefine U_k as $r_0 U_k$ and redefine $G_k(r)$ as $G_k(r/r_0)$, thereby making $G_k(1) = y_0$. These particular normalizations are most convenient for proving Theorem 1 below, but in empirical applications alternative normalizations may be more natural, e.g., choosing location to make $E(U_0) = 0$.

What follows is our first identification theorem, which as noted in the literature review is closely related to, but is not a direct corollary of, results in Matzkin (2003).

THEOREM 1: Let $Y = \sum_{k=1}^{K} G_k(X_k U_k) + U_0$ and let Assumptions A1, A2, and A3 hold. Then the functions $G_1, G_2, ..., G_K$ and the distributions $F_{U_0|Z}, F_{U_1|Z}, ... F_{U_K|Z}$ are all nonparametrically identified.[2]

It follows immediately from Assumption A1 and Theorem 1 that the marginal distributions $F_{U_0}, F_{U_1}, ... F_{U_K}$ are also nonparametrically identified. In applications we would generally assume that Assumptions A2 and A3 hold for all $k \in \{1, ..., K\}$.

As noted earlier, Theorem 1 depends on an identification at zero argument, however, there is considerable overidentifying information in the data that Theorem 1 does not use. For ex-

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[2]The proof of Theorem 2 involves evaluating the distribution of Y given X where either $X = 0$ or all but one element of X equals zero. This means conditioning on a set of measure zero. The same applies to Theorem 1 regarding conditioning on Z at a point if Z is continuous. Note, however, that issues of nonuniqueness of the limiting argument (the Borel-Kolmogorov paradox) do not arise here, since the identification proof depends only on transformations of smooth conditional density and expectation functions. It would be possible to recast the proofs in terms of conditioning on sets $\| X \| \leq c$ and taking limits as $c \to 0$.

ample, given a G_k and $F_{U_k|Z}$ function identified by Theorem 1, one piece of overidentifying information would be to construct the function $\int_{u \in supp(U_k|Z=z)} \left[\partial G_k (x_k u) / \partial x_k \right] dF_{U_k|Z} (u \mid z)$ and verify that this equals the function $\partial E (Y \mid X = x, Z = z) / \partial x_k$ where x_k is the $k'th$ element of the vector x. Later Theorem 2 below will provide considerably more overidentifying information.

In our notation, Matzkin (2003) considers models of the form $Y = \sum_{k=0}^{K} M_k (S, X_k, U_k)$ where M_k are unknown functions and S are additional observed covariates. Our Theorem 1 fits this general framework with $M_k (S, X_k, U_k) = G_k (X_k U_k)$ and $G_0 (X_0 U_0) = U_0$ with $X_0 = 1$. As we do, Matzkin assumes that M_k is monotonic in U_k and that U and X are continous and conditionally independent. She proposes alternative restrictions or normalizations that suffice to identify each function M_k and distribution F_{U_k}, but none of her proposed restrictions fit our generalized random coefficients framework, and as a result Theorem 1 is not a corollary of her results. Her closest result to our framework is the assumption that $M_k (S, X_k, U_k) = N_k (S, X_k - U_k)$ for some function N_k. By redefining X_k and U_k as $\ln X_k$ and $\ln U_k$, we can replace our $G_k (X_k U_k)$ with $g_k (X_k - U_k)$ for a suitably redefined function g_k. However, we still cannot apply Matzkin's result by equating $N_k (S, X_k - U_k) = g_k (X_k - U_k)$, because for this specification Matzkin (2003) requires the presence of at least one additional regressor S that has certain properties, and our model has no such additional regressors inside the G_k functions.[3] Another point of difference between our Theorem 1 and Mazkin (2003) is that she assumes the U_k functions are mutually (conditionally) independent, while our Theorem 1 does not impose either conditional or unconditional independence.

A small extension to Theorem 1 is the following.

COROLLARY 1: Let $Y = G (X_1 U_1, ..., X_K U_K) + U_0$ for any function G that includes $X_k U_k e_k$ in its domain, for $k = 1, ..., K$. Then
i) There exists functions $G_1, ..., G_K$, and $\widetilde{G}$ such that

$$Y = \widetilde{G} (X_1 U_1, ..., X_K U_K) + \sum_{k=1}^{K} G_k (X_k U_k) + U_0 \qquad (3)$$

where the function $\widetilde{G} (X_1 U_1, ..., X_K U_K)$ equals zero when all but one of its elements equal zero, and
ii). Theorem 1 holds replacing $Y = \sum_{k=1}^{K} G_k (X_k U_k) + U_0$ with equation (3).

In Corollary 1, the function $\widetilde{G}$ is not identified, so the main points of this corollary are first that any function G can be decomposed into an additive part $\sum_{k=1}^{K} G_k$ and an interactions part $\widetilde{G}$, and second that the presence of the interaction function $\widetilde{G}$ does not interfere with identification of the G_k and $F_{U_k|Z}$ functions. We later apply Corollary 1 in a context where $\widetilde{G}$ can be identified by other means.

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[3] An alternative identifying restriction Matzkin (2003) proposes has S empty, but assumes $M_k (\widetilde{x}_k, U_k) = U_k$ for some known value $\widetilde{x}_k$. However, in our model this would require $G_k (\widetilde{x}_k U_k) = U_k$, which only holds when G_k is proportional to the identity function. One other restriction she considers is that M_k be linearly homogeneous in X_k and U_k, which cannot hold for our G_k.

3.2 Joint Distribution Identification

Theorem 1 only identifies the distribution function of each U_k. We now make additional assumptions sufficient to identify the joint distribution $F_{U|Z}(U_1,...,U_K \mid Z)$. Theorem 2 below shows identification when G is known in equation (1). We will then combine both Theorems to obtain overidentification of equation (2). The fact that we will have overidentification should mitigate some concern about making use of identification at zero arguments in Theorem 1.

ASSUMPTION A4: The conditional distribution $F_{\widetilde{Y}|X,Z}(y \mid x, z)$ and the marginal distribution $F_Z(z)$ are identified. $U \perp X \mid Z$. The distribution of $\left(U_1^{-1}, ..., U_K^{-1} \mid Z\right)$ is identified from its integer moments. $supp(X)$ is rectangular, $supp(X \mid Z) = supp(X)$, and the closure of $supp(X \mid Z)$ equals the closure of $supp(U_1 X_1, ..., U_K X_K \mid Z, U)$.

Necessary and sufficient conditions for integer moments to identify a distribution are known, and are weaker than the conditions needed for existence of a moment generating function. See, e.g., Assumption 7 of Fox, Kim, Ryan, and Bajari (2012). In our empirical application it will be reasonable to assume that each U_k is bounded and bounded away from zero, which is sufficient and stronger than necessary.

Given $U \perp X \mid Z$, the support condition in Assumption A4 could be satisfied in a few different ways. Most usefully for our application, the support condition holds if $supp(X \mid Z) = \mathbb{R}_+^K$ and $supp(U \mid Z) \subseteq \mathbb{R}_+^K$. It will also hold if $supp(X \mid Z) = \mathbb{R}^K$ and U has any support that excludes the origin.

ASSUMPTION A5: $\widetilde{Y} = G(X_1 U_1, ..., X_K U_K)$ for some identified function G.

Let $t = (t_1, ..., t_K)$ denote a K vector of positive integers. For a given function h_ρ and vector t, define $\kappa_{\rho,t}$ by

$$\kappa_{\rho,t} = \int_{supp(X)} h_\rho [G(s_1, ...s_K), t_1, ..., t_K] s_1^{t_1-1} s_2^{t_2-1} ... s_K^{t_K-1} ds_1 ds_2 ... ds_K \qquad (4)$$

The subscript ρ is present because we will later consider multiple h_ρ functions, in which case ρ will index the choice of function.

ASSUMPTION A6: Given G, for any K vector of positive integers t we can find a nonnegative, bounded function h_ρ such that $\kappa_{\rho,t}$ is finite and nonzero.

Assumption A6 is a rather mild restriction on G, because the function h_ρ is freely chosen, based on knowing G and t. For example, in our later application where $supp(X) = \mathbb{R}_+^K$, choosing h_ρ to be nonnegative and somewhere nonzero will ensure $\kappa_{\rho,t}$ is nonzero, and finiteness will follow if h_ρ has sufficiently thin tails when any s_k goes to infinity. For example, if $G(s_1, ...s_K)$ is a polynomial, then it goes to plus or minus infinity if and only if any s_k goes to infinity, so in this case we may choose h_ρ to be a thin tailed density function on $\mathbb{R}_+^K$ to satisfy Assumption A6. The same choice of h_ρ would also work if $G(s_1, ...s_K)$ equaled a polynomial multiplied by any nonnegative bounded function, or indeed any finite sum of products of polynomials and nonnegative bounded functions.

THEOREM 2: Let Assumptions A4, A5, and A6 hold. Then the joint distribution function $F_{U|Z}(U_1,...,U_K \mid Z)$ is identified.

Define $\lambda_{\rho,t}(Z)$ by

$$\lambda_{\rho,t}(Z) = \int_{X \in supp(X)} E\left[h_\rho\left(\widetilde{Y}, t_1, ..., t_K\right) \mid X_1, X_2...X_K, Z\right] X_1^{t_1-1} X_2^{t_2-1}...X_K^{t_K-1} dX_1 dX_2...dX_K \tag{5}$$

$\lambda_{\rho,t}(Z)$ is an integral of a known conditional expectation, and so is identified. The proof of Theorem 2 works by showing that the moment $E\left(U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} \mid Z\right)$ is identified for integers $t_1,...,t_k$ by equaling the ratio of identified objects $\lambda_{\rho,t}(Z)/\kappa_{\rho,t}$. Identification of $F_{U|Z}(U_1,...,U_K \mid Z)$ then follows from identification of the distribution of $U_1^{-1} U_2^{-1}...U_K^{-1} \mid Z$ using its MGF. We could have instead directly identified $F_{U|Z}$ using it's MGF by working with negative values of $t_1,...,t_k$, or identified its characteristic function replacing $t_1,...,t_k$ with the square root of minu one times $t_1,...,t_k$ for reals $t_1,...,t_k$, but Assumption A6 would be more difficult to satisfy than with positive integers t.

3.3   Full Model Identification and Over Identification

Here we combine Theorems 1 and 2 to completely identify equation (2), and then show how the model is overidentified, and might be identified even without the use of Theorem 1.

COROLLARY 2: Let $Y = \sum_{k=1}^K G_k(X_k U_k) + U_0$. Define $G(X_1 U_1, ..., X_K U_K) = \sum_{k=1}^K G_k(X_k U_k)$. Let Assumptions A1, A2, A3, A4, and A6 hold. Then functions $G_1, G_2, ..., G_K$ and the joint distribution function $F_{U|Z}(U_1,...,U_K \mid Z)$ are identified.

Corollary 2 shows not just identification but substantial over identification of the model. In particular, both Theorem 1 and Theorem 2 identify the functions $F_{U_k|Z}$ for $k = 1, ..., K$, and so overidentifying restrictions on the functions $G_1, G_2, ..., G_K$ and $F_{U_1|Z}, ..., F_{U_K|Z}$ are obtained by equating the construction of the functions $F_{U_k|Z}$ from each of the two theorems for each k.

Additional restrictions are obtained by applying Theorem 2 using different h functions. Let P denote a set such that, for any $\rho \in P$, Assumption A6 holds for the function h_ρ. It then follows immediately from the proof of Theorem 2 that

$$\frac{\lambda_{\rho,t}(z)}{\kappa_{\rho,t}} = \frac{\lambda_{\widetilde{\rho},t}(z)}{\kappa_{\widetilde{\rho},t}} \quad \text{for all } t \in \mathbb{R}_+^K \text{ and all } \rho \in P, \widetilde{\rho} \in P. \tag{6}$$

Theorem 2 only needed and used positive integers t, but the same equations hold for reals t. For a given choice of t, h_ρ, and $h_{\widetilde{\rho}}$, equation (6) depends only on conditional expectations of data that can be readily estimated, and on the functions $G_1, ...G_K$. Equation (6) therefore provides a continuum of equations in the unknown functions $G_1, ...G_K$ for each pair of ρ and $\widetilde{\rho}$. For example, we might take h_ρ to be normal density functions with mean μ and variance σ^2, and then each value of ρ could correspond to a different (μ, σ^2) vector, providing an infinite number of such equations for an infinite number of choices of t.

If there exists only one set of functions $G_1, ...G_K$ that jointly satisfies equation (6) for all real vectors t and all $\rho, \widetilde{\rho} \in P$, (which is essentially a completeness assumption), then they suffice to completely $G_1, ...G_K$. In this case Theorem 1 is not needed at all, since this result combined with Theorem 2 would then identify the entire model.

We do not know of low level sufficient conditions to ensure that this collection of equations (6) has a unique solution, so we do not present this result alone, without Theorem 1, as a formal identification theorem for the functions $G_1, ...G_K$. However, this infinite set of restrictions shows that we have far more identifying information regarding $G_1, ...G_K$ and $F_{U_1|Z}, ..., F_{U_K|Z}$ than just the thin set based information from Theorem 1.

3.4   Models With Interaction Terms

The overidentifying information discussed in the previous subsection suggests that Theorems 1 and 2 may be combined to potentially identify richer models than equation (2). Suppose that

$$Y = \widetilde{G}\left(X_1 U_1, ..., X_K U_K, \xi\right) + \sum_{k=1}^{K} G_k\left(X_k U_k\right) + U_0 \tag{7}$$

where the function $\widetilde{G}$ is known, but the vector of parameters ξ is unknown, and $\widetilde{G}$ equals zero whenever all but one of its first K elements equals zero. Then the functions $G_1, G_2, ..., G_K$ are still identified by Corollary 1. We can then identify and construct the function $\kappa_{\rho,t}(\xi)$ defined by

$$\kappa_{\rho,t}(\xi) = \int_{supp(X)} h_\rho\left[\widetilde{G}\left(s_1, ..., s_K, \xi\right) + \sum_{k=1}^{K} G_k\left(s_k\right), t_1, ..., t_K\right] s_1^{t_1-1} s_2^{t_2-1}...s_K^{t_K-1} ds_1 ds_2...ds_K$$

and it follows from equation (6) that

$$\frac{\lambda_{\rho,t}(z)}{\kappa_{\rho,t}(\xi)} = \frac{\lambda_{\widetilde{\rho},t}(z)}{\kappa_{\widetilde{\rho},t}(\xi)} \quad \text{for all } t \in \mathbb{R}_+^K \text{ and all } \rho \in P, \widetilde{\rho} \in P. \tag{8}$$

This then provides infinitely many equations that can be potentially used to identify the finite parameter vector ξ. If ξ can be identified from these equations, then we can then apply Theorem 2 to identify $F_{U|Z}(U_1,...,U_K \mid Z)$, and so the entire model will be identified.

To illustrate, consider the model

$$Y = G_1\left(X_1 U_1\right) + G_2\left(X_2 U_2\right) + \xi X_1 U_1 X_2 U_2 + U_0 \tag{9}$$

for some unknown scalar constant ξ, so in this example $\widetilde{G}\left(s_1, s_2, \xi\right) = \xi s_1 s_2$. Apply Theorem 1 to identify G_1, G_2, and $F_{U_0|Z}\left(U_0 \mid Z\right)$. Assume there exists a value of $\rho, \widetilde{\rho}$, and t such that $\kappa_{\rho,t}(\xi) / \kappa_{\widetilde{\rho},t}(\xi)$ is strictly monotonic in the scalar ξ. This then suffices to ensure that equation (8) has a unique solution and hence that ξ is identified. Finally, apply Theorem 2 with $G\left(s_1 s_2\right) = \xi s_1 s_2 + G_1\left(s_1\right) + G_2\left(s_2\right)$ to identify $F_{U|Z}\left(U_1,U_2 \mid Z\right)$, and so the entire model given by equation (9) is then identified.

For either equation (2) or more general models like those above, constructing an estimator based on mimicking the steps of our identification arguments would likely be both inefficient

and difficult to implement. Inefficiency is likely because Theorem 1 uses thin set identification and Theorem 2 provides equations based on specific choices of the function h_ρ, and it is hard to see how one might choose the function h_ρ to maximize efficiency. Indeed, different h_ρ functions might be optimal for each moment and each function to be estimated. Also, Theorem 2 identifies the MGF of $U_1^{-1}, ...U_K^{-1}$, so an inversion would be needed to directly obtain the distribution function of U. Finally, sequentially applying Theorem 1 to estimate $G_1, G_2, ..., G_K$ and Theorem 2 to estimate F_U would ignore the overidentifying information discussed in the previous subsection. We will therefore use parametric maximum likelihood to estimate our models.

4   Barten Scales in Utility Functions

Let a "consumer" refer to an individual or household that maximizes a single well behaved utility function. Let Q_j denote the quantity purchased of a good j, and let $S(Q, U)$ denote the direct utility function over the bundle of goods $Q = (Q_1, ..., Q_J)$ of a consumer having a vector of preference heterogeneity parameters $U = (U_1, ..., U_J)$. Assume S is continuous, non-decreasing, twice differentiable in Q and quasi-concave in Q. Define the reference consumer to be a consumer that has heterogeneity parameters U normalized to equal one, and let $\overline{S}(Q_1, ..., Q_J)$ denote the direct utility function of a reference consumer. Each consumer chooses quantities to maximize utility subject to the standard linear budget constraint $\sum_{j=1}^{J} P_j Q_j = M$ where P_j is the price of good j and M is the total amount of money the consumer spends on this bundle of goods. Let $W_j^* = Q_j P_j / M$ be the share of the money budget M that is spent on good j (called the budget share of good j). Write the Marshallian budget share functions that result from maximizing the reference utility function $\overline{S}$ subject to the budget constraint as $W_j^* = \omega_j(P_1/M, ..., P_J/M)$. Let $V(P_1/M, ..., P_J/M)$ denote the indirect utility function corresponding to $\overline{S}$, obtained by substituting $Q_j = \omega_j(P_1/M, ..., P_J/M) M/P_j$ into $\overline{S}(Q_1, ..., Q_J)$ for $j = 1, ..., J$.

Our empirical application is based on Barten (1964) scales. Barten scales are a longstanding method used to bring preference heterogeneity on the basis of observed variables into continuous demand models. Barten scales are consequently a natural starting point for the incorporation of random utility parameters representing unobserved preference heterogeneity. See, e.g., Lewbel (1997) for a survey of various types of equivalence scales in the consumer demand literature, including Barten scales, and see Jorgenson, Lau, and Stoker (1982) for a prominent empirical application of traditional Barten scales. Deaton and Muellbauer (1980) includes an extensive discussion of parametric identification of Barten Scales.

Barten (1964) proposed the model in which consumers have utility functions of the form $S(Q_1, .., Q_J; \alpha_{h1}, ..., \alpha_{hJ}) = \overline{S}(Q_1/\alpha_{h1}, ..., Q_J/\alpha_{hJ})$, where the Barten scales $\alpha_{h1}, ..., \alpha_{hJ}$ are positive functions of observable household attributes h, such as age or family size, that embody variation in preferences across consumers. For households with multiple members, Barten scales can be interpreted as representing the degree to which each good is shared or jointly consumed. The smaller the Barten scale α_{hj} is, the greater the economies of scale to consumption of good j within the household. This is then reflected in the demand functions, where smaller Barten scales have the same effect on demands as lower prices. For example, if a

couple with one car rides together some of the time, then in terms of total distance each travels by car, sharing has the same effect as making gasoline cheaper. The more they drive together instead of alone, the lower is the effective cost of gasoline, and the smaller is the couple's Barten scale for gasoline.

More generally, Barten scales provide a measure of the degree to which different households get utility from different goods. This is how we will employ them. Although Barten scales have long been a popular method of modeling preference heterogeneity in empirical work, up until now Barten scales have always been modeled as deterministic functions of observable characteristics of consumers. Here we consider using Barten scales to embody unobserved heterogeneity of preferences across consumers.

