Identification and estimation of average partial effects in ‘irregular’ correlated random coefficient panel data models, supplemental material: Additional proofs and application details

This supplemental web appendix contains proofs of some auxiliary Lemmas used to show Theorem 2.1. It also contains a proof of Theorem 2.2 and additional details regarding the empirical application. All notation is as established in the main paper unless noted otherwise. Equation numbering continues in sequence with that established in the main paper. References not included in the bibliography to the main paper are listed below.

B Additional proofs

Proof of Lemma A.3 To verify this result, we must first establish that the defined limit exists. For h_N sufficiently small, we can decompose the expectation Ξ_N as

$$\begin{aligned}\Xi_N &= \mathbb{E}\left[\mathbf{1}\left(|D_i| \geq u_0\right)\mathbf{X}_i^{-1}\mathbf{W}_i\right] + \mathbb{E}\left[\mathbf{1}\left(h_N < |D_i| < u_0\right)\mathbf{X}_i^{-1}\mathbf{W}_i\right]\\ &= O(1) + \mathbb{E}\left[\mathbf{1}\left(h_N < |D_i| < u_0\right)D_i^{-1}\mathbf{W}_i^*\right]\\ &= O(1) + \int_{h_N}^{u_0}\left(\frac{\xi(u)-\xi(-u)}{u}\right)du,\end{aligned}$$

where

$$\xi(u) \equiv \phi(u)\mathbb{E}[W_i^*|D_i = u]$$

is twice continuously differentiable for $|u| < u_0$ by Assumption 1.2. Using the Taylor's series expansion

$$\xi(u) - \xi(-u) = \xi(0) - \xi(0) + 2\frac{d\xi(0)}{du}\cdot u + \left[\frac{d^2\xi(u^*)}{du^2} - \frac{d^2\xi(-u^*)}{du^2}\right]\cdot\left(\frac{u^2}{2}\right),$$

for u^* some intermediate value between h_N and u_0, it follows that

$$\left|1\left(h_N < u < u_0\right)\cdot\left(\frac{\xi(u)-\xi(-u)}{u}\right)\right| \leq 1(u \leq u_0)\left[2\left\|\frac{d\xi(0)}{du}\right\| + \max_{|u|\leq u_0}\left\|\frac{d^2\xi(u)}{du^2}\right\|\cdot u_0\right],$$

and since the right-hand side is integrable, $\Xi_N \to \Xi_0$ by dominated convergence. Then, taking λ to be an arbitrary (fixed) q-vector, verification that $\widehat{\Xi}_N \xrightarrow{p} \Xi_0$ follows from the convergence of the covariance matrix of the numerator of $\widehat{\Xi}_N\lambda$ to zero:

$$\begin{aligned}\left\|\mathbb{V}\left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| > h_N\right)\mathbf{X}_i^{-1}\mathbf{W}_i\lambda\right]\right\| &= \left\|\mathbb{V}\left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| > h_N\right)D_i^{-1}\mathbf{W}_i^*\lambda\right]\right\|\\ &\leq \frac{2\left\|\lambda\right\|^2}{N}\mathbb{E}\left[\mathbf{1}\left(|D_i| > h_N\right)|D_i|^{-2}\left\|\mathbf{W}_i^*\right\|^2\right]\\ &= \frac{2\left\|\lambda\right\|^2}{N}\left[\mathbb{E}\left[\mathbf{1}\left(|D_i| \geq u_0\right)|D_i|^{-2}\left\|\mathbf{W}_i^*\right\|^2\right]\right.\\ &\qquad\left.+\mathbb{E}\left[\mathbf{1}\left(h_N < |D_i| < u_0\right)|D_i|^{-2}\left\|\mathbf{W}_i^*\right\|^2\right]\right]\\ &= O\left(\frac{1}{N}\right) + \frac{2\left\|\lambda\right\|^2}{N}\int_{h_N}^{u_0}\left(\frac{\zeta(u)+\zeta(-u)}{u^2}\right)du,\end{aligned}$$

where $\|\mathbf{W}_i^*\|^2 \equiv tr\left[\mathbf{W}^{*\prime}\mathbf{W}^*\right]$ and

$$\zeta(u) \equiv \phi(u)\mathbb{E}[\|W_i^*\|^2 \,|D_i = u]$$

is bounded for $|u| < u_0$, so

$$\begin{aligned}
\left\|\mathbb{V}\left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| > h_N\right)\mathbf{X}_i^{-1}\mathbf{W}_i\lambda\right]\right\| &\leq O\left(\frac{1}{N}\right) + \frac{4\,\|\lambda\|^2}{N}\left(\max_{|u|\leq u_0}\|\zeta(u)\|\right)\int_{h_N}^{u_0}\left(\frac{1}{u^2}\right)du \\
&= O\left(\frac{1}{N}\right) + \frac{4\,\|\lambda\|^2}{N}\left(\max_{|u|\leq u_0}\|\zeta(u)\|\right)\left[\frac{1}{h_N} - \frac{1}{u_0}\right] \\
&= O\left(\frac{1}{Nh_N}\right) \\
&= o(1)
\end{aligned}$$

under Assumption 2.5.