We propose random Barten scales, assuming that consumers have utility functions of the form $S(Q_1, .., Q_J; U_1, ..., U_{hJ}) = \overline{S}(Q_1/U_1, ..., Q_J/U_J)$, where $U_1, ..., U_J$ are positive random utility parameters embodying preference heterogeneity (both observed and unobserved) across consumers. More formally, we could write each random Barten scale as $U_j(h)$, since for each good j, the distribution function that U_j is drawn from could depend on observable household attributes h. Barten's original model is then the special case where the distribution of each $U_j(h)$ is degenerate with a mass point at α_{hj}.

Define normalised prices $X_j = P_j/M$ for each good j and rewrite the budget constraint as $\sum_{j=1}^{J} X_j Q_j = 1$. Now $\overline{S}(Q_1, ..., Q_J)$ and $V(X_1, ..., X_J)$ are the direct and indirect utility functions of the reference consumer, and $\omega_j(X_1, ..., X_J)$ is the Marshallian budget share demand function of the reference consumer. It can be immediately verified from the first order conditions for utility maximization that a consumer will have Marshallian demand functions of the form $W_j^* = \omega_j(U_1 X_1, ..., U_J X_J)$ for each good j if and only if the consumer's direct and indirect utility function equal, up to an arbitrary monotonic transformation, $\overline{S}(Q_1/U_1, ..., Q_J/U_J)$ and $V(U_1 X_1, ..., U_J X_J)$, respectively. Also, given a specification of reference indirect utility $V(X_1, ..., X_J)$, the corresponding Barten scaled demand functions can be obtained by the logarithmic form of Roy's identity:

$$\omega_j(U_1 X_1, ..., U_J X_J) = \frac{\partial V(U_1 X_1, ..., U_J X_J)}{\partial \ln X_j} / \left(\sum_{\ell=1}^{J} \frac{\partial V(U_1 X_1, ..., U_J X_J)}{\partial \ln X_\ell} \right) \quad (10)$$

Notice that the functional form of each ω_j only depends on the functional form of $\overline{S}$ or equivalently of $\overline{V}$, so $U_1, ... U_J$ can vary independently of $X_1, ..., X_J$ across consumers. These derivations are exactly those given by Barten (1964) and by later authors who applied Barten scales, e.g., Jorgenson, Lau, and Stoker (1982), except that we put unobserved random variables U_j in place of deterministic functions α_{hj} of observed household characteristics. Random Barten scaled Marshallian demand functions then have precisely the form of our generalized random coefficients given in equation (1).

4.1 Indirectly Additively Separable Utility

In our empirical application, we let ω_1 be the budget share of a single good of interest, energy, and we let ω_2 denote the budget share of all other goods, corresponding to the general Barten scaled model with $J = 2$. This case only requires estimating a single equation for ω_1, since the

equation for ω_2 is automatically determined by construction as $\omega_2 = 1 - \omega_1$. If we had $J > 2$, then we would have $J - 1$ separate equations to estimate, and we would have further overidentification because the same Barten scales, with the same joint distribution $F_{U|Z}(U_1,...,U_K \mid Z)$, would appear in each equation.

Matzkin, (2007a), (2007b), (2008) discusses identification of systems of equations where the number of equations equals the number of random parameters, assuming it is possible to invert the reduced form of the system to express the random parameters as functions of observables. Although our model has J Barten scales U_j and J demand equations, Matzkin's identification method for systems of equations cannot be applied here because there are actually only $J - 1$ distinct demand functions $\omega_1,...,\omega_{J-1}$, with the remaining demand function ω_J determined by the adding up constraint that $\sum_{j=1}^{J} \omega_j = 1$.

The decomposition of total consumption into $J = 2$ goods is often done in empirical work when one wishes to focus on the welfare effects of a price change on a particular good, as we will do empirically. See, e.g., Hausman (1981), Hausman and Newey (1995), Blundell, Horowitz, and Parey (2010), and Hoderlein and Vanhems (2011). This construction is formally rationalizable by assuming utility is separable into good 1 and a subutility function of all other goods. See, e.g., Blackorby, Primont, and Russell (1978). Alternatively, Lewbel (1996) provides conditions on the distribution of prices (stochastic hicksian aggregation) under which Marshallian demand functions have the same properties with nonseparable utility as with separable utility.

With $J = 2$ goods, our model is $W_1^* = \omega_1(U_1 X_1, U_2 X_2)$ and $W_2^* = 1 - W_1^*$, and with $J = 2$ we can rewrite equation (10) as

$$\lambda\left(W_1^*\right) = \ln\left(\frac{\partial V\left(U_1 X_1, U_2 X_2\right)}{\partial \ln X_1}\right) - \ln\left(\frac{\partial V\left(U_1 X_1, U_2 X_2\right)}{\partial \ln X_2}\right) \tag{11}$$

where $\lambda\left(W_1^*\right)$ is the logit transformation function $\lambda\left(W_1^*\right) = \ln\left[W_1^* / \left(1 - W_1^*\right)\right]$.

Due to the constraints of Slutsky symmetry, imposing additivity directly on the Marshallian budget share function $\omega_1(X_1, X_2)$ would result in extreme restrictions on behavior. See, e.g., Blackorby, Primont, and Russell (1978). So we instead impose additivity on the logit transformation of $\omega_1(X_1, X_2)$ (later this will be relaxed to allow for interaction terms), thereby assuming demands have the additive form

$$\lambda\left(W_1\right) = \lambda\left[\omega_1\left(U_1 X_1, U_2 X_2\right)\right] + U_0 = g_1\left(U_1 X_1\right) + g_2\left(U_2 X_2\right) + U_0 \tag{12}$$

Here the functions g_1 and g_2 are nonparametric and U_0 is interpreted as measurement error in the observed budget share W_1 relative to the true budget share W_1^*. This implies that the underlying demand function for good 1 is given by

$$W_1^* = \omega_1\left(U_1 X_1, U_2 X_2\right) = \left(1 + e^{-g_1(U_1 X_1) - g_2(U_2 X_2)}\right)^{-1} \tag{13}$$

Use of the logit transformation here, and assumed additivity in logit transformed budget shares, has as far as we know not been considered before in the estimation of continuous demand functions. However, this logit transformed model has a number of advantages. First, $\lambda(W_1)$ has support on the whole real line, so the measurement error U_0 has unrestricted support, instead of a support that necessarily depends on covariates. Second, with this transform no constraints

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need to be placed on the range of values the nonparametric functions g_1 and g_2 take on. Third, unlike all other semiparametric or nonparametric applications of the Hausman (1981) consumer surplus type methodology (such as those cited above), a closed form expression for the indirect utility function that gives rise Marshallian demands (13) and hence (12) exists, and is given by Theorem 3.

THEOREM 3: The demand function ω_1 satisfies $\lambda\left[\omega_1\left(U_1 X_1, U_2 X_2\right)\right] = g_1\left(U_1 X_1\right) + g_2\left(U_2 X_2\right)$ for some functions g_1 and g_2 if and only if ω_1 is derived from an indirect utility function of the form

$$V\left(U_1 X_1, U_2 X_2\right) = H\left[h_1\left(U_1 X_1\right) + h_2\left(U_2 X_2\right), U_1, U_2\right].$$

for some montonic in its first element function H and some differentiable functions h_1 and h_2. The functions g_1, g_2, h_1, and h_2 are related by

$$h_1\left(U_1 X_1\right) + h_2\left(U_2 X_2\right) = \int e^{g_1\left(U_1 x_1\right)} d \ln x_1 + \int e^{-g_2\left(U_2 x_2\right)} d \ln x_2 \tag{14}$$

and

$$g_1\left(U_1 X_1\right) + g_2\left(U_2 X_2\right) = \ln\left(\frac{\partial h_1\left(U_1 X_1\right)}{\partial \ln X_1}\right) - \ln\left(\frac{\partial h_2\left(U_2 X_2\right)}{\partial \ln X_2}\right) \tag{15}$$

Also, the functions $h_1\left(U_1 P_1 / M\right)$ and $h_2\left(U_2 P_2 / M\right)$ are each nonincreasing, and their sum is strictly increasing in M and quasiconvex in P_1, P_2, and M.

The function H has no observable implications for individual consumer's demand functions, and is present only because utility functions are ordinal and therefore unchanged by monotonic transformations.[4] We can therefore just write the indirect utility function in Theorem 3 as

$$V\left(U_1 X_1, U_2 X_2\right)^{-1} = h_1\left(U_1 X_1\right) + h_2\left(U_2 X_2\right). \tag{16}$$

which takes H to be the reciprocal function (this is a convenient normalization since we later take h_1 and h_2 to be increasing functions, and utility must be nondecreasing in total expenditures).

Preferences $V\left(X_1, X_2\right)$ are defined to be indirectly additively separable (see, e.g., Blackorby, Primont, and Russell 1978) if, up to an arbitrary monotonic transformation H, $V\left(X_1, X_2\right) = H\left[h_1\left(X_1\right) + h_2\left(X_2\right)\right]$ for some functions h_1, h_2. So an equivalent way to state the first part of Theorem 3 is that ω_1 satisfies equation (13) if and only if preferences are given by a Barten scaled indirectly additively separable utility function. The second part of Theorem 3 then provides closed form expressions for the indirect utility function given the nonparametric (additive in the logit transformation) demand function and vice versa.

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[4]Later we will reintroduce the function H to construct a money metric representation of utility for use in social welfare calculations.

4.2 Random Barten Scales: Identification

From equation (12) we have the demand model

$$\lambda\left(W_1\right) = g_1\left(U_1 X_1\right) + g_2\left(U_2 X_2\right) + U_0 \tag{17}$$

Identification of this model can be obtained by Corollary 3, letting $Y = \lambda\left(W_1\right)$ and $G_k = g_k$. A condition that suffices to make the monotonicity of Assumption A3 hold is that the goods not be Giffen goods.[5] Having good 1 not be Giffen guarantees monotonicity of g_1, and similarly the restriction that good 2 is not Giffen means that ω_2 is monotonic in X_2, which by the adding up constraint $\omega_1 + \omega_2 = 1$ implies monotonicity of g_2.

Next consider Assumptions A1 and A2. As discussed earlier, Barten scales are traditionally modeled as deterministic functions of demographic characteristics, so in our extension to random Barten scales we take Z to be demographic characteristics. Continuity of each X_k and U_k is straightforward. U_0 is assumed to be measurement error in Y and hence independent of the other variables. U_1 and U_2 are preference parameters, and it is common to assume that tastes are determined independently of regressors in partial equilibrium analyses. In our application W_1 will be energy demand by Canadian households and the regressors are scaled prices. One might therefore be concerned about correlations between U and X caused by endogeneity of prices, however, Canadian households comprise a very small fraction of world energy demand, and so the likely effect of U on energy prices should be very small. In our empirical work we verify statistically that endogeneity, if any is present due to this or other sources (such as potential measurement error in total expenditures) appears too small to significantly change our results.

Each X_k is by construction nonnegative so to satisfy Assumption A4 we assume the support of each X_k is $\mathbb{R}^+$, the closure of which includes zero. We are therefore assuming that prices can be arbitrarily close to zero and/or total expenditures can be arbitrarily large. Given our overidentification results, we are not relying exclusively on data in this region for identification. Barten scales are positive, so for Assumption A4 we assume their distribution is both bounded and bounded away from zero to ensure the required MGF exists. Finally, the earlier discussions regarding Assumption A6 carry over directly to these demand applications.

4.3 Random Barten Scales: Specification and Estimation

Given nonparametric identification, we originally attempted to estimate the model using Sieve Maximum Likelihood as in Chen (2007). However, we found that, given our sample size, complexity of the model, and the dimensionality of our data, attempts to estimate more than second or third order expansions for the unknown functions in the model yielded results that were numerically unstable. We have therefore opted to primarily follow a parametric Maximum Likelihood modeling strategy, though using functional forms based on series expansions. We then follow these estimates with a battery of robustness checks and tests of model adequacy.

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[5]A Giffen good is a good that has a positive own price elasticity in its Marshallian quantity demand function, and hence an upward sloping demand curve. While possible in theory, very little empirical evidence has been found for the existence of Giffen goods, and particularly not for the types of goods we consider in our application. The only example we know of is Jensen and Miller (2008), who show that some grains may have been Giffen goods for extremely poor households in rural China.

Marshallian budget shares are commonly modeled as equal to, or proportional to, polynomials, almost always of third or lower order in terms of flexibility. See, e.g., Lewbel (2008) and references therein. We therefore specify the functions g_1 and g_2 as squared third-order polynomials

$$e^{g_1(X_1)} = \left(\beta_{10} + \beta_{11}X_1 + \beta_{12}X_1^2 + \beta_{13}X_1^3\right)^2 \tag{18}$$

$$e^{-g_2(X_2)} = \left(\beta_{20} + \beta_{21}X_2 + \beta_{22}X_2^2 + \beta_{23}X_2^3\right)^2 \tag{19}$$

with constants β_{ks} for $k = 1, 2$ and $s = 0, .., 3$. We square these polynomials, analogous to Gallant and Nychka (1987), to ensure that the resulting demand functions will not entail taking logs of a negative number. This specification also has the advantage that we can analytically evaluate the integrals that define the corresponding indirect utility function in Theorem 3. Specifically, by equation (14) we get $V^{-1} = h_1(U_1X_1) + h_2(U_2X_2)$ where[6]

$$h_k(X_k) = \int_{\ln X_k} \left(\beta_{k0} + \beta_{k1}e^r + \beta_{k2}e^{2r} + \beta_{k3}e^{3r}\right)^2 dr \tag{20}$$

As noted earlier, it is both unusual and convenient to have closed form expressions for utility functions corresponding to arbitrary demand function components like these.

We impose the free normalization $\beta_{20} = 1$. This is imposed without loss of generality, because if $\beta_{20} \neq 0$ then we can rescale the indirect utility function and redefine the remaining β_{ks} parameters appropriately to get an observationally equivalent representation of indirect utility that has $\beta_{20} = 1$. This, and assuming $E(U_0) = 0$, are free normalizations that take the place of the normalizations of the g functions described in Theorem 1. These observationally equivalent normalizations are used in place of the ones used to prove Theorem 1, because they're more natural and easier to impose in our particular application.

Applying Theorem 3 and equation (12) to this model gives the demand function

$$\begin{aligned}
\lambda(W_1) &= \lambda\left[\omega_1\left(U_1X_1, U_2X_2, \beta\right)\right] + U_0 \tag{21} \\
&= \ln\left[\left(\sum_{s=0}^{3}\beta_{1s}(U_1X_1)^s\right)^2\right] - \ln\left[\left(\sum_{s=0}^{3}\beta_{2s}(U_2X_2)^s\right)^2\right] + U_0. \tag{22}
\end{aligned}$$

We also tried estimating the models with higher-order polynomials (see the Appendix) but found that, as noted above, the resulting parameter estimates became numerically highly unstable.

We next need to specify the distribution of U. To simplify the model, instead of letting the dependence of U on a vector of demographic characteristics Z be entirely unrestricted, we assume each Barten scale takes the form

$$U_k = \alpha_k(Z)\,\widetilde{U}_k \tag{23}$$

where $\alpha_k(Z)$ is a traditional deterministic Barten scale that depends on demographic household characteristics Z, and the remaining random variation $\widetilde{U}_k$ in each Barten scale is assumed to be

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[6]The integral in equation (20) is readily evaluated, e.g., for our squared cubic functions we have
$\int \left(\beta_{k0} + \beta_{k1}e^r + \beta_{k2}e^{2r} + \beta_{k3}e^{3r}\right)^2 dr =$
$\frac{1}{6}\beta_{k3}^2 e^{6r} + \frac{2}{5}\beta_{k2}\beta_{k3}e^{5r} + \frac{1}{4}\left(2\beta_{k1}\beta_{k3} + \beta_{k2}^2\right)e^{4r} + \frac{2}{3}\left(\beta_{k0}\beta_{k3} + \beta_{k1}\beta_{k2}\right)e^{3r} + \frac{1}{2}\left(2\beta_{k0}\beta_{k2} + \beta_{k1}^2\right)e^{2r} + 2\beta_{k0}\beta_{k1}e^r + r\beta_{k0}^2$ Each $h_k(U_kX_k)$ function is given by substituting $r = \ln(U_kX_k)$ into this expression.

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independent of covariates. We model $\ln[\alpha_k(Z)]$ as linear in a vector of demographic charac-teristics Z, having a vector of coefficients denoted θ_k. This index does not include a constant term, because the scaling of $\alpha_k(Z)$ is freely absorbed into the β_{ks} parameters.

For our main results we specify f_0, the density of U_0, as a mean zero normal with variance σ_0^2. We specify the joint distribution of the random component of the Barten scales, $\widetilde{U} = (\widetilde{U}_1, \widetilde{U}_2)$, to be a log normal, specifically, the density of $\ln \widetilde{U}$ is

$$f_{\ln \widetilde{U}}\left(\widetilde{U}_1, \widetilde{U}_2, \sigma, \rho\right) = \frac{1}{2\pi\sigma_1\sigma_2\left(1-\rho^2\right)^{1/2}} \exp\left(\frac{\left(\frac{\ln \widetilde{U}_1}{\sigma_1}\right)^2 - 2\rho\left(\frac{\ln \widetilde{U}_2}{\sigma_2}\right)\left(\frac{\ln \widetilde{U}_1}{\sigma_1}\right) + \left(\frac{\ln \widetilde{U}_2}{\sigma_2}\right)^2}{-2\left(1-\rho^2\right)}\right).$$
(24)

As our estimation method uses numerical integration, in the empirical work we use a finite support for the integration over $\widetilde{U}$. We integrate $\ln \widetilde{U}$ over the box defined by $\pm 3\sigma_1, \pm 3\sigma_2$ with a stepsize of $0.06\sigma_j$, yielding a 10,000 point grid for numerical integration.[7]

To check the adequacy of these parametric error distribution assumptions, In an appendix we later consider more general specifications in which these densities are multiplied by polynomials as in Gallant and Nychka (1987), corresponding to the types of hermite polynomial expansions that are commonly used nonparametric sieve density specifications. As we show later, we find that far richer specifications of these densities, with two or three times as many parameters as the above, yields estimates that are numerically close to the above normal and log normal models.

For a given consumer with observed values x_1 and x_2, the conditional density function of W_1 is then given by

$$f_{W_1|X_1,X_2,Z}\left(w_1 \mid x_1, x_2, z; \alpha, \beta, \sigma, \rho\right) \tag{25}$$
$$= \int_0^\infty \int_0^\infty f_0\left[\lambda\left(W_1\right) - \lambda\left[\omega_1\left(\alpha_1(z)\widetilde{u}_1 x_1, \alpha_2(z)\widetilde{u}_2 x_2, \beta\right)\right], \sigma_0\right] f_{\widetilde{U}}\left(\widetilde{u}_1, \widetilde{u}_2, \sigma, \rho\right) d\widetilde{u}_1 d\widetilde{u}_2$$

which for the above specification of preferences is

$$f_{W_1|X_1,X_2,Z}\left(w_1 \mid x_1, x_2, z; \alpha, \beta, \sigma, \rho\right) =$$
(26)
$$\int_{-\infty}^\infty \int_{-\infty}^\infty \frac{f_{\ln \widetilde{U}}\left(\widetilde{u}_1, \widetilde{u}_2, \sigma, \rho\right)}{(2\pi)^{1/2}\sigma_0} \exp\left(\frac{-1}{2\sigma_0^2}\left[\lambda\left(W_1\right) - \ln\left(\left(\frac{\sum_{s=0}^3 \beta_{1s}\left(\widetilde{u}_1\alpha_1(z)x_1\right)^s}{\sum_{s=0}^3 \beta_{2s}\left(\widetilde{u}_2\alpha_2(z)x_2\right)^s}\right)^2\right)\right]^2\right) d\ln\widetilde{u}_1 d\ln\widetilde{u}_2.$$

Assuming n independently, identically distributed observations w_{1i}, x_{1i}, x_{2i} of consuming households i, estimation proceeds by searching over parameter vectors α, β, σ, and ρ to maxi-mize the log likelihood function

$$\sum_{i=1}^n \ln f_{W_1|X_1,X_2,Z}\left(w_{1i} \mid x_{1i}, x_{2i}, z_i; \alpha, \beta, \sigma, \rho\right). \tag{27}$$

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[7]Bounding the support of $\widetilde{U}$ in this way satisfies the assumptions of Theorem 2 that $\widetilde{U}$ be bounded away from zero and ensuring that a moment generating function exists.