Proof of Theorem 2.2: Rewriting equation (30) we have

$$\widehat{V} = \left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{R}_i\right]^{-1} \times \left[\frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\widehat{\mathbf{U}}_i^{+}\widehat{\mathbf{U}}_i^{+\prime}\mathbf{Q}_i\right] \times \left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{R}_i\right]^{-1} \tag{54}$$

for $\mathbf{U}_i^{+} = \mathbf{Y}_i^* - \mathbf{R}_i\boldsymbol{\theta}_0$, with $\widehat{\boldsymbol{\theta}} = \left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{R}_i\right]^{-1} \times \left[\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{Y}_i^*\right]$, for $\boldsymbol{\theta} = (\boldsymbol{\delta}', \boldsymbol{\beta}')'$ and

$$\underset{T\times q+p}{\mathbf{Q}_i} = \left(h_N^{-1}\mathbf{1}\left(|D_i| \leq h_N\right)\mathbf{W}_i^*, \frac{\mathbf{1}\left(|D_i| > h_N\right)}{D_i}I_p\right), \quad \underset{T\times q+p}{\mathbf{R}_i} = \left(\mathbf{W}_i^*, \mathbf{1}\left(|D_i| > h_N\right)D_iI_p\right).$$

The dependence of $\mathbf{Q}_i$ and $\mathbf{R}_i$ on h_N is suppressed to simplify the notation.

Starting with the 'Jacobian' term in $\widehat{V}$ we get, by the definitions of $\mathbf{Q}_i$, $\mathbf{R}_i$, and $\widehat{\Xi}_N$,

$$\begin{aligned}
\frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{R}_i &= \frac{1}{N}\sum\nolimits_{i=1}^{N}\begin{pmatrix} h_N^{-1}\mathbf{1}\left(|D_i| \leq h_N\right)\mathbf{W}_i^{*\prime} \\ \frac{\mathbf{1}(|D_i|>h_N)}{D_i}I_p \end{pmatrix}\begin{pmatrix} \mathbf{W}_i^* & \mathbf{1}\left(|D_i| > h_N\right)D_iI_p \end{pmatrix} \\
&= \begin{pmatrix} \frac{1}{Nh_N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| \leq h_N\right)\mathbf{W}_i^{*\prime}\mathbf{W}_i^* & \underline{0}_q\underline{0}_p' \\ \frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| > h_N\right)\mathbf{X}_i^{-1}\mathbf{W}_i & \frac{1}{N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| > h_N\right)I_p \end{pmatrix} \\
&\xrightarrow{p} \begin{pmatrix} 2\mathbb{E}\left[\mathbf{W}_i^{*\prime}\mathbf{W}_i^*\,|\,D_i = 0\right]\phi_0 & \underline{0} \\ \Xi_0 & I_p \end{pmatrix}
\end{aligned}$$

by (50) and A.3. Decomposing the middle term $h_N\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\widehat{\mathbf{U}}_i^{+}\widehat{\mathbf{U}}_i^{+\prime}\mathbf{Q}_i/N$, yields

$$\begin{aligned}
&\left\|\frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\widehat{\mathbf{U}}_i^{+}\widehat{\mathbf{U}}_i^{+\prime}\mathbf{Q}_i - \frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{U}_i^{+}\mathbf{U}_i^{+\prime}\mathbf{Q}_i\right\| \\
\leq\ & \frac{2h_N}{N}\sum\nolimits_{i=1}^{N}\left\|\mathbf{U}_i^{+}\right\|\|\mathbf{Q}_i\|^2\left\|\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}_0\right\| + \frac{h_N}{N}\sum\nolimits_{i=1}^{N}\|\mathbf{Q}_i\|^2\left\|\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}_0\right\|^2 \\
=\ & \frac{2}{Nh_N}\sum\nolimits_{i=1}^{N}\left(\mathbf{1}\left(|D_i| \leq h_N\right)\|\mathbf{W}_i^*\|^2\left\|\mathbf{U}_i^{+}\right\| + \mathbf{1}\left(|D_i| > h_N\right)\frac{h_N^2\left\|\mathbf{U}_i^{+}\right\|}{|D_i|^2}\right)\left\|\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}_0\right\| \\
&+\frac{1}{Nh_N}\sum\nolimits_{i=1}^{N}\left(\mathbf{1}\left(|D_i| \leq h_N\right)\|\mathbf{W}_i^*\|^2 + \mathbf{1}\left(|D_i| > h_N\right)\frac{h_N^2}{|D_i|^2}\right)\left\|\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}_0\right\|^2.
\end{aligned} \tag{55}$$