5   Empirical Results

5.1   Data

We estimate a baseline parametric specification as above, and with it undertake Engel curve, cost-of-living and social welfare analyses. Later, in an Appendix, we consider the impact of possible misspecification by undertaking a series of robustness checks. These include consideration of interaction terms in the utility function (relaxing additivity in functions of X_j), higher order polynomial demand functions, allowing for non-normal error terms with sieve-like hermite series density expansions, checking impacts of remaining heteroskedasticity, and controlling for possible endogeneity in X_j.

We estimate the model using Canadian household expenditure microdata from the 1997 to 2008 Surveys of Household Spending. We consider households comprised of one adult (as of 31 Dec) aged 25-64 residing in cities of 30,000 or more residents in provinces other than Prince Edward Island (due to data masking). We drop observations whose expenditures on energy goods are zero, and eliminate a few extreme outliers by removing those whose total nondurable expenditures are in the top or bottom percentile of the total nondurable expenditure distribution. This leaves 9971 observations for estimation.

We consider the budget share of energy goods, W_1, defined as the share of total nondurable expenditures devoted to energy goods. Total nondurable expenditures are constructed as the sum of household spending on food, clothing, health care, alcohol and tobacco, public transportation, private transportation operation, and personal care, plus the energy goods defined as fuel oil, electricity, natural gas and gasoline (reported in thousands of dollars). We include eight demographic characteristics, comprising the vector Z, as observed preference shifters: a dummy for female individuals; age of the individual (on an 8 unit integer scale for 5 year age groups with age 40 to 44 coded as 0); calendar year minus 2002; a dummy for residence in the francophone province of Quebec; Environment Canada ex poste records of the number of days requiring heating and cooling in each province in each year (normalized as z-scores from the full sample of all households in all provinces in all city sizes); an indicator that the household is a renter (spending more than $100 on rent in the year); and an indicator that the household received more than 10% of its gross income from government transfers. These demographic characteristics equal zero for the reference consumer (whose utility function is $\overline{S}$ and indirect utility function is V): a single male aged 40-44 with less than 10% transfer income living in owned accomodation outside Quebec in 2002 with average heating and cooling days.

Table 1: Summary Statistics

9971 Observations	mean	std dev	min	max
logit energy share, Y	-1.949	0.766	-7.140	1.005
energy share, W_1	0.146	0.085	0.001	0.732
nondurable expenditure, M	15.661	7.104	2.064	41.245
energy price, P_1	1.039	0.230	0.426	1.896
non-energy price, P_2	0.965	0.075	0.755	1.284
energy normalized price, X_1	0.082	0.049	0.015	0.570
non-energy normalized price, X_2	0.077	0.045	0.020	0.476
female indicator	0.482	0.500	0.000	1.000
age group-4	0.549	2.262	-3.000	4.000
year-2002	0.363	3.339	-5.000	6.000
Quebec resident	0.168	0.374	0.000	1.000
heat days, normalized	-0.102	0.990	-2.507	2.253
cooling days, normalized	0.014	1.007	-1.729	4.013
renter indicator	0.512	0.500	0.000	1.000
transfer income indicator	0.184	0.387	0.000	1.000

Prices vary by province (9 included) and year (12 years) yielding 108 distinct price vectors for the underlying commodities comprising nondurable consumption. These underlying commodity prices are normalised to equal one in Ontario in 2002. To account for the impact on prices of individual variation in compositional differences of these aggregate commodities, we follow the methodology of Lewbel (1989) and Hoderlein and Mihaleva (2008) in constructing P_1 as the Stone price index using within group household specific budget shares of energy goods, and P_2 is constructed similarly for non-energy goods. This construction has the feature of further increasing relative price variation across households at the aggregate group level. These price indices are normalized (a free normalization) to take the value of one in Ontario in 2002. The budget, M, is equal to the total nondurable expenditures of the household. The regressors X_j are then given by $X_j = P_j/M$. Finally, the regressand, Y, is the logit transformation of the energy budget share, so $Y = \lambda(W_1)$. Table 1 gives summary statistics for these budget shares, expenditures, prices, normalised prices and demographic preference (Barten scale) shifters.

5.2 Baseline: Random Barten Scales in Indirectly Additively Separable Model

Our main analyses are based on two models. The first, Model 1, imposes the restriction that $\widetilde{U}_1 = \widetilde{U}_2 = 1$, which makes the distribution of $\widetilde{U}$ degenerate, and thereby makes equation (25) simplify to

$$f_{W_1|X_1,X_2,Z}(w_1 \mid x_1, x_2, z; \alpha, \beta, \sigma_0) = \frac{\exp\left(\frac{-1}{2\sigma_0^2}\left[\lambda(W_1) - \ln\left(\left(\frac{\sum_{s=0}^{3}\beta_{1s}(\widetilde{u}_1\alpha_1(z)x_1)^s}{\sum_{s=0}^{3}\beta_{2s}(\widetilde{u}_2\alpha_2(z)x_2)^s}\right)^2\right)\right]^2\right)}{(2\pi)^{1/2}\sigma_0}.$$

(28)

This is just a traditional deterministic Barten scale model, having $U_j = \alpha_j(z)$, estimated using our general functional form for energy demand. Model 1 is then compared to our real specification, Model 2, which is equation (26) with the distribution of $\widetilde{U}$ given by equation (24), and therefore contains our random Barten scales $U_j = \alpha_j(z)\widetilde{U}_j$. Both models are estimated using maximum likelihood in Stata, with likelihood functions given by substituting equation (28) or (26) into equation (27). Estimated coefficients are given in Table 2 below.

Our estimates do not impose the theoretical restriction that the g functions be monotonic.[8] Despite not imposing monotonicity, we find that our Model 2 estimates satisfy monotonicity at 97.7% of the data points in our sample.

Model 2 has three more parameters than Model 1 (σ_1, σ_2, ρ). The likelihood ratio test statistic for the restriction that these parameters are all zero is 672, so the parameters that allow for random Barten scales parameters are highly jointly significant. One can see in Table 2 that they are also individually highly significant.

Figures 1 and 2 show the estimated joint distribution of $\ln\alpha_1(z)$ and $\ln\alpha_2(z)$ (logged deterministic portion of the Barten scales) in Model 1 and Model 2, respectively. Summary statistics for these distributions are provided in the bottom panel of Table 2. In both Models the estimated distributions of $\ln\alpha_1(z)$, $\ln\alpha_2(z)$ are bimodal. The two modes are driven almost entirely by the renter variable; conditioning on just renters or just owners produces unimodal distributions. In Canada, most renters do not pay for their own home heating or electricity (this is included in rents and doesn't depend on usage), causing a ceteris paribus reduction in their energy shares relative to home owners.

Looking at the bottom of Table 2, we see in both models that the standard deviation of $\ln\alpha_2(z)$ is much larger than that of $\ln\alpha_1(z)$. This indicates that heterogeneity in preferences due to observables is larger for non-energy than for energy goods. This is not surprising, it just says that people vary more (based on observable characteristics) in their taste for non-energy goods than in their taste for energy goods. Unconditionally, $\ln\alpha_1(z)$, $\ln\alpha_2(z)$ are slightly negatively correlated in both models. However, conditional on rental tenure, the deterministic components of Barten scales are strongly positively correlated.

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[8]In a parametric setting like ours, failing to impose inequality constraints on estimation that are satisfied by the true model does not affect standard limiting distribution theory, assuming that the true parameter values do not on the boundary of the parameter space.

Table 2: Estimated Parameters

Parameter		Model 1 llf=-10043.1 Estimate	*Std Err*	Model 2 llf=-9706.9 Estimate	*Std Err*
β_{10}		0.145	*0.010*	0.185	*0.007*
β_{11}		8.113	*0.487*	7.623	*0.287*
β_{12}		-37.563	*2.924*	-32.871	*2.147*
β_{13}		51.576	*5.650*	40.630	*4.390*
β_{21}		2.484	*0.568*	1.805	*0.266*
β_{22}		-1.743	*0.663*	1.053	*0.314*
β_{23}		0.152	*0.141*	-0.996	*0.139*
α_1	female	-0.214	*0.031*	-0.228	*0.015*
	agegp	0.002	*0.009*	0.013	*0.004*
	time	-0.013	*0.004*	-0.003	*0.002*
	PQ	0.085	*0.043*	0.043	*0.021*
	heat	0.036	*0.016*	0.026	*0.008*
	cool	-0.062	*0.015*	-0.035	*0.007*
	renter	-0.292	*0.058*	-0.440	*0.026*
	social	0.034	*0.038*	0.054	*0.020*
α_2	female	-0.130	*0.076*	-0.117	*0.010*
	agegp	-0.068	*0.023*	-0.038	*0.002*
	time	0.018	*0.010*	0.044	*0.001*
	PQ	0.402	*0.100*	0.217	*0.017*
	heat	0.015	*0.040*	-0.021	*0.008*
	cool	-0.077	*0.043*	-0.014	*0.006*
	renter	0.943	*0.155*	0.605	*0.008*
	social	-0.085	*0.091*	-0.110	*0.011*
σ_0		0.663	*0.005*	0.469	*0.009*
σ_1				0.165	*0.036*
σ_2				1.336	*0.011*
ρ				0.883	*0.100*
std dev	$\ln(\alpha_1)$	0.197		0.252	
	$\ln(\alpha_2)$	0.568		0.380	
correlation	$\ln(\alpha_1), \ln(\alpha_2)$	-0.479		-0.700	
(all obs)	$\ln U_1, \ln U_2$			0.293	
correlation	$\ln(\alpha_1), \ln(\alpha_2)$	0.426		0.105	
(renter=0)	$\ln U_1, \ln U_2$			0.699	
correlation	$\ln(\alpha_1), \ln(\alpha_2)$	0.420		0.087	
(renter=1)	$\ln U_1, \ln U_2$			0.691	

In Model 1 the log Barten scales equal $\ln \alpha_j(z)$, but in Model 2 the log Barten scales are given by $\ln U_j = \ln \alpha_j(z) + \ln \widetilde{U}_j$. The components $\ln \alpha_j(z)$ and $\ln \widetilde{U}_j$ are, respectively, the observed deterministic and unobserved random components of these Barten scales. Thus the mean, variance, and correlations of the $\ln \widetilde{U}_j$ terms in Model 2 are directly comparable to the

corresponding statistics of the $\ln \alpha_j(z)$ terms. The estimated parameters of the distribution of $\ln \widetilde{U}_j$ show some similar features to that of $\ln \alpha_j$. The standard deviation $\ln \widetilde{U}_2$ (σ_2) is much larger than that of $\ln \widetilde{U}_1$ (σ_1), so both observed and unobserved Barten scales vary more across consumers for nonenergy goods than for energy goods. Likewise, $\ln \widetilde{U}_1$ and $\ln \widetilde{U}_2$ are positively correlated (like $\ln \alpha_1$ and $\ln \alpha_2$ after conditioning on the rental dummy). That the estimated effects of the unobserved random components of Barten scales have similar patterns to the estimated effects of observed preference shifter on Barten scales is a reasuring indicator of the sensibility and reasonableness of our model. Overall, we find that unobserved preference heterogeneity is about important as observed preference heterogeneity in driving variation in Barten scales. The estimated joint distribution of $\ln U_1$, $\ln U_2$, summing the effects of $\ln \alpha_j(z)$ and $\ln \widetilde{U}_j$, is shown in Figure 3. Comparing Figures 1 and 3 shows that accounting for unobserved heterogeneity substantially increases the estimated total heterogeneity in tastes across individuals.

The unobserved preference heterogeneity terms $\widetilde{U}_j$ partly pick up unobserved variation that would otherwise be subsumed by the non behavioral error term U_0, making the estimated standard deviation of U_0 fall from 0.666 in Model 1 to 0.469 in Model 2. But more significantly, U_1 and U_2 also pick up a substantial portion of what would otherwise be unexplained heteroskedasticity in demand. This can be seen in a variety of ways. For example, appropriately modeling the heteroskedasticity driven by unobserved preference heterogeneity should yield increased precision in parameter estimates, just as appropriate generalized least squares estimation usually reduces standard errors relative to ordinary least squares estimation in regression models. Empirically, we do see an improvement in estimation precision, comparing across the columns in Table 2. The parameter estimates in Model 2 generally have standard errors about 20 per cent smaller than those of model 1. Other evidence that random Barten scales explain a substantial portion of both the size and systematic components of otherwise unexplained variability in demand (by making U_0 both smaller and more homoskedastic) is provided in the Appendix.

Table 3 gives summary statistics on predicted values of the logit transformed budget share Y and of the budget share itself, W_1. To show the impact of various model components on these fits, we first evaluate predicted Y and W_1 at reference level prices $P = 1$, reference level random Barten scale components $\widetilde{U}_j = 1$, average demographics $\overline{\alpha}_j$ and observed budgets M. The result, given in the first row of Table 3, shows the estimated impact of just varying total expenditures M. Then, in the succeeding rows of Table 3, we add back observed variation first due to observed demographics, then to prices, and finally in the unobserved preference heterogeneity terms $\widetilde{U}_j$.

Since Model 1 has just a single additively separable error term, the average prediction from Model 1 including variation from all regressors (M, α_j, P) exactly equals the mean of the observed Y (-1.949). The predicted average mean of Y in Model 2 is somewhat larger for the cases where we don't account for unobserved preference heterogeneity $\widetilde{U}_j$, with Y averaging about -1.78. But, when we account for unobserved preference heterogeneity (in the bottom row), Model 2 predicts the average -1.997 which is very close to the mean of the observed Y.

Looking across rows in Table 3 shws that the predictions of Models 1 and 2 have about the same standard deviation when $\widetilde{U}_j$ is ignored, but the standard deviation of predicted Y in Model 2 more than doubles to 0.896 when variation in $\widetilde{U}_j$ is accounted for. Predicted energy budget shares have a similar property. When including all sources of variation, the standard

deviation of Model 1 predictions of W_1 is 0.044, while that of Model 2 prediction is 0.065. Comparing these Table 3 predictions to Table 1 shows that, for both Y and W_1, Model 2 gives closer predictions to the actual empirical standard deviation of these variables than does Model 1.

Table 3: Estimated Energy Budget Shares				
Logit Budget Shares Y	Model 1		Model 2	
	Mean	Std Dev	Mean	Std Dev
$M, \alpha_j = \overline{\alpha}_j, P = 1, \widetilde{U}_j = 1$	-1.943	0.251	-1.779	0.250
$M, \alpha_j, P = 1, \widetilde{U}_j = 1$	-1.978	0.397	-1.799	0.379
$M, \alpha_j, P, \widetilde{U}_j = 1$	-1.949	0.387	-1.775	0.371
$M, \alpha_j, P, \widetilde{U}_j$			-1.997	0.896
Budget Shares W_1				
$M, \alpha_j = \overline{\alpha}_j, P = 1, \widetilde{U}_j = 1$	0.128	0.026	0.147	0.030
$M, \alpha_j, P = 1, \widetilde{U}_j = 1$	0.128	0.044	0.148	0.048
$M, \alpha_j, P, \widetilde{U}_j = 1$	0.131	0.044	0.151	0.047
$M, \alpha_j, P, \widetilde{U}_j$			0.137	0.065

Figure 4 shows estimated Engel curves from the models, showing W_1 as a function of $\ln M$. These are estimated demand functions evaluated at the reference price vector $P_1 = P_2 = 1$ and at average demographics $\overline{\alpha}_j$. Model implies a different Engel curve for every value that $\widetilde{U}_1$ and $\widetilde{U}_2$ can take on. The single Engel curve for Model 1, which imposes $\widetilde{U}_j = 1$, is shown as a thick gray line, while that for Model 2 evaluated at $\widetilde{U}_j = 1$ is shown as a thick black line. To illustrate the range of Engel curves implied by our model, we also evaluate Model 2 at each quartile of the distribution of $\widetilde{U}_1$ paired with each quartile of the distribution of $\widetilde{U}_2$, for a total of nine pairs of values. This yields eight Engel curves (in addition to the one thick black line), which are depicted by thin gray lines in Figure 4. Finally, we show the estimated marginal density of $\ln M$ (divided by 10 to fit in the graph) as a thick light gray line at the bottom of Figure 4.

On average, richer consumers tend to spend a smaller fraction of their budget on energy goods than poorer consumers. This can be seen in the mostly downward slope of the Engel curves in Figure 4. Comparing these curves to the depicted density function of $\ln M$ shows that only a small fraction of all consumers, the poorest ones, are on the upward sloping parts of these curves.

A striking feature of Figure 4 is that for each given value of M, there is substantial variation in the level and slope of Engel curves, due entirely to variation in $\widetilde{U}_j$. To reduce clutter we did not include standard error bars on this graph, but the differences between these estimated Engel curves are statistically significant. For example, at the mean value of $\ln M$ ($\ln M = 2.64$), the top Engel curve displayed is that of the top quartile of both $\widetilde{U}_1$ and $\widetilde{U}_2$, and the bottom Engel curve is that of the bottom quartile of both $\widetilde{U}_1$ and $\widetilde{U}_2$. The estimated levels of these Engel curves at $\ln M = 2.64$ are 0.166 and 0.104, respectively, with standard errors of 0.002 and 0.008, repectively. The sample value of the z-test statistic testing the hypothesis that these levels are the same is 7.4, so the levels are both economically and statistically significantly different from each other. The estimated derivatives of the budget shares with respect to $\ln M$ at this level of $\ln M$ are -0.116 and -0.017, respectively, and the sample value of z-test statistic for the hypothesis that these slopes are the same is 11.4, strongly rejecting this hypothesis.

Thus variation in the random components $\widetilde{U}_j$ of Barten scales, corresponding to unobserved variation in tastes across consumers, yields significant differences in both the levels and slopes of the estimated Engel curves.

Figure 4 showed the effects on W_1 of just the random component of the Barten scales at all different budget levels M. In contrast, Figure 5 illustrates the total effect of Barten scales on W_1. Specifically, Figure 5 shows a contour plot of the joint distribution of W_1 (predicted by Model 2) and observed M, evaluated at reference prices $P_1 = P_2 = 1$ and observed demographics z. The vertical variation in this graph therefore shows the estimated variation in W_1 due to Barten scales (both observed and unobserved components) at different M levels. This wide variation in tastes will have important implications for welfare analyses below.

Taken together, all of the above results show that Model 2's inclusion of random Barten scales accounts for more and richer variation in observed behaviour than does Model 1. This is due to the fact that budget shares are highly variable and heteroskedastic, and Model 1 treats this variance and heteroskedasticity entirely as meaningless variation in a nonbehavioral error term, while Model 2 captures much of this variation in a behaviorally sensible, structural way, via random Barten scales.

The next subsections show the price implications of our estimates and provide welfare calculations, based primarily on Model 2. Additional checks of the adequacy of the specification of Model 2 and robustness of our results against a range of more general alternatives is provided in the Appendix

5.3 Consumer Surplus Effects of a Carbon Tax

We now apply our model to evaluate the partial equilibrium effects of a large change in the price of energy, as might result from a carbon tax.[9] Using equation (35), we have a closed form expression for indirect utility. We can therefore directly compute consumer surplus effects, just by numerically inverting the indirect utility function (35) to obtain the cost of living impact of a price change. Without Theorem 3, we would otherwise require numerical approximation instead of exact solutions such as in Vartia (1984), or we need to numerically solve a differential equation as in Hausman and Newey (1995), but such a solution would need to be calculated for every value on the continuum of points that U_1 and U_2 can take on.

Recall the indirect utility function defined over normalized prices X_j and Barten scales U_j is $V(U_1 X_1, U_2 X_2) = V(U_1 P_1/M, U_2 P_2/M)$. For an individual facing initial prices $\overline{P}_1, \overline{P}_2$, having total expenditures M, and having preferences indexed by Barten scales U_1, U_2, the cost-of-living impact of moving to new prices P_1, P_2 is $\pi\left(U_1, U_2, M, P_1, P_2, \overline{P}_1, \overline{P}_2\right)$, defined as the solution to

$$V(U_1 X_1, U_2 X_2) = V\left(\frac{U_1 \overline{P}_1}{M}, \frac{U_2 \overline{P}_2}{M}\right) = V\left(\frac{U_1 P_1}{\pi M}, \frac{U_2 P_2}{\pi M}\right).$$

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[9]Our model is not a general equilibrium model, so we are only estimating the consumer's responses to a change in energy prices. Moreover, these should only be interpreted as short run responses, since in the longer run consumers could change their energy elasticities and demand by, e.g., buying more energy efficient cars and appliances. Also, we just consider a change in the overall price of energy, and so do not consider impacts of possible changes in the composition of energy goods.

Here π is the proportionate change in costs M needed to compensate for the price change, that is, the amount by which M would need to be scaled up to bring an individual facing prices P_1, P_2 (and having preferences given by U_1, U_2) back to the same indifference curve they were on when facing prices $\overline{P}_1$, $\overline{P}_2$.

The cost-of-living index π should be weakly positive for a price increase, reflecting the fact utility is decreasing in normalized prices X_j and that the Slutsky matrix is symmetric and negative semidefinite. Our model and associated estimator imposes all of these conditions except negative semidefiniteness. However, this condition turns out to be satisfied at all but six of our data points in Model 2 (the condition is violated at 36 points in Model 1).

To show price effects clearly, we consider a large price change: a 50% increase in the price of energy. This price increase is chosen to approximate the effect of a \$300 per ton CO2 tax (see, e.g., Rhodes and Jaccard 2014)[10]. We solve for the π function given the initial price vector $\overline{P}_1 = \overline{P}_2 = 1$ and the new price vector $P_1 = 1.5$, $P_2 = 1$. Figure 6 shows the resulting estimated joint distribution (contour plot) of $\ln\pi$ and $\ln M$ from Model 2 evaluated at the observed demographics Z and budgets M. This plot is constructed by calculating the surplus for each observation in the data, with draws from the estimated distribution of $\widetilde{U}_1$, $\widetilde{U}_2$, and using observed values of the preference shifters Z.

Analogous to Table 3, to assess the contribution of variation in π due to observed and unobserved preference heterogeneity, we also calculate the π distribution imposing $\widetilde{U}_j = 1$ and imposing $\alpha_j = \overline{\alpha}_j$, and do so for both Models 1 and 2. Table 4 gives summary statistics (means and standard deviations) of these distributions. For ease of presentation this Table reports per cent changes in cost of living, that is, $100(\pi - 1)$. Standard errors for these statistics that account for the sampling variability of the parameter estimates (estimated via simulation) are provided in *italics*.

Table 4: Cost of Living Impacts: 50% Energy Price Increase					
Per Cent Increase		Model 1		Model 2	
$\pi - 1$, per cent		Estimate	*Std Err*	Estimate	*Std Err*
$\alpha_j = \overline{\alpha}_j, \widetilde{U}_j = 1$	Mean	5.34	*0.22*	5.66	*0.17*
	Std Dev	1.26	*0.06*	1.30	*0.05*
$\alpha_j, \widetilde{U}_j = 1$	Mean	5.31	*0.24*	5.64	*0.17*
	Std Dev	1.85	*0.21*	1.69	*0.08*
$\alpha_j, \widetilde{U}_j$	Mean			5.37	*0.20*
	Std Dev			4.31	*0.46*

It has long been known that first order approximations to the cost of living effects of marginal price changes can be evaluated without estimating demand functions, essentially by ignoring substitution effects (see, e.g., Stern 1987). These theoretical results have been used to argue that demand function estimation is not required for marginal policy analyses. In our data, the average value of the budget share for energy is 0.146, so if there were no substitution effects in response to a price change, doubling the price of energy would increase the cost of living by 7.3 per cent. This would be the first order approximation based estimate of π. However,

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[10]British Columbia has a CO2 tax. It charges 6.7 cent/liter of gasoline for 30\$/ton. A 67 cent/liter CO2 tax is about half as large as the pump price of gasoline in 2002 (the base year for this analysis).

the estimated cost-of-living impacts given in Table 4, averaging about 5.4 per cent, are much smaller than 7.3 per cent, showing substantial relative price substitution effects. This difference of nearly 2 percentage points is very large relative to the standard error of estimated means in Models 1 and 2 of about 0.20 percentage points, so the hypothesis that the model estimates have a mean of 7.3 is strongly rejected at conventional levels. These results supports findings in, e.g., Banks, Blundell, and Lewbel (1996) that, contrary to the first order approximation theory, it is empirically necessary to estimate demand functions and associated price elasticities to properly evaluate the consumer surplus and welfare effects of large price changes. Moreover, one goal of an energy tax would be to reduce energy consumption (a substitution effect), so it's important to account for the impact on welfare of this reduction.

Models 1 and 2, with or without variation in α_j or $\widetilde{U}_j$, deliver similar estimates of the mean effects of the energy tax on cost of living. However, as can be seen by going down the columns of Table 4, each source of variation increases the estimated standard deviation of π, and in particular the inclusion of the random Barten scale components $\widetilde{U}_j$ in Model 2 more than doubles the estimated standard deviation of π across consumers (this difference is statistically significant as well as being economically large; the z-test statistic for the hypothesis that these standard deviations are the same has value of 6.1). Though less substantial economically, the difference in mean effects between the Model 2 estimates without and with unobserved preference heterogeneity (5.64% and 5.37%, respectively) is also statistically significant with a z-test statistic of 3.9.

The large estimated standard deviation of π in Model 2 (which is mostly due to unobserved variation in tastes $\widetilde{U}_j$) has substantial welfare implications. The larger is the variation in π, the larger is the variation in impacts of an energy tax. Although the average consumer would need to have their budget M increased by 5.37% to compensate for the tax, some consumers (those near the bottom of Figure 6) would only need a slight increase in their budget to be made whole, while others (those near the top of Figure 6) would be greatly harmed by the tax, needing an over 10% increase in M to compensate.

What makes this substantial variation in cost of living impacts particularly relevant economically is that it mostly impacts poorer consumers. As can be seen in Figure 6, both the mean and the variation in cost of living impacts is larger at low values of M than at high values, so those consumers who are hurt the most by the tax in percentage terms are also predominantly the poorer consumers, who can least afford the increase in costs. Not only do richer consumers spend a smaller fraction of their budget on energy goods (as seen in Figure 4), but they also appear to have a greater ability to substitute away from energy when the relative price of energy increases.

5.4 Social Welfare Implications of a Carbon Tax

The above analysis showed the distribution across consumers of the cost of living effects of an energy tax. We now evaluate the implications of these results for aggregate welfare, based on a range of social welfare functions. To evaluate social welfare functions, we require interpersonally comparable, cardinalized measures of individual utility. We follow the standard procedure in this literature of constructing money metric cardinalizations of utility. A money metric utility cardinalization $\widetilde{V}$ of a given indirect utility function V is the monotonic transformation of V

having the property that, evaluated at base prices $\overline{P}_1, \overline{P}_2$, the function $\widetilde{V} = M$. We therefore may define cardinalized utility $\widetilde{V}$ by

$$\widetilde{V}\left(U_1 P_1/M, U_2 P_2/M\right) = H\left[V\left(U_1 P_1/M, U_2 P_2/M\right), U_1, U_2, \overline{P}_1, \overline{P}_2\right]$$

where, by definition, the monotonic transformation function H is chosen so that

$$\widetilde{V}\left(U_1 \overline{P}_1/M, U_2 \overline{P}_2/M\right) = M$$

for all values of M, U_1, U_2. We let base prices be $\overline{P}_1 = \overline{P}_2 = 1$. Strict monotonicity of H as a function of V, and of V as a function of M, ensures that such a function H exists and that the resulting function $\widetilde{V}$ is unique.

The money-metric function $\widetilde{V}$ gives a utility level that may be interpreted as the number of dollars that, at base prices, delivers the same level of utility that the consumer can achieve with a budget of M dollars when facing prices P_1, P_2. Attained utility depends on U_1, U_2, so the function H that yields $\widetilde{V}$ will likewise depend on these Barten scales. This construction of a cardinalized, interpersonally comparable utility function $\widetilde{V}$ as a transformation of V, while required for evaluating social welfare, has no observable affects on demand functions. It is therefore not possible to test the assumption (standard in this literature) that welfare can be evaluated based on a social welfare function of the form $S(\widetilde{V}_1...\widetilde{V}_N)$, where $\widetilde{V}_i$ denotes the money metric utility of consumer i.

The range of social welfare functions that we consider are in the Atkinson (1970) Mean-of-Order-r class, defined by

$$S_r(\widetilde{V}_1...\widetilde{V}_N) = \left(\frac{1}{N}\sum\left(\widetilde{V}_i\right)^r\right)^{1/r} \text{ for } r \neq 0, \text{ and } S_r(\widetilde{V}_1...\widetilde{V}_N) = \exp\left(\frac{1}{N}\sum\ln\widetilde{V}_i\right) \text{ for } r = 0.$$

We use $r = -1, 0, 1$ corresponding to the harmonic, geometric and arithmetic mean of individual money metric utility. The resulting social welfare function S_1 is inequality neutral, while S_0 and S_{-1} are inequality averse. We compute proportionate welfare losses $\triangle S_r$ equal to welfare at base prices minus welfare at new prices divided by welfare at base prices. The money metric at base prices $(\overline{P}_1 = \overline{P}_2 = 1)$ equals the budget M_i assigned to each consumer. Letting $\widetilde{V}_i$ be the money metric at new prices $P_1 = 1.5$, $P_2 = 1$, we have

$$\triangle S_r = \left(S_r(M_1...M_N) - S_r(\widetilde{V}_1...\widetilde{V}_N)\right)/S_r(M_1...M_N)$$

A first cut at welfare analysis is to employ a first-order approximation of the money metric. A standard approximation of the individual money-metric utility associated with a given price vector, which in our case is (1.5, 1), is given by $\widetilde{V}_i = M_i/(W_{1i} * 1.5 + (1 - W_{1i}) * 1)$. This is just the Laspeyres index approximation to the true cost of living index for consumer i. The welfare loss associated with this approximate money metric accounts for some heterogeneity across individuals (because W_{1i} differs across individuals) but does not account for substitution responses. The arithmetic mean welfare index computed using this approximate money-metric utility shows a welfare loss of 6.30 per cent. The inequality-averse geometric and harmonic mean welfare indices show higher welfare losses of 6.74 and 7.18 per cent, respectively. This

is because energy budget shares W_{1i} are negatively correlated with budgets M_i, so the welfare indices that upweight poorer households show a greater welfare loss associated with lower budgets and hence with higher energy budget shares.

These approximate social welfare loss numbers can be compared with the estimates from Models 1 and 2, provided in Table 5. Standard errors accounting for the sampling variability of the estimated parameters were generated via simulation and are shown in *italics*. In Table 5 we account for both substitution effects and heterogeneity using our models. In the upper panel, we account for only observed heterogeneity using the estimates of Models 1 and 2. In the lower panel of Table 5, we also account for unobserved preference heterogeneity (the random component of Barten scales) using Model 2.

Table 5: Social Welfare Loss: 50% Energy Price Increase		Model 1		Model 2	
Welfare Loss, Per Cent	r	Estimate	*Std Err*	Estimate	*Std Err*
$\alpha_j, \widetilde{U}_j = 1$					
Arithmetic Mean	1	4.81	*0.19*	5.14	*0.14*
Geometric Mean	0	5.27	*0.26*	5.43	*0.16*
Harmonic Mean	-1	6.14	*0.54*	5.48	*0.17*
$\alpha_j, \widetilde{U}_j$					
Arithmetic Mean	1			4.91	*0.17*
Geometric Mean	0			5.37	*0.22*
Harmonic Mean	-1			5.85	*0.30*

Every estimate in Table 5 is lower than its corresponding first-order approximation. For example, the welfare loss given the geometric mean welfare index is about 1.5 percentage points lower (about one one fourth lower) than the first-order approximation. This shows that accounting for substitution effects has a substantial effect on welfare. All the estimates also show welfare losses increasing with the inequality aversion of the welfare index. As discussed above, this is primarily due to the downward sloping Engel curves as seen in Figure 4.

Another feature seen in Table 5 is that the estimates based on Model 2, which account for unobserved preference heterogeneity, have smaller standard errors. For the more inequality-averse welfare measures, this improvement in precision is substantial, e.g., the estimated harmonic mean welfare loss for Model 2 has about half the standard error of that for Model 1. This is due to the fact that the Model 2 treatment of unobserved heterogeneity increases precision of estimated parameters over model 1, as discussed earlier.

Finally, we find that dealing with unobserved preference heterogeneity affects both the level and pattern of estimated welfare losses. There are two ways in which the misspecification of Model 1 compared to Model 2 matters for welfare analysis. First, Model 2 Engel Curves are on average less downward sloping than those of Model 1. This means that, even without accounting for unobserved preference heterogeneity, when we consider welfare functions that are very inequality-averse, Model 1 will tend to overstate welfare losses. For example, in the upper panel of Table 5, the harmonic mean index shows a welfare loss of 6.14 per cent for Model 1 but only 5.48 per cent for Model 2.

A second difference is that Model 2 has greater variance in individual utility losses than

does Model 1, and inequality-averse welfare functions will tend to penalize such variance. This effect can be seen in the right side of Table 5 comparing the upper and lower panels. In the lower panel we account for unobserved preference heterogeneity, which increases the welfare loss for inequality-averse welfare functions. However, this increase is relatively modest.

The welfare loss in Model 2 not accounting for unobserved preference heterogeneity (top panel) is 0.34 percentage points higher with the harmonic versus the arithmetic mean welfare index. Because these welfare loss measures are highly positively correlated, this difference is statistically significant, with a standard error of 0.05. Similarly, due to all of the above described effects, the Model 1 estimates of welfare are substantially biased, especially when the welfare index is inequality-averse. In the bottom panel of Table 5, where we add variation due to unobserved preference heterogeneity, the difference in estimated welfare loss measures is 0.94 percentage points, with a standard error of the difference of 0.15. Thus, appropriately accounting for unobserved preference heterogeneity dramatically increases the sensitivity of estimated welfare loss to the inequality-aversion of the welfare index.

Overall, our energy tax experiment yields two major conclusions. First, accounting for unobserved preference heterogeneity has a big impact on how much variation we find in the impacts of price changes. In our example, the standard deviation of cost-of-living impacts due to a 50% energy tax is more than doubled by accounting for unobserved preference heterogeneity by random Barten scales. Second, we find that accounting for unobserved preference heterogeneity changes welfare analyses in a variety of ways, and in particular decreases estimated welfare loss when the welfare index is inequality-averse.

In the Appendix we provide a large number of comparisons of our Model 2 against a variety of more general specifications and modeling assumptions. These results confirm that the patterns of price, expenditure, and welfare effects summarized above are robust to a wide range of possible sources of specification error.

6   Conclusions

We show nonparametric identification of a generalized random coefficients model. We also provide an empirical application in which the generalized random coefficient structure arises naturally from extending an existing commonly used economic model of observed heterogeneity to a model allowing for unobserved heterogeneity. In this Barten scales application, we allow for general forms of unobserved preference heterogeneity that are shown to be important for empirically evaluating the welfare effects of potential policy interventions such a carbon tax. For example, we find that accounting for unobserved preference heterogeneity more than doubles the estimated variation in impacts of an energy tax (as measured by the standard deviation across consumers of the cost of living impact of the tax).

Accounting for unobserved preference heterogeneity via Barten scales is economically important because, by including them, we reveal that the variance in the impacts of energy prices is particularly large among poorer consumers. This can be seen in Figure 6, where the variation in cost of living impacts of an energy price increase is largest at low values of total expenditures. This figure also confirms the previously known result that mean cost of living impacts of energy price increases are also higher for the poor. Roughly, these results say that not only are energy

taxes regressive on average (the previously known result), but also the degree of regressivity varies more among poorer consumers than among the rich.

This result has important implications for social welfare. Empirically, the unobserved preference heterogeneity revealed by our model strongly interacts with inequality aversion in social welfare calculations, reversing conclusions that would have been made using almost all previous demand models, which fail to account for such preference heterogeneity.

Useful areas for further work on the theory of generalized random coefficients would be finding conditions under which Theorem 2 alone could be used to identify the model without the identification at zero assumptions used in Theorem 1, formalizing the extent to which overidentification stemming from these theorems might be used to identify more general structures or to test the model, and investigating how the assumptions for identification might interact with assumptions needed for possible nonparametric estimators such as sieve maximum likelihood.

Our application focuses on consumers with single utility functions, that is, unitary households. A possible extension would be to consider collective household models. For example, Barten scales are used in the collective household models of Browning, Chiappori, and Lewbel (2013). It would also be useful to extend our carbon tax analyses into a general equilibrium setting, which would affect the conditional independence assumptions we used for identification of the random Barten scales.

7 Appendix A: Proofs

Before proving Theorem 1, we prove a couple of lemmas.

LEMMA 1: Let $\widetilde{Y}_k = G_k(X_k U_k)$ where G_k is a strictly monotonically increasing, function. Assume $U_k \perp X \mid Z$. The marginal distributions of U_k and X_k are continuous. The support of X_k includes zero, the support of U_k is a subset of the support of $\widetilde{Y}_k$, and for every r such that $G_k(r)$ is on the support of $\widetilde{Y}_k$ there exist an $x_k \neq 0$ on the support of X_k such that $f_{U_k}\left(x_k^{-1} r\right) \neq 0$. Assume the location and scale normalizations $G_k(0) = 0$ and $G_k(1) = y_0$ for some known y_0 in the support of $\widetilde{Y}_k$ are imposed. Let $r = H_k(\widetilde{y}_k)$ be inverse of the function G_k where $\widetilde{y}_k = G_k(r)$. Assume H_k is differentiable. Define $X_{(k)}$ to be the vector of all the elements of X except for X_k. Define the function $S_k(\widetilde{y}_k, \widetilde{x})$ by

$$S_k(\widetilde{y}_k, \widetilde{x}) = E\left[F_{\widetilde{Y}_k | X_k, X_{(k)}, Z}\left(\widetilde{y}_k \mid \widetilde{x}^{-1}, 0, Z\right)\right] = \int_{supp(Z)} F_{\widetilde{Y}_k | X_k, X_{(k)}, Z}\left(\widetilde{y}_k \mid \widetilde{x}^{-1}, 0, z\right) f_z(z) \, dz.$$

Then, for all x_k and $\widetilde{y}_k$ such that $x_k \neq 0$ and $f_{U_k}\left(x_k^{-1} H_k(\widetilde{y}_k)\right) \neq 0$,

$$H_k(\widetilde{y}_k) = sign\left(sign(x_k) \frac{\partial S_k\left(\widetilde{y}_k, x_k^{-1}\right)}{\partial x_k^{-1}}\right) \exp\left(\int_{y_0}^{\widetilde{y}} \frac{x_k \partial S_k\left(\widetilde{y}_k, x_k^{-1}\right) / \partial \widetilde{y}}{\partial S_k\left(\widetilde{y}_k, x_k^{-1}\right) / \partial x_k^{-1}} d\widetilde{y}_k\right) \qquad (29)$$

Note that if Z is discretely distributed, then the integral defining S_k becomes a sum. If Z is empty (so U_k and X are unconditionally independent) then $S_k(\widetilde{y}_k, \widetilde{x}) = F_{\widetilde{Y}_k | X_k, X_{(k)}}\left(\widetilde{y}_k \mid \widetilde{x}^{-1}, 0\right)$.