By a similar argument as for (49) above,

$$\begin{aligned}\mathbb{E}\left[h_N^{-1}\mathbf{1}\left(|D_i| \leq h_N\right)\|\mathbf{W}_i^*\|^2\left(\left\|\mathbf{U}_i^+\right\|+1\right)\right] &= \mathbb{E}\left[h_N^{-1}\mathbf{1}\left(|D_i| \leq h_N\right)\mathbb{E}\left[\|\mathbf{W}_i^*\|^2\left(\left\|\mathbf{U}_i^+\right\|+1\right)\ \mid D_i\right]\right] \\ &\rightarrow 2\mathbb{E}\left[\|\mathbf{W}_i^*\|^2\left(\left\|\mathbf{U}_i^+\right\|+1\right)\ \mid D_i = 0\right]\phi_0,\end{aligned}$$

and the same reasoning as (43) yields

$$\begin{aligned}\mathbb{E}\left[\mathbf{1}\left(|D_i| > h_N\right)\|\mathbf{W}_i^*\|^2\left(\left\|\mathbf{U}_i^+\right\|+1\right)\frac{h_N}{|D_i|^2}\right] &= \mathbb{E}\left[h_N^{-1}\mathbf{1}\left(|D_i| > h_N\right)\frac{h_N}{|D_i|^2}\mathbb{E}\left[\|\mathbf{W}_i^*\|^2\left(\left\|\mathbf{U}_i^+\right\|+1\right)\ \mid D_i\right]\right] \\ &\rightarrow 2\mathbb{E}\left[\|\mathbf{W}_i^*\|^2\left(\left\|\mathbf{U}_i^+\right\|+1\right)\ \mid D_i = 0\right]\phi_0.\end{aligned}$$

Thus, by Markov's inequality, (55) yields

$$\begin{aligned}\frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\widehat{\mathbf{U}}_i^+\widehat{\mathbf{U}}_i^{+\prime}\mathbf{Q}_i &= \frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{U}_i^+\mathbf{U}_i^{+\prime}\mathbf{Q}_i + O_p\left(\left\|\hat{\boldsymbol{\theta}}-\boldsymbol{\theta}_0\right\|\right) \\ &= \frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{U}_i^+\mathbf{U}_i^{+\prime}\mathbf{Q}_i + o_p\left(\mathbf{1}\right).\end{aligned}$$

Finally,

$$\frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{U}_i^+\mathbf{U}_i^{+\prime}\mathbf{Q}_i = \begin{pmatrix} \frac{1}{Nh_N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| \leq h_N\right)\mathbf{W}_i^{*\prime}\left(\mathbf{Y}_i^*-\mathbf{W}_i^*\boldsymbol{\delta}_0\right)\left(\mathbf{Y}_i^*-\mathbf{W}_i^*\boldsymbol{\delta}_0\right)'\mathbf{W}_i^* & \underline{0}_p\underline{0}_q' \\ \underline{0}_q\underline{0}_p' & \frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{1}\left(|D_i| > h_N\right)\left\{\mathbf{X}_i^{-1}\left(\mathbf{Y}_i-\mathbf{W}_i\boldsymbol{\delta}_0\right)-\boldsymbol{\beta}_0\right\}\left\{\mathbf{X}_i^{-1}\left(\mathbf{Y}_i-\mathbf{W}_i\boldsymbol{\delta}_0\right)-\boldsymbol{\beta}_0\right\}' \end{pmatrix}.$$

By the calculations yielding (53), the expected value of the first diagonal submatrix is

$$\mathbb{E}\left[h_N^{-1}\mathbf{1}\left(|D_i| \leq h_N\right)\mathbf{W}_i^{*\prime}\left(\mathbf{Y}_i^*-\mathbf{W}_i^*\boldsymbol{\delta}_0\right)\left(\mathbf{Y}_i^*-\mathbf{W}_i^*\boldsymbol{\delta}_0\right)'\mathbf{W}_i^*\right] = 2\mathbb{E}\left[\mathbf{W}_i^{*\prime}\mathbf{X}^*\Sigma\left(\mathbf{X}\right)\mathbf{X}^{*\prime}\mathbf{W}_i^*\middle| D_i = 0\right]\phi_0 + o\left(1\right),$$

while the variance of any term in the matrix is $O\left((Nh_N)^{-1}\right) = o(1)$ by Assumptions 2.3 and 2.5 (with $r \leq 8$). Similarly, the expectation of the second diagonal submatrix is

$$\begin{aligned}\mathbb{E}\left[h_N\ \mathbf{1}\left(|D_i| > h_N\right)\left\{\mathbf{X}_i^{-1}\left(\mathbf{Y}_i-\mathbf{W}_i\boldsymbol{\delta}_0\right)-\boldsymbol{\beta}_0\right\}\left\{\mathbf{X}_i^{-1}\left(\mathbf{Y}_i-\mathbf{W}_i\boldsymbol{\delta}_0\right)-\boldsymbol{\beta}_0\right\}'\right] &= 2\mathbb{E}\left[\mathbf{X}^*\Sigma\left(\mathbf{X}\right)\mathbf{X}^{*\prime}\middle| D = 0\right]\phi_0 + o\left(1\right) \\ &= 2\Upsilon_0\phi_0 + o(1),\end{aligned}$$

while similar calculations to those leading to (45) yield

$$\begin{aligned}\left\|\mathbb{V}\left[h_N\ \mathbf{1}\left(|D_i| > h_N\right)\left\{\mathbf{X}_i^{-1}\left(\mathbf{Y}_i-\mathbf{W}_i\boldsymbol{\delta}_0\right)-\boldsymbol{\beta}_0\right\}\left\{\mathbf{X}_i^{-1}\left(\mathbf{Y}_i-\mathbf{W}_i\boldsymbol{\delta}_0\right)-\boldsymbol{\beta}_0\right\}'\right]\right\| &\leq \frac{4m_8(0)\phi_0}{Nh_n} + o\left(\frac{1}{Nh_n}\right) \\ &= o(1),\end{aligned}$$

where $m_8(u)$ is defined in Assumption 2.3, Thus

$$\frac{h_N}{N}\sum\nolimits_{i=1}^{N}\mathbf{Q}_i'\mathbf{U}_i^{+}\mathbf{U}_i^{+\prime}\mathbf{Q}_i \xrightarrow{p} \begin{pmatrix} 2\mathbb{E}\left[\mathbf{W}_i^{*\prime}\mathbf{X}^{*}\Sigma\left(\mathbf{X}\right)\mathbf{X}^{*\prime}\mathbf{W}_i^{*}\right|D_i=0\right]\phi_0 & \underline{0}_q\underline{0}_p' \\ \underline{0}_p\underline{0}_q' & 2\Upsilon_0\phi_0 \end{pmatrix},$$

ensuring $\widehat{V}\xrightarrow{p}V_0$.

C Additional details on empirical application

Data description: The preparation of our estimation sample from the raw public release data files involved some complex and laborious data-processing. We outline the procedures used to construct our sample in this appendix. A sequence of heavily commented STATA do files, which read in the IFPRI (2005) release of the data and output a text file of our estimation sample are available online at `https://files.nyu.edu/bsg1/public/`.

As noted in the main text, we use data collected in conjunction with an external evaluation of the Nicaraguan conditional cash transfer program Red de Protección Social (RPS) (see IFPRI, 2005). The RPS evaluation sample is a panel of 1,581 households from 42 rural communities in the departments of Madriz and Matagalpa, located in the northern part of the Central Region of Nicaragua. Each sampled household was first interviewed in August/September 2000 with follow-ups attempted in October of both 2001 and 2002. A total of 1,359 households were successfully interviewed in all three waves. One of these households reports zero food expenditures (and hence calorie availability) in one wave and is dropped from our sample. Our estimation sample therefore consists of a balanced panel of 1,358 households from all three waves.

The survey was fielded using an abbreviated version of the 1998 Nicaraguan Living Standards Measurement Survey (LSMS) instrument. As such it includes a detailed consumption module with information on household expenditure, both actual and in kind, on 59 specific foods and several dozen other common budget categories (e.g., housing and utilities, health, education, and household goods). The responses to these questions were combined to form an annualized consumption aggregate, C_{it}. In forming this variable we followed the algorithm outlined by Deaton and Zaidi (2002).

In addition to recording food expenditures, actual quantities of foods acquired are available. Using conversion factors listed in the World Bank (2002) and Instituto Nacional de Estadísticas y Censos (2005) (henceforth INEC) we converted all food quantities into grams. We then used the caloric content and edible percent information in the Instituto de Nutrición de Centro América y Panamá (2000) (henceforth INCAP) food composition tables to construct a measure of daily total calories available for each household. In forming our measure of calorie availability we followed the general recommendations of Smith and Subandoro (2007). The logarithm of this measure divided by household size, Y_{it}, serves as the dependent variable in our analysis.