The main implication of Lemma 1 is that if the distribution $F_{\widetilde{Y}_k|X,Z}$ is identified, then the function H_k is identified by construction.

PROOF of Lemma 1: For any $\widetilde{y}_k = G_k(x_k U_k)$ and any $x_k > 0$ we have

$$
\begin{aligned}
F_{\widetilde{Y}_k|X_k,X_{(k)},Z}(\widetilde{y}_k \mid x_k, 0, z) &= \Pr\left(G_k(x_k U_k) \le \widetilde{y} \mid X_k = x_k, X_{(k)} = 0, Z = z\right) \\
&= \Pr\left(U_k \le x_k^{-1} H_k(\widetilde{y}) \mid X_k = x_k, X_{(k)} = 0, Z = z\right) \\
&= F_{U_k|X_k,X_{(k)},Z}\left[x_k^{-1} H_k(\widetilde{y}) \mid x_k, 0, z\right] = F_{U_k|Z}\left[x_k^{-1} H_k(\widetilde{y}) \mid z\right]
\end{aligned}
$$

where the last equality uses $U_k \perp X \mid Z$. Similarly for any $x_k < 0$ we have

$$
\begin{aligned}
F_{\widetilde{Y}_k|X_k,X_{(k)},Z}(\widetilde{y}_k \mid x_k, 0, z) &= \Pr\left(G_k(x_k U_k) \le \widetilde{y} \mid X_k = x_k, X_{(k)} = 0, Z = z\right) \\
&= \Pr\left(U_k \ge x_k^{-1} H_k(\widetilde{y}) \mid X_k = x_k, X_{(k)} = 0, Z = z\right) \\
&= 1 - F_{U_k|Z}\left[x_k^{-1} H_k(\widetilde{y}) \mid z\right]
\end{aligned}
$$

Together these equations say

$$
F_{U_k|Z}\left[x_k^{-1} H_k(\widetilde{y}_k) \mid z\right] = I(x_k < 0) + sign(x_k) F_{\widetilde{Y}_k|X_k,X_{(k)},Z}(\widetilde{y}_k \mid x_k, 0, z).
$$

So

$$
\begin{aligned}
F_{U_k}\left[x_k^{-1} H_k(\widetilde{y}_k)\right] &= \int_{supp(Z)} \left[I(x_k < 0) + sign(x_k) F_{\widetilde{Y}_k|X_k,X_{(k)},Z}(\widetilde{y}_k \mid x_k, 0, z)\right] f(z) \, dz. \\
&= I(x_k < 0) + sign(x_k) S\left(\widetilde{y}_k, x_k^{-1}\right)
\end{aligned}
$$

It follows that for any $x_k \ne 0$,

$$
\frac{\partial S\left(\widetilde{y}_k, x_k^{-1}\right)}{\partial x_k^{-1}} = sign(x_k) f_{U_k}\left[x_k^{-1} H_k(\widetilde{y}_k)\right] H_k(\widetilde{y}_k)
$$

and

$$
\frac{\partial S\left(\widetilde{y}_k, x_k^{-1}\right)}{\partial \widetilde{y}_k} = sign(x_k) f_{U_k}\left[x_k^{-1} H_k(\widetilde{y}_k)\right] x_k^{-1} \frac{\partial H_k(\widetilde{y}_k)}{\partial \widetilde{y}_k}
$$

So for $f_{U_k}\left[x_k^{-1} H_k(\widetilde{y}_k)\right] \ne 0$ it follows that

$$
\frac{x_k \partial S\left(\widetilde{y}_k, x_k^{-1}\right)/\partial \widetilde{y}_k}{\partial S\left(\widetilde{y}_k, x_k^{-1}\right)/\partial x_k^{-1}} = \frac{\partial H_k(\widetilde{y}_k)/\partial \widetilde{y}_k}{H_k(\widetilde{y}_k)} = \frac{\partial \ln|H_k(\widetilde{y}_k)|}{\partial \widetilde{y}_k}
$$

so

$$\exp\left(\int_{\widetilde{y}_0}^{\widetilde{y}_k} \frac{x_k \partial S\left(\widetilde{y}, x_k^{-1}\right)/\partial \widetilde{y}}{\partial S\left(\widetilde{y}, x_k^{-1}\right)/\partial x_k^{-1}} d\widetilde{y}\right) = \exp\left(\int_{\widetilde{y}_0}^{\widetilde{y}_k} \frac{\partial \ln|H_k\left(\widetilde{y}\right)|}{\partial \widetilde{y}} d\widetilde{y}\right)$$

$$= \exp\left(\ln|H_k\left(\widetilde{y}_k\right)| - \ln|H_k\left(\widetilde{y}_0\right)|\right) = |H_k\left(\widetilde{y}_k\right)|$$

where $H_k\left(\widetilde{y}_0\right) = 1$ follows from $G_k\left(1\right) = \widetilde{y}_0$. Finally

$$sign\left(sign\left(x_k\right)\frac{\partial S\left(\widetilde{y}_k, x_k^{-1}\right)}{\partial x_k^{-1}}\right) = sign\left(sign\left(x_k\right)sign\left(x_k\right)f_{U_k}\left[x_k^{-1}H_k\left(\widetilde{y}_k\right)\right]H_k\left(\widetilde{y}_k\right)\right)$$

$$= sign\left(f_{U_k}\left[x_k^{-1}H_k\left(\widetilde{y}_k\right)\right]H_k\left(\widetilde{y}_k\right)\right) = sign\left(H_k\left(\widetilde{y}_k\right)\right)$$

So the right side of equation (29) equals $sign\left(H_k\left(\widetilde{y}_k\right)\right)|H_k\left(\widetilde{y}_k\right)| = H_k\left(\widetilde{y}_k\right)$ as claimed.

LEMMA 2: If Assumption A1 holds and the normalization $G_k\left(0\right) = 0$ for all k holds, then $F_{U_0|Z}$ and the distribution function $F_{\widetilde{Y}|X,Z}\left(\widetilde{Y} \mid x, z\right)$ are identified, where $\widetilde{Y} = \sum_{k=1}^{K} G_k\left(X_k U_k\right)$.

PROOF of Lemma 2:

$$F_{Y|X,Z}\left(y \mid 0, z\right) = \Pr\left(G\left(0\right) + U_0 \leq y \mid X = 0, Z = z\right) = F_{U_0|X,Z}\left(y \mid 0, z\right) = F_{U_0|Z}\left(y \mid z\right)$$

identifies the distribution function $F_{U_0|Z}$ on the support of Y, which contains the support of U_0. Next define $\widetilde{Y} = Y - U_0$. Then since $Y = \widetilde{Y} + U_0$ and the distributions of $Y \mid X, Z$ and $U_0 \mid X, Z$ are identified, for each value of $X = x$, $Z = z$ apply a deconvolution (using the nonvanishing characteristic function of U_0) to identify the distribution of $\widetilde{Y} \mid X, Z$, where $\widetilde{Y} = \sum_{k=1}^{K} G_k\left(X_k U_k\right)$.

PROOF of Theorem 1: When $X_{(k)} = 0$ (equivalently, when $X = e_k x_k$ for some x_k) we get $\widetilde{Y} = G_k\left(X_k U_k\right) + \sum_{j \neq k} G_j\left(0\right) = G_k\left(X_k U_k\right)$. Define $\widetilde{Y}_k = G_k\left(X_k U_k\right)$. It follows that $F_{\widetilde{Y}_k|X_k,X_{(k)},Z}\left(\widetilde{y}_k \mid x_k, 0, z\right) = F_{\widetilde{Y}|X,Z}\left(\widetilde{y}_k \mid x_k e_k, z\right)$, so the distribution function on the left of this identity is identified, given by Lemma 2 that $F_{\widetilde{Y}|X,Z}$ is identified. Let $r = H_k\left(\widetilde{y}_k\right)$ denote the inverse of the function G_k where $\widetilde{y}_k = G_k\left(r\right)$. It follows by construction from Lemma 1 that $H_k\left(\widetilde{y}_k\right)$ is identified for every value of $\widetilde{y}_k$ on the support of $\widetilde{Y}_k$ satisfying the property that, for some x_k on the support of X_k, $f_{U_k}\left[x_k^{-1}H\left(\widetilde{y}_k\right)\right] \neq 0$. This identification of $H_k\left(\widetilde{y}_k\right)$ in turn means that the function $G_k\left(r\right)$ is identified for every r such that $G_k\left(r\right)$ is on the support of $\widetilde{Y}_k$ and there exist an x_k on the support of X_k such that $f_{U_k|Z}\left(x_k^{-1}r\right) \neq 0$. This then implies identification of G_k on its support. Finally, given identification of $F_{\widetilde{Y}|X,Z}$ and of $H_k\left(\widetilde{y}_k\right)$, the distribution function $F_{U_k|Z}$ is identified by $F_{U_k|Z}\left[H\left(\widetilde{y}\right)/x_k \mid z\right] = F_{\widetilde{Y}|X_k,X_{(k)},Z}\left(\widetilde{y} \mid x_k, 0, z\right)$ for $x_k > 0$ and $F_{U_k|Z}\left[H\left(\widetilde{y}\right)/x_k \mid z\right] = 1 - F_{\widetilde{Y}|X_k,X_{(k)},Z}\left(\widetilde{y} \mid x_k, 0, z\right)$ for $x_k < 0$.

PROOF of Corollary 1: For part i), Given any function G, for $k = 1, ..., K$, define $G_k\left(X_k U_k\right) = G\left(X_k U_k e_k\right) = G\left(0, ..., 0, X_k U_k, 0, ..., 0\right)$ and define $\widetilde{G}$ by $\widetilde{G}\left(X_1 U_1, ..., X_K U_K\right) = G\left(X_1 U_1, ..., X_K U_K\right)-$

$\sum_{k=1}^{K} G_k (X_k U_k)$. Then, by construction, part i) holds. For part ii), we have that the function $G (X_1 U_1, ..., X_K U_K)$ is zero when evaluated at $X = 0$ or at $X = X_k e_k$ for any k, so evaluated at any such value of X, equation (3) is equivalent to equation (2). For equation (2), the proof of Theorem 1 shows (for $k = 1, ..., K$) identification of $F_{U_k|Z}$ and of G_k only using $X = 0$ and $X = X_k e_k$, so these functions are also identified for equation (3).

PROOF of Theorem 2:

Let $\Omega_X = supp (X)$, $\Omega_{X|Z} = supp (X \mid Z)$, etc. and let $F (U \mid Z) = F (U_1, U_2...U_K \mid Z)$ By the definition of $\lambda_{\rho,t} (Z)$ we have $\lambda_{\rho,t} (Z) =$

$$\int_{X \in s\Omega_X} \int_{U \in \Omega_{U|Z}} h_\rho (G (X_1 U_1, ..., X_K U_K), t_1, ..., t_K) dF (U \mid Z) X_1^{t_1-1} X_2^{t_2-1}...X_K^{t_K-1} dX_1 dX_2...dX_K$$

$$= \int_{U \in \Omega_{U|Z}} \int_{X \in \Omega_{X|Z}} h_\rho (G (X_1 U_1, ..., X_K U_K), t_1, ..., t_K) X_1^{t_1-1} X_2^{t_2-1}...X_K^{t_K-1} dX_1 dX_2...dX_K dF (U \mid Z)$$

where the second equality follows from Fubini's theorem and $supp (X) = supp (X \mid Z)$. Do a change of variables on the inner integral, letting $s_k = X_k U_k$ for $k = 1, ..., K$ to get $\lambda_{\rho,t} (Z) =$

$$\int_{U \in s\Omega_{U|Z}} \int_{s \in supp(X_1 U_1,...,X_K U_K|Z,U)} h_\rho (G (s_1, ..., s), t_1, ..., t_K) s_1^{t_1-1} s_2^{t_2-1}...s_K^{t_K-1} U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K}$$

$$ds_1 ds_2...ds_K dF (U \mid Z)$$

$$= \int_{U \in \Omega_{U|Z}} \int_{s \in \Omega_X} h_\rho (G (s_1, ..., s), t_1, ..., t_K) s_1^{t_1-1} s_2^{t_2-1}...s_K^{t_K-1} U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} ds_1 ds_2...ds_K dF (U \mid Z)$$

$$= \int_{U \in \Omega_{U|Z}} \int_{s \in \Omega_X} h_\rho (G (s_1, ..., s), t_1, ..., t_K) s_1^{t_1-1} s_2^{t_2-1}...s_K^{t_K-1} ds_1 ds_2...ds_K U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} dF (U \mid Z)$$

$$= \int_{U \in \Omega_{U|Z}} \kappa_{\rho,t} U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} dF (U \mid Z)$$

$$= \kappa_{\rho,t} \int_{U \in \Omega_{U|Z}} U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} dF (U \mid Z) = \kappa_{\rho,t} E \left(U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} \mid Z \right)$$

where the second equality above uses $supp (U_1 X_1, ..., U_K X_K \mid Z, U) = supp (X)$. It therefore follows that moments $E \left(U_1^{-t_1} U_2^{-t_2}...U_K^{-t_K} \mid Z \right)$ are identified for all positive integers $t_1,...,t_k$ by equalling the ratio of identified objects $\lambda_{\rho,t} (Z) / \kappa_{\rho,t}$. To identify moments where $t_k = 0$ for one or more values of k, redefine $\lambda_{\rho,t} (Z)$ and $\kappa_{\rho,t}$ setting the corresponding X_k terms equal to zero. For example, to identify moments having $t_1 = 0$, replacing equation (5) with

$$\lambda_{\rho,t} (Z) = \int_{(X_2,...X_K) \in supp(X_2,...X_K)} E \left[h_\rho \left(\widetilde{Y}, 0, t_2, ..., t_K \right) \mid X_1 = 0, X_2...X_K, Z \right] X_2^{t_2-1}...X_K^{t_K-1} dX_2...dX_K$$

and correspondingly redefining $\kappa_{\rho,t}$ as

$$\kappa_{\rho,t} = \int_{(s_2,...s_K) \in supp(X_2,...X_K)} h_\rho [G (0, s_2, ...s_K), 0, ..., t_K] s_2^{t_2-1}...s_K^{t_K-1} ds_2...ds_K$$

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gives $\lambda_{\rho,t}(Z)/\kappa_{\rho,t}$ equal to $E\left(U_2^{-t_2}...U_K^{-t_K} \mid Z\right)$.

We have now shown that $E\left(U_1^{-t_1}U_2^{-t_2}...U_K^{-t_K} \mid Z\right)$ is identified for any nonnegative integers $t_1,...,t_k$. Since the MGF of $U_1^{-1}U_2^{-1}...U_K^{-1} \mid Z$ exists, it follows that the joint distribution of $U_1^{-1}U_2^{-1}...U_K^{-1} \mid Z$ is identified from these moments, and therefore that the joint distribution $F_{U|Z}(U_1,...,U_K \mid Z)$ is identified.

PROOF of Corollary 2: Theorem 1 identifies the functions $G_1, G_2, ..., G_K$, and shows that the distribution of $\widetilde{Y}$ defined by $\widetilde{Y} = G(X_1 U_1, ..., X_K U_K)$ is identified. Assumption A5 therefore holds, and by Theorem 2 the joint distribution function $F_{U|Z}(U_1,...,U_K \mid Z)$ is identified.

PROOF of Theorem 3: As discussed in the text, a property of Barten scales (which can be readily verified using Roys identity) is that, if $V(X_1, X_2)$ is the indirect utility function corresponding to the demand function $\omega_1(X_1, X_2)$, then up to an arbitrary monotonic transformation $H(V, U_1, U_2)$ of V, the indirect utility function corresponding to $\omega_1(U_1 X_1, U_2 X_2)$ is $V(U_1 X_1, U_2 X_2)$, and vice versa. It therefore suffices to prove that the theorem holds with $U_1 = U_2 = 1$.

By equation (10), given any indirect utility function V, the corresponding demand function ω_1 is given by

$$\omega_1(X_1, X_2) = \frac{\partial V(X_1, X_2)/\partial \ln X_1}{\left[\partial V(X_1, X_2)/\partial \ln X_1\right] + \left[\partial V(X_1, X_2)/\partial \ln X_2\right]} \tag{30}$$

This is just one way to write Roys identity in a demand system of two goods. Then by the definition of λ, we have that for any demand function ω_1, the corresponding indirect utility function V satisfies

$$\lambda[\omega_1(X_1, X_2)] = \ln\left(\frac{\partial V(X_1, X_2)}{\partial \ln X_1}\right) - \ln\left(\frac{\partial V(X_1, X_2)}{\partial \ln X_2}\right) \tag{31}$$

and similarly, given any V the corresponding ω_1 satisfies equation (31).

It follows immediately that, given any differentiable functions $h_1(X_1)$ and $h_2(X_2)$, if $V(X_1, X_2) = h_1(X_1) + h_2(X_2)$ then equation (31) equals

$$\lambda[\omega_1(X_1, X_2)] = \ln\left(\frac{\partial h_1(X_1)}{\partial \ln X_1}\right) - \ln\left(\frac{\partial h_2(X_2)}{\partial \ln X_2}\right) \tag{32}$$

which is in the form of equation

$$\lambda[\omega_1(X_1, X_2)] = g_1(X_1) + g_2(X_2), \tag{33}$$

showing that any additive indirect utility function generates a demand equation in the form of (33). Also, by equation (32) given the functions h_1 and h_2 we can define g_1 and g_2 by

$$g_1(X_1) + g_2(X_2) = \ln\left(\frac{\partial h_1(X_1)}{\partial \ln X_1}\right) - \ln\left(\frac{\partial h_2(X_2)}{\partial \ln X_2}\right). \tag{34}$$

To go the other direction, start by supposing that equation (33) holds for some functions g_1 and g_2. We will apply the following two component special case of Lemma 4.1 in Blackorby, Primont, and Russell (1978, p. 160): Assume a function $F(r_1, r_2)$ is twice continuously differentiable and strictly increasing in its elements. Then $F(r_1, r_2) = F_0(F_1(r_1) + F_2(r_2))$ for some functions F_0, F_1, and F_2 if and only if

$$\frac{\partial}{\partial r_1} \ln \left(\frac{\partial F(r_1, r_2)/\partial r_2}{\partial F(r_1, r_2)/\partial r_1} \right) = \psi(r_1)$$

for some function ψ. Taking the derivative of equation (31) with respect to $\ln X_1$ and using equation (33) gives

$$\frac{\partial \lambda [\omega_1(X_1, X_2)]}{\partial \ln X_1} = \frac{\partial g_1(X_1)}{\partial \ln X_1} = \ln \left(\frac{\partial V(X_1, X_2)/\partial \ln X_2}{\partial V(X_1, X_2)/\partial \ln X_1} \right).$$

Apply the Lemma with $r_j = \ln X_j$, $F(r_1, r_2) = V(X_1, X_2)$, $\psi(r_1) = \partial g_1(X_1)/\partial \ln X_1$, and $F_j(r_j) = h_j(X_j)$ to prove that $V(X_1, X_2)$ must equal $F_0(h_1(X_1) + h_2(X_2))$ for some function F_0, which by the properties of indirect utility functions must be monotonically increasing (recall also that twice differentiability was one of the assumed properties of our indirect utility functions). Finally, applying equation (30) to this indirect utility function shows that equation (34) holds, which we can integrate to obtain

$$\int e^{g_1(x_1)} d \ln x_1 + \int e^{-g_2(x_2)} d \ln x_2 = h_1(X_1) + h_2(X_2).$$

Together these results prove the first part Theorem 3. Adding back the Barten scales U_1 and U_2 to the functions g_1, g_2, h_1, and h_2 proves equations (15) and (14). The properties of the functions h_1 and h_2 given at the end of Theorem 2 follow from the fact that the indirect utility function $h_1(U_1 P_1/M) + h_2(U_2 P_2/M)$ must possess the standard properties of all indirect utility functions, i.e., homogeneity and quaisconvexity in P_1, P_2, and M, nondecreasing in each price, and increasing in M.

8   Appendix B: Empirical Model Robustness Checks

Given that our estimated model is parametrically specified, one may worry about the possible impacts of misspecification. In this appendix we provide an extensive set of analyses to verify the robustness of our empirical results to a wide variety of alternative model specifications, including adding complexity to the utility function specification, relaxing the parametric structure on preference and error distributions, and dealing with potential heteroskedasticity of errors or endogeneity of regressors.

A brief summary of these results is that, while some departures from our baseline Model 2 are statistically significant, none result in big changes in our economic analyses or conclusions, indicating that our results are robust to many different possible sources of misspecification.

8.1 Interactions in Utility

The indirect additive utility function $V^{-1} = h_1(X_1) + h_2(X_2)$ from Theorem 3 restricts price interaction effects. To relax the restrictiveness (in terms of cross effects) of the resulting additive demand functions we here consider adding interaction terms to the model of Theorem 3, giving an indirect utility function of the form

$$V^{-1} = h_1(X_1) + h_2(X_2) + S(X_1, X_2, \xi) \tag{35}$$

where the interaction function S has a simple parametric form (with parameter vector ξ), as in equation (9).

By applying Roy's identity to (35), then as before logit transforming the result, adding Barten scales, and adding the error term U_0 we obtain the demand model:

$$\lambda(W_1) = \ln\left[e^{g_1(U_1 X_1)} + M_1(U_1 X_1, U_2 X_2, \xi)\right] - \ln\left[e^{-g_2(U_2 X_2)} + M_2(U_1 X_1, U_2 X_2, \xi)\right] + U_0 \tag{36}$$

where

$$M_j(U_1 X_1, U_2 X_2, \xi) = \frac{\partial S(U_1 X_1, U_2 X_2, \xi)}{\partial \ln(U_j X_j)} \tag{37}$$

and, as in Theorem 3,

$$\frac{\partial h_1(X_1)}{\partial \ln X_1} = e^{g_1(X_1)} \quad \text{and} \quad \frac{\partial h_2(X_1)}{\partial \ln X_2} = e^{-g_2(X_2)}.$$

Keeping the same polynomial expansions for g_k as before, namely, equations (18) and (19), yields the same h_k as before, equation (20), and when substituted into the above gives the demand function

$$\begin{aligned}
\lambda(W_1) &= \omega_{S1}(U_1 X_1, U_2 X_2, \beta, \xi) + U_0 \tag{38}\\
&= \ln\left[\begin{array}{c}\left(\beta_{10} + \beta_{11}\ln(U_1 X_1) + \beta_{12}(\ln(U_1 X_1))^2 + ... + \beta_{1S}(\ln(U_1 X_1))^S\right)^2 \\ +M_1(U_1 X_1, U_2 X_2, \xi)\end{array}\right] \\
&\quad - \ln\left[\begin{array}{c}\left(\beta_{20} + \beta_{21}\ln(U_2 X_2) + \beta_{22}(\ln(U_2 X_2))^2 + ... + \beta_{2S}(\ln(U_2 X_2))^S\right)^2 \\ +M_2(U_1 X_1, U_2 X_2, \xi)\end{array}\right] + U_0
\end{aligned}$$

The demand function given by equation (38) is the same as (21), except for the addition of the M_j functions given by equation (37), which embody the desired price interaction terms.

We consider two models for S. One is the simple multiplicative interaction term

$$S = U_1 X_1 U_2 X_2 \xi_{11},$$

which has

$$M_1 = M_2 = U_1 X_1 U_2 X_2 \xi_{11}.$$

This is the most obvious model for interactions, but it forces $M_1 = M_2$, pushing the g_1 and g_2 terms by the same amount. We therefore also consider

$$S = U_1 X_1 U_2 X_2 \xi_{11} + (U_1 X_1)^2 U_2 X_2 \xi_{21} + U_1 X_1 (U_2 X_2)^2 \xi_{12}$$

37

which allows M_1 and M_2 to differ:

$$M_1 = U_1 X_1 U_2 X_2 \left(\xi_{11} + 2U_1 X_1 \xi_{21} + U_2 X_2 \xi_{12}\right),$$
$$M_2 = U_1 X_1 U_2 X_2 \left(\xi_{11} + U_1 X_1 \xi_{21} + 2U_2 X_2 \xi_{12}\right).$$

For both models, we restrict the parameters ξ to be weakly positive because negative values of the parameters ξ induce violations of regularity (that is, budget shares outside 0, 1, or equivalently, taking logs of negative numbers) for large values of U_j.[11]

Table 6: Interactions Terms in Utility

Parameter		Model 2 llf=-9706.9 Estimate	Std Err	Model 2 w interaction llf=-9644.3 Estimate	Std Err
σ_0		0.469	0.009	0.476	0.016
σ_1		0.165	0.036	0.229	0.044
σ_2		1.336	0.011	1.851	0.024
ρ		0.883	0.100	0.839	0.199
ξ_{11}				0.671	0.093
std dev	$\ln \alpha_1$	0.252		0.261	
	$\ln \alpha_2$	0.380		0.694	
correlation	$\ln \alpha_1, \ln \alpha_2$	-0.700		-0.540	
(all obs)	$\ln U_1, \ln U_2$	0.293		0.361	
correlation	$\ln \alpha_1, \ln \alpha_2$	0.105		0.434	
(renter=0)	$\ln U_1, \ln U_2$	0.699		0.729	
correlation	$\ln \alpha_1, \ln \alpha_2$	0.087		0.419	
(renter=1)	$\ln U_1, \ln U_2$	0.691		0.722	
budget shares	$M, \alpha_j = \overline{\alpha}_j, P = 1, \widetilde{U}_j = 1$	0.147	0.030	0.155	0.036
(mean and std dev)	$M, \alpha_j, P = 1, \widetilde{U}_j = 1$	0.148	0.048	0.153	0.049
	$M, \alpha_j, P, \widetilde{U}_j = 1$	0.151	0.047	0.155	0.049
	$M, \alpha_j, P, \widetilde{U}_j$	0.137	0.065	0.138	0.061
Cost-of-Living	$M, \alpha_j = \overline{\alpha}_j, \widetilde{U}_j = 1$	5.66	1.30	5.67	1.39
(mean and std dev)	$M, \alpha_j, \widetilde{U}_j = 1$	5.64	1.69	5.55	1.61
	$M, \alpha_j, \widetilde{U}_j$	5.37	4.31	5.35	4.87
Welfare Loss	Arithmetic Mean	4.91	0.17	4.79	0.23
(estimate and std err)	Geometric Mean	5.37	0.22	5.24	0.29
	Harmonic Mean	5.85	0.30	5.70	0.41

Table 6 gives results for the model with the simple interaction $S = U_1 X_1 U_2 X_2 \xi_{11}$ on the right hand side, and for our baseline model on the left. We do not present estimates for the more complicated interaction model because the estimated values of ξ_{21} and ξ_{12} in that case

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[11]To allow for possible negative interactions, we also tried the model $S = \alpha / \left(1 + U_1 X_1 U_2 X_2\right)$, the magnitude of which is bounded by $\alpha \geq 0$. This specification has $M_1 = M_2 = -U_1 X_1 U_2 X_2 \alpha / \left(1 + U_1 X_1 U_2 X_2\right)$. We then estimate equation (38) imposing the same normalizations and using the same sieve maximum likelihood method as before. Empirically, this did not yield any different results.

were zero (that is, the positivity restrictions on those parameters were binding), reducing that specification to the simple interaction model. Table 6 gives a subset of estimated coefficients, summary statistics on budget shares and cost of living impacts, and social welfare analysis analogous to those results in Tables 2, 3, 4 and 5.

Looking at the top panel of Table 6, which shows the estimated parameters governing the distributions of U_0, $\widetilde{U}_1$ and $\widetilde{U}_2$, we see that the inclusion of the interaction term ξ_{11} does not much affect the estimated variance of U_0 or $\widetilde{U}_1$, or the correlation of $\widetilde{U}_1$ and $\widetilde{U}_2$, but it does seem to increase the estimated variance of $\widetilde{U}_2$. Further, the inclusion of the interaction term ξ_{11} reduces the precision of these parameter estimates. In particular, it doubles the estimated standard error of ρ.

The magnitude of the estimated interaction term appears relatively large and significant.[12] However, since we have no estimated difference between M_1 and M_2, the model becomes $\lambda(W_1) = \ln(e^{g_1} + M_1) - \ln(e^{-g_2} + M_1) + U_0$, which approximately equals $\ln(e^{g_1}) - \ln(e^{-g_2}) + U_0$ using $\ln(e^{g_1} + M_1) \approx \ln(e^{g_1}) + \ln(M_1)$. As a result, the overall impact of the interaction terms on behavior is rather small for most consumers. This can be seen by comparing the left and right sides of the lower portions of Table 6. For the same reason, inclusion of interactions generally reduces the precision with which objects like welfare losses are measured, while having almost no impact on their estimated magnitude. For example, the bottom lines of Table 6 show that including the interaction term changes estimated welfare impacts by less than 3%, and increases standard errors of the welfare effects by more than that amount.

Our overall conclusion is that possible interaction effects (and hence potential violations of the additivity assumed by Theorem 3) may be present, but if so, they do not materially affect our conclusions.

8.2 Higher Order Polynomials in Demand

Our model specifies demand functions using squared third-order polynomials for functions g_1 and g_2. This is consistent with a large theoretical and empirical literature on demand estimation finding that three terms in total expenditures per demand equation are sufficient to capture the shapes of various demand functions. See, e.g., Lewbel (2008) and references therein. This specification yields a total of seven parameters (the beta's) governing the shape of the demand function for energy over X_1, X_2.

To check for adequacy of this specification we also estimated models with fourth and fifth order polynomials. We find that the fourth and fifth order terms for energy goods (β_{14}, β_{15}) are small, and are both individually and jointly insignificant. The estimated fourth and fifth-order terms for nonenergy goods (β_{24}, β_{25}) are very large in magnitude but also have very large standard errors, indicating numerical parameter instability, which is a common problem that indicates overfitting when estimating high order polynomials with small data sets. To save space we do not reproduce a full summary of alternative estimates as in Table 6, but simply note that these high-magnitude high-variance parameters associated with higher order polynomials introduce a large amount of noise into the fitted model. For example, this alternative specification

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[12]The product $\widetilde{U}_1\widetilde{U}_2$ has a median of 1 and a right-skewed distribution, which makes the distribution of $M_1 = M_2$ also right-skewed, with a mean of 0.047 but a median of 0.021 and an inter-decile range of $(0.005, 0.098)$.

nearly triples the standard errors of estimated welfare losses. We therefore confine our empirical analysis to the more stable squared third-order polynomials in our baseline specification.

8.3 Specification of the Distributions of Unobservables

Our estimated models assume that the measurement error term U_0 is distributed normally, and our baseline Model 2 assumes $\widetilde{U} = (\widetilde{U}_1, \widetilde{U}_2)$ has a bivariate log-normal distribution. Here we consider more general specifications for these distributions, based on hermite polynomial expansions of the type commonly used for sieve estimation. Denote the joint density of $\widetilde{U}$ by $f_{\widetilde{U}J}$ and the density for U_0 by f_{0J}, where J indexes the order of our polynomial expansions. Specifically, we let $J = 0$ index our baseline parameterizations where $\ln \widetilde{U}$ is a mean zero bivariate normal and U_0 is a mean zero normal. Higher values of J then correspond to richer distribution specfications based on polynomial expansions that multiply the base normal distributions.

We employ the seminonparametric density models used by Gallant and Nychka (1987). Ignoring trimming for now, the joint density of $\widetilde{U}$ is specified as

$$f_{\widetilde{U}J}\left(\widetilde{U}_1, \widetilde{U}_2, \gamma, \sigma, \rho\right) = \frac{\left(\sum_{j_1=0}^{J} \sum_{j_2=0}^{J} \gamma_{j_1 j_2} \left(\ln \widetilde{U}_1\right)^{j_1} \left(\ln \widetilde{U}_2\right)^{j_2}\right)^2}{\widetilde{U}_1 \widetilde{U}_2 \widetilde{C}_J (\gamma, \sigma, \rho)} \cdot \tag{39}$$

$$\exp\left(\frac{\left(\frac{\ln \widetilde{U}_1}{\sigma_1}\right)^2 - 2\rho \left(\frac{\ln \widetilde{U}_2}{\sigma_2}\right)\left(\frac{\ln \widetilde{U}_1}{\sigma_1}\right) + \left(\frac{\ln \widetilde{U}_2}{\sigma_2}\right)^2}{-2\left(1 - \rho^2\right)}\right) \tag{40}$$

where γ is the vector of polynomial coefficients $\gamma_{j_1 j_2}$, $\sigma = (\sigma_1, \sigma_2)$, and $\widetilde{C}_J (\gamma, \sigma, \rho)$ is the constant necessary to make $f_{\widetilde{U}J}$ integrate to one. Because we scale by $\widetilde{C}_J$, we can without loss of generality let $\gamma_{00} = 1$. In the simplest model when $J = 0$ this reduces to our baseline specification of a bivariate log normal $\widetilde{U}$ with $\widetilde{C}_J (\gamma, \sigma, \rho) = 2\pi \sigma_1 \sigma_2 \left(1 - \rho^2\right)$, each $\ln \widetilde{U}_k$ having mean zero and variance σ_k^2, and correlation coefficient ρ. Note that we don't need to explicitly model the mean of $\ln \widetilde{U}$ and hence the scaling of each $\widetilde{U}_k$, because the scaling of $\widetilde{U}_k$, like that of $\alpha_k (Z)$, is freely absorbed into the β_{ks} parameters. As in our baseline specification, we trim this distribution of $\ln \widetilde{U}_k$ at ± 3 standard deviations. We trim to bound the support of $\widetilde{U}_k$ away from zero and to ensure existence of a moment generating function, both as assumed by Theorem 2. This trimming was found to have minimal numerical effects on our empirical estimates.

We analogously model the density of U_0 as

$$f_{0J} (U_0, \delta, \sigma_0) = \frac{\left(\sum_{j=0}^{J} \delta_j U_0^j\right)^2}{C_J (\delta, \sigma_0)} \exp\left(\frac{-1}{2}\left(\frac{U_0}{\sigma_0}\right)^2\right) \tag{41}$$

where $\delta_0 = 1$, and δ_1 is set to ensure that the error term U_0 has mean zero[13].

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[13]It is straightforward to verify that the function C_J needed for estimation is given by $C_J (\delta, \sigma_0) = (2\pi)^{1/2} \sum_{j=0}^{J} \sum_{k=0}^{J} \delta_j \delta_k \sigma_0^{j+k+1} \mu_{j+k}$ where $\mu_{j+k} = E(e^{j+k})$ for a standard nor-

Estimation then proceeds by replacing f_0 and $f_{\widetilde{U}}$ with f_{0J} and $f_{\widetilde{U}J}$ in equations (25) and (27), including γ and δ as additional parameters to estimate. Table 7 summarizes results with $J = 3$ for U_0 and with $J = 2$ for $\ln \widetilde{U}_1$, $\ln \widetilde{U}_2$. Just like Table 6, Table 7 gives selected estimates from this model on the right, along with the corresponding estimates from our baseline model ($J = 0$) on the left for comparison, including summary statistics on estimated budget shares, cost-of-living indices, and social welfare losses. $J = 2$ was the most flexible model we could estimate for the joint distribution of $\ln \widetilde{U}_1$, $\ln \widetilde{U}_2$, in that attempts at estimating models with $J > 2$ produced very large standard errors and numerical instability of the sort discussed in the previous subsection.

Figures 7 and 8 show the estimated distributions of U_0 and $\ln U$ based on these polynomial expansions. Figure 7 shows only a very small departure from our baseline normal model for the density of U_0. The estimated joint distribution of $\ln \widetilde{U}$ differs from our baseline lognormal estimate mainly in having a larger variance (particularly for $\ln \widetilde{U}_1$), and has a small second mode. The resulting estimate of the joint density of $\ln U$ is shown in Figure 8. Comparing Figures 3 and 8, the larger variance of $\ln \widetilde{U}$ results in the two modes of Figure 3 largely merging into one, while the second mode in $\ln \widetilde{U}$ puts some extra mass at smaller values of U.

The estimates in these figures and in Table 7 show departures from the baseline model that are statistically significant, but turn out to have numerically modest and insignificant affects on our economic analyses. The estimates of every summary measure and welfare implication of the model (see the bottom section of Table 7) are changed by less than than one standard error, meaning that we could not reject the null hypothesis that any one of the economic implications of the more richly specified model actually equal the values estimated from our baseline model. Moreover, to the extent that the results from these more general specifications differ from baseline, they all strengthen rather than weaken our economic conclusions. For example, the estimated variation in cost-of-living impacts due to a 50% increase in the price of energy is larger in this model than in the baseline model, and the extent to which increasing inequality aversion increases welfare losses is also somewhat larger.

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mal e. Similarly, $E(U_0) = 0$ if and only if $\zeta_J(\delta, \sigma_0) = \sum_{j=0}^{J} \sum_{k=0}^{J} \delta_j \delta_k \sigma_0^{j+k} \mu_{j+k+1} = 0$, which we imposed on estimation by appropriately specifying δ_1. For example, when $J = 3$ we get $C_3(\delta, \sigma_0) = (2\pi)^{1/2} \left(15\delta_3^2 \sigma_0^7 + \left(6\delta_1 \delta_3 + 3\delta_2^2 \right) \sigma_0^5 + \left(2\delta_2 + \delta_1^2 \right) \sigma_0^3 + \sigma_0 \right)$ and $\zeta_3(\delta, \sigma_0) = 2 \left(\delta_1 \sigma_0 + 3 (\delta_1 \delta_2 + \delta_3) \sigma_0^3 + 15 \delta_2 \delta_3 \sigma_0^5 \right)$, so $E(U_0) = 0$ is imposed by setting $\delta_1 = -\left(3\delta_3 \sigma_0^3 + 15 \delta_2 \delta_3 \sigma_0^5 \right) / (\sigma_0 + 3\delta_2)$.

Table 7: Non-Normal Unobserved Preference Heterogeneity

Parameter		Model 2 llf=-9706.88 Estimate	Std Err	Model 2 w squared poly llf=-9561.92 Estimate	Std Err
σ_0		0.469	0.009	0.459	0.017
δ_1				-0.946	0.074
δ_2				0.563	0.110
σ_1		0.165	0.036	0.532	0.009
σ_2		1.336	0.011	1.401	0.062
γ_{10}				-2.431	0.404
γ_{01}				-0.109	0.113
γ_{11}				0.361	0.215
γ_{20}				-0.225	0.286
γ_{21}				1.299	0.260
γ_{22}				-0.236	0.070
γ_{02}				0.037	0.043
γ_{1122}				-0.133	0.063
ρ		0.883	0.100	0.691	0.034
std dev	$\ln \widetilde{U}_1$	0.165		0.627	
	$\ln \widetilde{U}_2$	1.336		1.715	
correlation	$\ln \widetilde{U}_1, \ln \widetilde{U}_2$	0.883		0.740	
std dev	$\ln \alpha_1$	0.252		0.190	
	$\ln \alpha_2$	0.380		0.725	
	$\ln U_1$	0.298		0.631	
	$\ln U_2$	1.350		1.802	
correlation	$\ln \alpha_1, \ln \alpha_2$	-0.700		-0.488	
(all obs)	$\ln U_1, \ln U_1$	0.293		0.574	
correlation	$\ln \alpha_1, \ln \alpha_2$	0.105		0.397	
(renter=0)	$\ln U_1, \ln U_1$	0.699		0.707	
correlation	$\ln \alpha_1, \ln \alpha_2$	0.087		0.384	
(renter=1)	$\ln U_1, \ln U_1$	0.691		0.710	
budget shares	$M, \alpha_j = \overline{\alpha}_j, P = 1, \widetilde{U}_j = 1$	0.147	0.030	0.152	0.027
(mean, sd)	$M, \alpha_j, P = 1, \widetilde{U}_j = 1$	0.148	0.048	0.150	0.062
	$M, \alpha_j, P, \widetilde{U}_j = 1$	0.151	0.047	0.154	0.063
	$M, \alpha_j, P, \widetilde{U}_j$	0.137	0.065	0.130	0.068
Cost-of-Living	$M, \alpha_j = \overline{\alpha}_j, \widetilde{U}_j = 1$	5.66	1.30	6.10	1.75
(mean, sd)	$M, \alpha_j, \widetilde{U}_j = 1$	5.64	1.69	5.83	2.44
	$M, \alpha_j, \widetilde{U}_j$	5.37	4.31	6.06	5.98
Welfare Loss	Arithmetic Mean	4.91	0.17	5.51	0.24
(est, std err)	Geometric Mean	5.37	0.22	5.84	0.37
	Harmonic Mean	5.85	0.30	6.35	0.51

8.4 Heteroskedasticity of U_0

While the previous subsection considered possible non normality of U_0, this subsection considers potential heteroskesdasticity in U_0. Heteroskedasticity of the Barten scales U_1 and U_2 is already modeled by the deterministic components $a_1(z)$ and $a_2(z)$. As noted in the main text, if our Barten scale model of taste variation is valid, then U_1 and U_2 should pick up a substantial portion of what would otherwise be unexplained variation, including heteroskedasticity, in demand. As a result, if our model is appropriate, then the nonbehavioral error term U_0 should be much smaller and more homoskedastic in Model 2 than in Model 1.

Table 2 showed that U_0 is indeed much smaller in Model 2, with an estimated standard deviation σ_0 falling from 0.666 in Model 1 to 0.469 in Model 2. To measure how much of the heteroskedasticity that would end up in U_0 is captured by the random coefficients $\widetilde{U}_j$, we consider a maximum likelihood analog to the Breusch-Pagan test. Specifically, we reestimate the models allowing the standard deviation of U_0, σ_0, to depend on normalized prices and demographics, and then examine the significance of these covariates in σ_0. For this test we replace σ_0 with a function $\sigma_0(X_1, X_2, Z)$ that is linear in the ten variables comprising, X_1, X_2, and Z. We then examine the estimated size and significance of the function $\sigma_0(X_1, X_2, Z)$ in Models 1 and 2.

In Model 1, the sample value of the Wald test of the joint significance of the coefficients comprising $\sigma_0(X_1, X_2, Z)$ is 327, while in Model 2 the Wald test statistic drops by over 50% to 145. So by this measure more than half of the heteroskedasticity in the residual U_0 in Model 1 is captured (and hence explained by) the preference heterogeneity embodied by our random Barten scale components $\widetilde{U}_j$. These Wald statistics remain above the critical values of the χ^2_{10} distribution at conventional levels, so although our model is big improvement over traditional Barten scales in this dimension, there remains some residual systematic heterogeneity that our model does not explain.[14]

Another way to measure how much heteroskedasticity is explained by random Barten scales is to compare the estimated values of $\sigma_0(X_1, X_2, Z)$ under the two models. In Model 1, the mean and standard deviation of $\widehat{\sigma}_{0i} = \widehat{\sigma}_0(X_{1i}, X_{2i}, Z_i)$ are 0.656 and 0.087, respectively. In Model 2, the mean and standard deviation of $\widehat{\sigma}_{0i}$ are 0.472 and 0.068, respectively. This again shows that random Barten scales explain a substantial portion of both the size and systematic components of otherwise unexplained variability in demand, by making U_0 both smaller and more homoskedastic. As in Tables 6 and 7, attempting to generalize our baseline model by modeling remaining heteroskedasticity in U_0 results in almost no change in the resulting economic analyses.

8.5 Endogeneity of X_j

All of our estimates treat normalized prices X_j as exogenous regressors. Our model has no constant term, and X_j does not enter Y linearly, so endogeneity with respect to the measure-

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[14]An assumption of the model from Theorem 1 is that U_0 is independent of X after conditioning on Z. In Model 1, the coefficients of X_1, X_2 in $\sigma_0(X_1, X_2, Z)$ are small in magnitude but highly statistically significant with t statistics over 5. In contrast, the coefficients of X_1, X_2 in $\sigma_0(X_1, X_2, Z)$ are even smaller and much less significant in Model 2, with t statistics of 1.04 and 2.52, respectively.

ment error U_0 is difficult to motivate. In contrast, endogeneity of X_j with respect to the unobserved preference heterogeneity parameters $\widetilde{U}_1$ and $\widetilde{U}_2$ is certainly possible. For example the distribution of $\widetilde{U}$ is assumed to be independent of regressors, so the presence of unobserved characteristics that correlate with both $\widetilde{U}_j$ and other regressors could induce bias, specifically bias in the estimates of the a_j functions (since the $\widetilde{U}_j$ terms are specifically the unobserved components of Barten scales and hence of preference heterogeneity, while the a_j functions are the observed components of Barten scales).

As discussed in the text, Canadian households are a small fraction of the world energy market, and so are unlikely to have much effect on energy prices. However, $\widetilde{U}_j$ could be correlated with X_j to the extent that the distribution of Canadian household preferences affects total Canadian household demand, and such demand contributes to differences between local (within Canada) energy prices and world energy prices. Similar arguments may apply to nonenergy prices, and measurement error in M could similarly cause correlations between the errors and regressors.

To test for the possibility that budgets M or prices P_1 and P_2 are endogenous, we try instrumenting these variables with gross annual income by source (for budgets), world oil price (for energy prices) and the US residential Case-Shiller index (for nonenergy prices), all interacted with observed demographics Z and a constant. The first stage for these regressions is very strong, with F statistics in excess of 100 in all 3 cases, so the instruments are not weak. We cannot test for exogeneity of these instruments, since we do not have observed residuals analogous to $\widetilde{U}_1$ and $\widetilde{U}_2$ with which to construct an overidentification test. However, we note that the use of income to instrument total expenditures M is standard in the consumer demand literature. The world oil price ought to be a valid instrument for energy prices in Canada because Canadian residential demand is such a small component of the world energy market. It is also a relatively strong instrument, especially since we interact it with Z which includes the calendar year, thereby obtaining sufficient variation in the prediction (R^2 is about 0.75 in this first stage regression). The US residential price index is likewise to be a good instrument for Canadian non-energy prices because Canadian residential prices are correlated with USA residential prices, but are not a big driver of them, and because local goods and service prices are highly correlated with residential prices (see, e.g., Moretti 2012).

Given the nonlinearities and nonseparable errors in our model, we use control function methods to test for endogeneity. As observed in, e.g., Wooldridge (2011), testing the significance of the coefficients of control function residuals provides a valid test of endogeneity, even when the resulting control function estimator does not satisfy all of the assumptions necessary to completely cure the potential endogeneity problem. This is relevant because endogeneity due to simultaneity of supply and demand does not in general lead to the triangular structures necessary for validity of control function based estimation.

We obtain control function residuals by regressing $\ln M$, $\ln P_1$ and $\ln P_2$ on our instrument vector and on demographic characteristics Z. We then include the residuals from these three regressions as control function regressors inside the functions α_1 and α_2. The sample value of the Wald test statistic for the hypothesis that all six control functions are zero is 713, and the likelihood ratio test statistic for the hypothesis is 124, which are both above conventional chi squared statistic critical values. We therefore cannot reject the hypothesis that no endogeneity is present, however, as with our other robustness checks, we can examine whether including these

control function residuals materially affects our estimates, and hence verify whether attempting to correct for potential endogeneity would be economically relevant. Note, however, that these estimates, which are provided in Table 8 in the same form as Tables 6 and 7, are only consistent if endogeneity takes the specific control function form.

Table 8: Estimates with Control Function Corrections

Parameter		Model 2 llf=-9706.88 Estimate	Std Err	Model 2 w ctrl fns llf=-9644.41 Estimate	Std Err
σ_0		0.469	*0.009*	0.394	*0.022*
σ_1		0.165	*0.036*	0.288	*0.030*
σ_2		1.336	*0.011*	1.208	*0.038*
ρ		0.883	*0.100*	0.465	*0.069*
α_1	$v_{\ln M}$			0.234	*0.037*
	$v_{\ln P_1}$			-0.187	*0.064*
	$v_{\ln P_2}$			-0.100	*0.169*
α_2	$v_{\ln M}$			0.337	*0.030*
	$v_{\ln P_1}$			-0.238	*0.028*
	$v_{\ln P_2}$			1.249	*0.084*
std dev	$\ln \alpha_1$	0.252		0.244	
	$\ln \alpha_2$	0.380		0.344	
	$\ln U_1$	0.298		0.372	
	$\ln U_2$	1.350		1.235	
correlation	$\ln \alpha_1, \ln \alpha_2$	-0.700		-0.444	
	$\ln U_1, \ln U_2$	0.293		0.258	
correlation:renter=0	$\ln \alpha_1, \ln \alpha_2$	0.105		0.464	
	$\ln U_1, \ln U_2$	0.699		0.456	
correlation: renter=1	$\ln \alpha_1, \ln \alpha_2$	0.087		0.459	
	$\ln U_1, \ln U_2$	0.691		0.444	
budget shares	$M, \alpha_j = \overline{\alpha}_j, P = 1, \widetilde{U}_j = 1$	0.147	0.030	0.147	0.036
(mean, sd)	$M, \alpha_j, P = 1, \widetilde{U}_j = 1$	0.148	0.048	0.148	0.051
	$M, \alpha_j, P, \widetilde{U}_j = 1$	0.151	0.047	0.151	0.051
	$M, \alpha_j, P, \widetilde{U}_j$	0.137	0.065	0.139	0.074
Cost-of-Living	$M, \alpha_j = \overline{\alpha}_j, \widetilde{U}_j = 1$	5.66	1.30	6.29	1.69
(mean, sd)	$M, \alpha_j, \widetilde{U}_j = 1$	5.64	1.69	6.28	2.08
	$M, \alpha_j, \widetilde{U}_j$	5.37	4.31	5.80	3.83
Welfare Loss	Arithmetic Mean	4.91	*0.17*	5.20	*0.18*
(mean, *std err*)	Geometric Mean	5.37	*0.22*	5.67	*0.22*
	Harmonic Mean	5.85	*0.30*	6.12	*0.31*

Most of the impact of the control function residuals is due to the log budget residual $v_{\ln M}$. This is consistent with our a priori belief that unobserved preference heterogeneity is unlikely to have much correlation with prices, but might well be correlated with the budgets of households, possibly due to measurement error in M. As in our previous robustness checks, the correction

for potential endogeneity does not change our economic analyses much. One difference is that the estimated correlation between $\ln\alpha_1$ and $\ln\alpha_2$ is increased, but this is offset by an estimated decrease in the correlation between $\ln\widetilde{U}_1$ and $\ln\widetilde{U}_2$, leaving the estimated correlation between the Barten scales $\ln U_1$ and $\ln U_2$ little changed.

As with our experiments estimating more complicated error distributions, to the extent that controlling for potential endogeneity differs from baseline, all of the results in this alternative specification strengthen rather than weaken our economic conclusions. For example, the estimated variation in cost-of-living impacts due to a 50% increase in the price of energy is larger when including control function residual than in the baseline model, and the extent to which increasing inequality aversion increases welfare losses is also increased a small amount.

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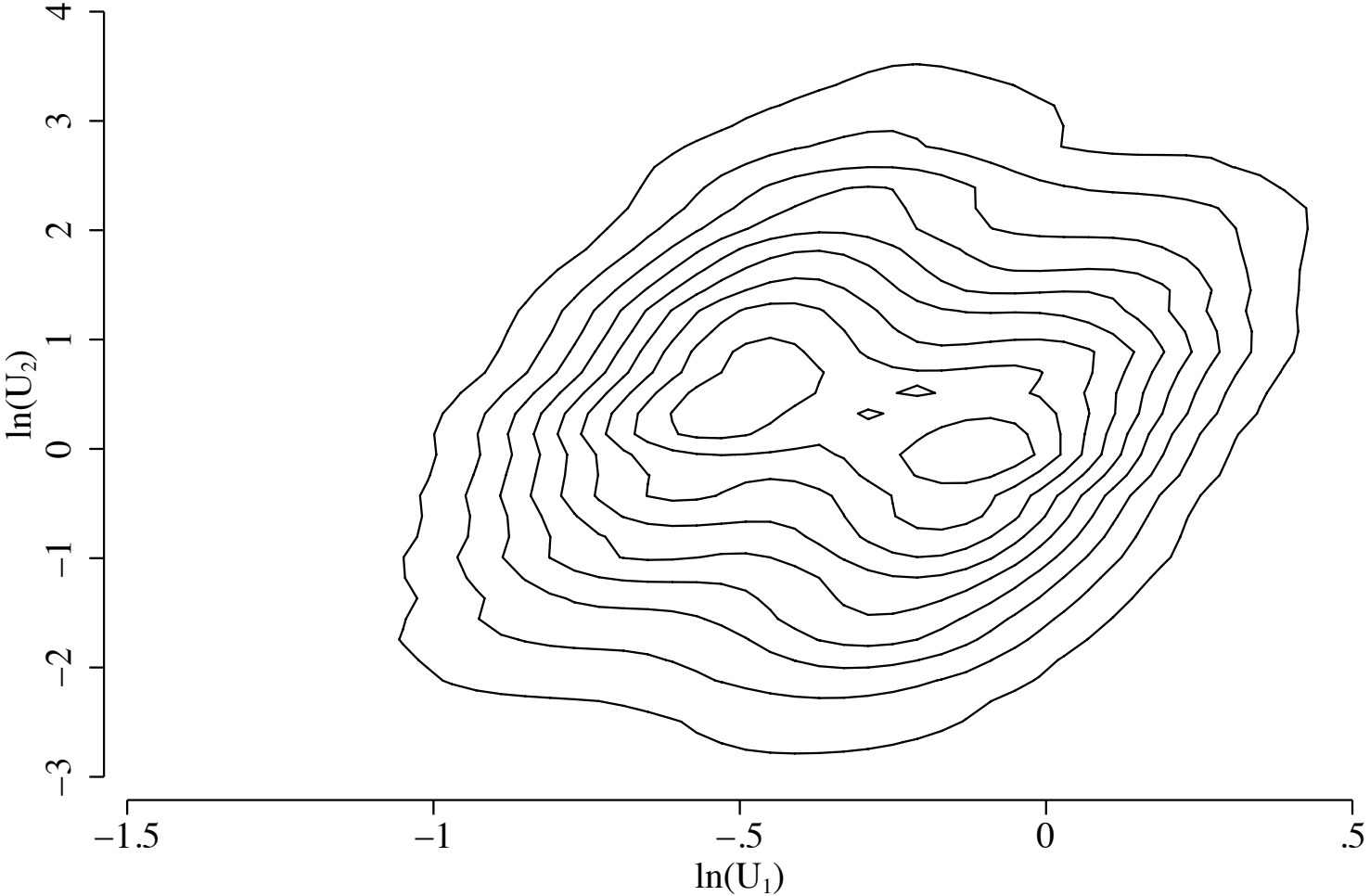
Estimated Distribution of $\ln(\alpha_1), \ln(\alpha_2)$: Model 1
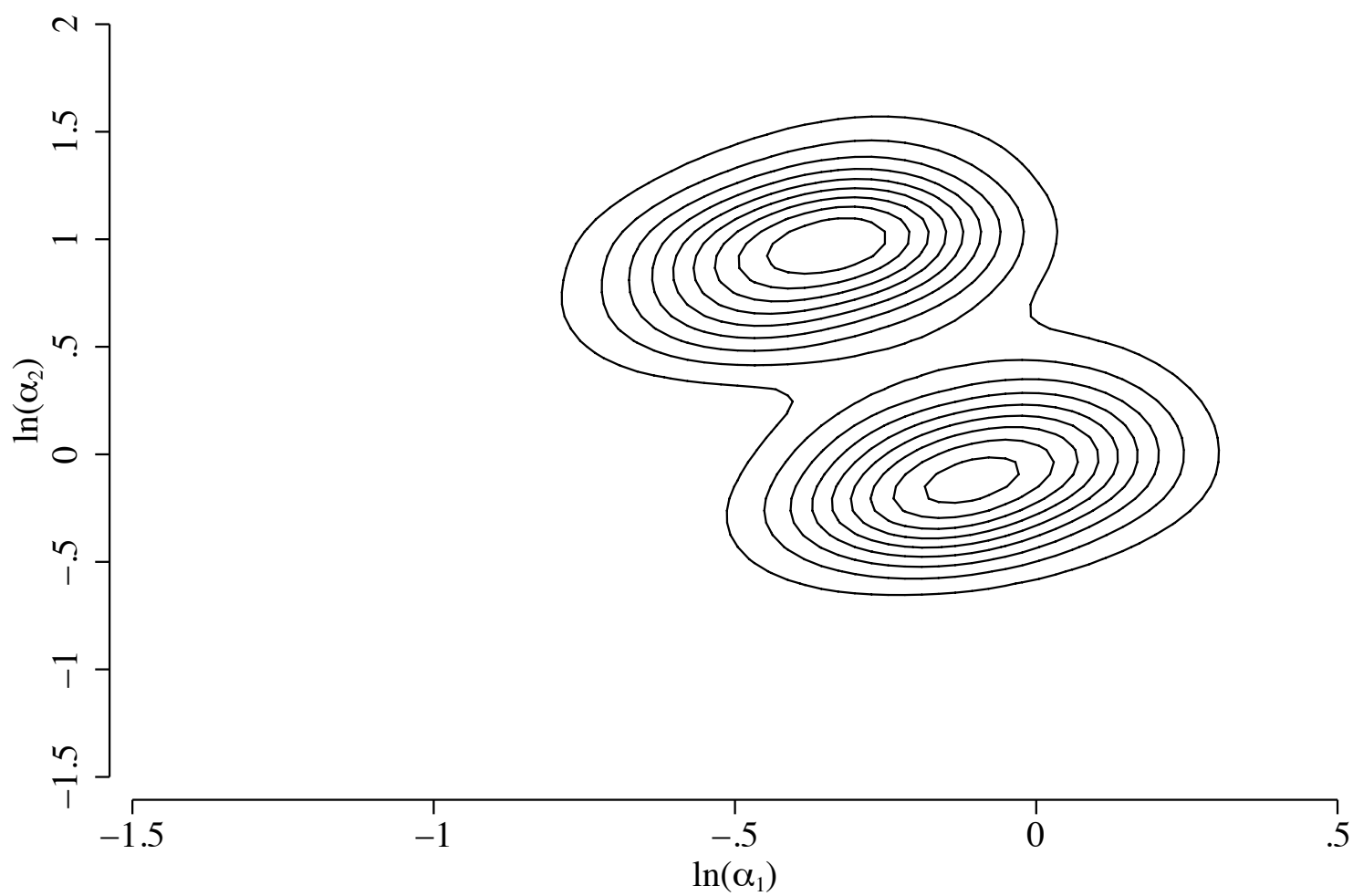
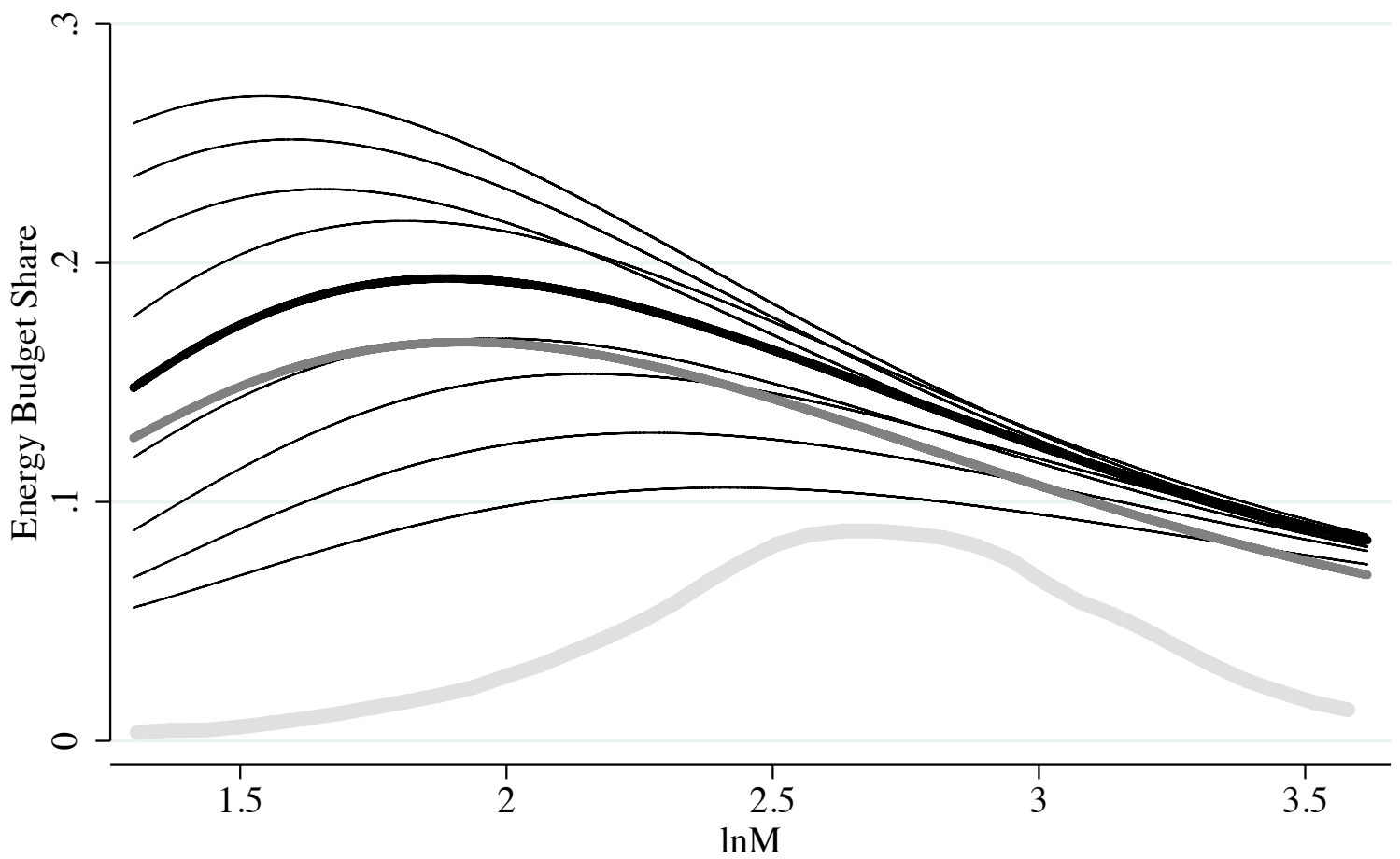Estimated Distribution of $\ln(\alpha_1), \ln(\alpha_2)$: Model 2
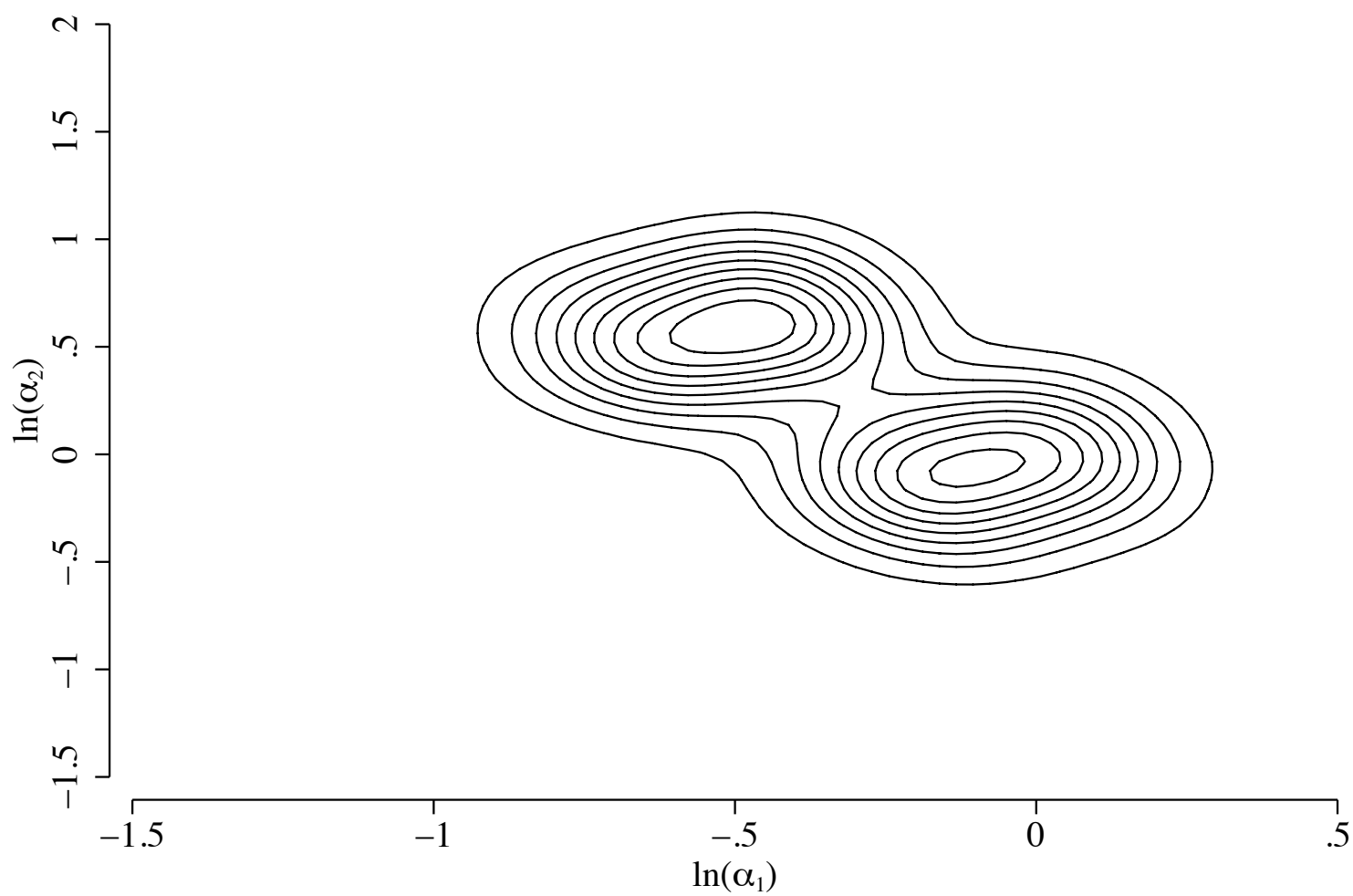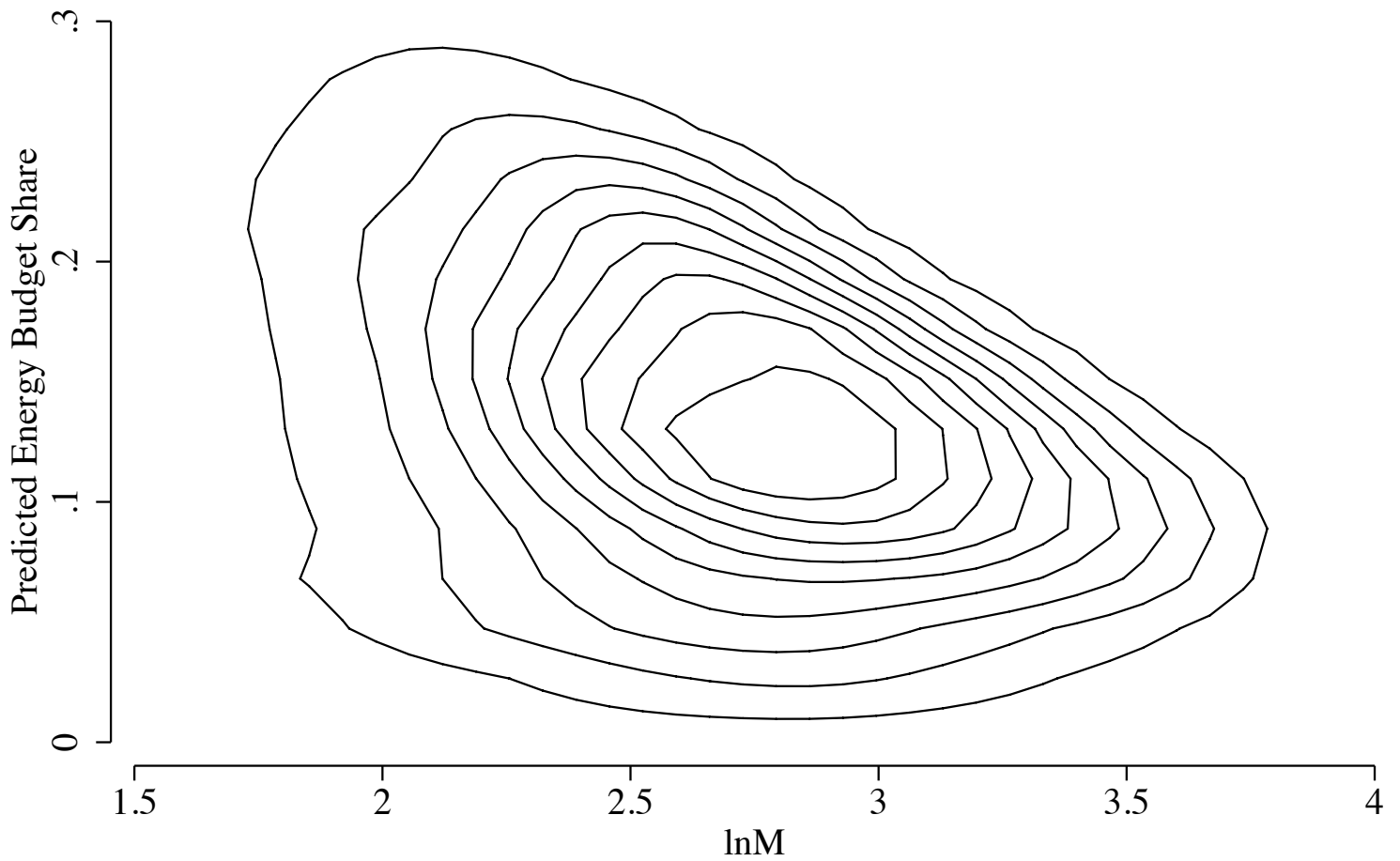
Estimated Distribution of $\ln(U_1),\ln(U_2)$: Model 2

Estimated Budget Shares, Models 1 and 2

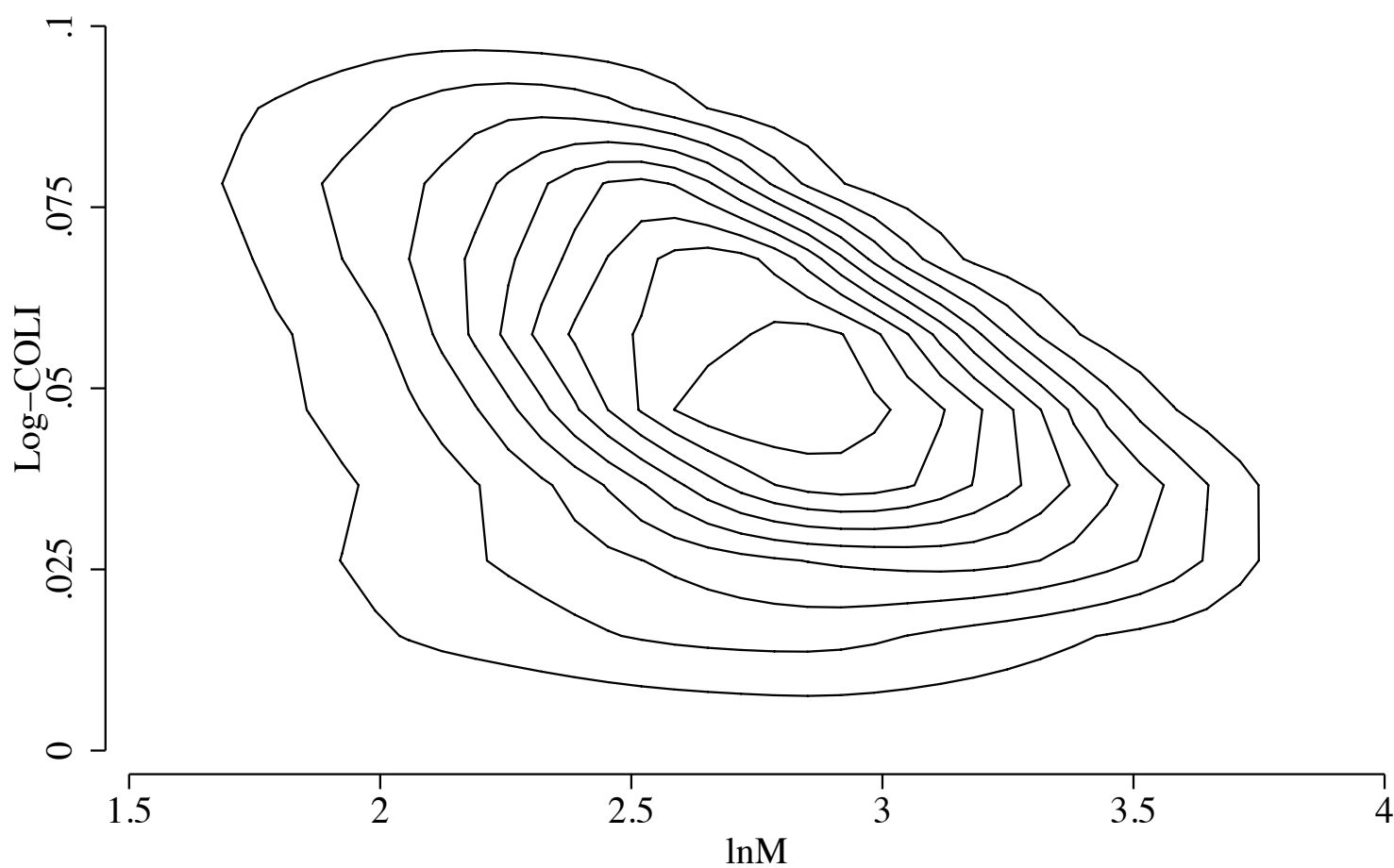Model 2 at quartiles of U_1, U_2; evaluated at base prices and mean α_j

Distribution of Predicted Budget Shares, Model 2
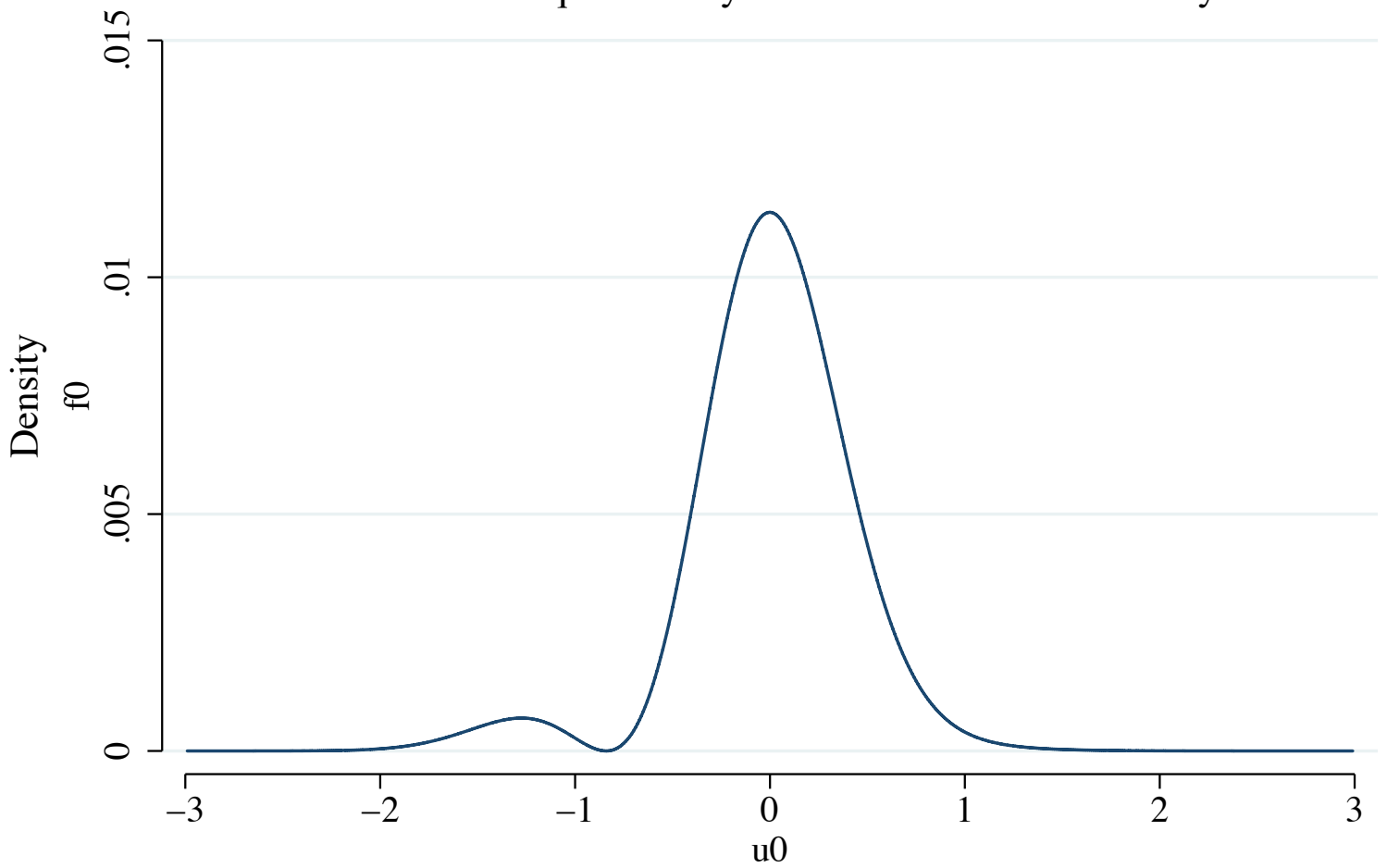at base prices and estimated α_j and U_j distribution

Distribution of Log Cost of Living Impacts, Model 2

Given base prices, 50% increase in Energy Price, and estimated α_j and U_j

Estimated Distribution of U0: Model 2

Third Order Squared Polynomial times Normal Density

Estimated Distribution of $\ln(U_1), \ln(U_2)$: Model 2

Second Order Squared Polynomial times Bivariate Normal Density