The combination of both expenditure and quantity information at the household-level also allowed us to estimate unit prices for foods. These unit values were used to form a Paasche cost-of-living index for the i^{th} household in year t of

$$I_{it}=\left[S_{it}\left\{\sum\nolimits_{f=1}^{F}W_{f,it}\left(P_f^b/P_{f,it}\right)\right\}+\left(1-S_{it}\right)J_{it}\right]^{-1}, \tag{56}$$

where S_{it} is the fraction of household spending devoted to food, $W_{f,it}$ the share of overall food spending devoted to the f^{th} specific food, $P_{f,it}$ the year t unit price paid by the household for food f, and P_f^b its 'base' price (equal to the relevant 2001 sample median price). We use 2001 as our base year since it facilitates comparison with information

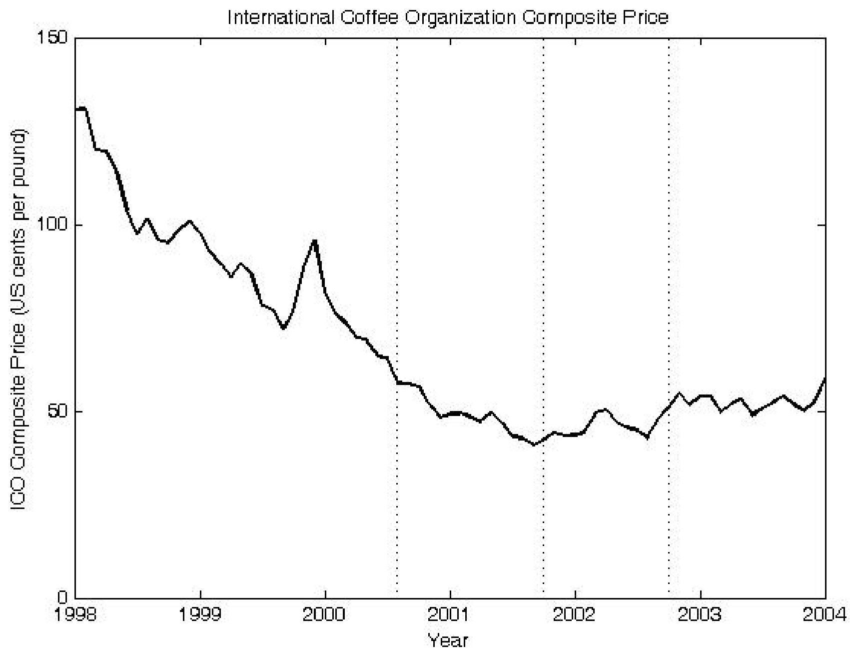

Figure 2: International Coffee Organization Composite Price Index, 1998 to 2004
NOTES: Source data downloaded from http://www.ico.org/ on 21 June 2011.

from a nationwide LSMS survey fielded that year. Following the suggestion of Deaton and Zaidi (2002) we replace household-level unit prices with village medians in order to reduce noise in the price data. In the absence of price information on nonfood goods we set J_{it} equal to one in 2001 and to the national consumer price index (CPI) in 2000 and 2002. Our independent variable of interest is real per capita consumption in thousands of Cordobas: $\mathrm{Exp}_{it} = ([C_{it}/I_{it}])/M_{it}$; M_{it} is total household size.

A nonlinear model: As we have three periods of data we can modify our model to allow the calorie elasticity to vary non-linearly with income. Nonlinearity in the calorie demand curve has been emphasized by Strauss and Thomas (1990, 1995) and Subramanian and Deaton (1996). We consider the model

$$\ln(\mathrm{Cal}_t) = b_{0t}(A, U_t) + b_{1t}(A, U_t)\ln(\mathrm{Exp}_t) + b_{2t}(A, U_t)\mathrm{Exp}_t^{-1},$$

so that a household's period-specific demand elasticity is given by $b_{1t}(A, U_t) - b_{2t}(A, U_t)\mathrm{Exp}_t^{-1}$. We estimate the average of this elasticity in 2000, 2001 and 2002 using the approach outlined in Section 3. Unlike in our linear analysis, we only allow for common intercept shifts across periods when fitting the nonlinear model. Figure 3 plots the a histogram of D with $\mathbf{X}_t = \left(1, \ln(\mathrm{Exp}_t), \mathrm{Exp}_t^{-1}\right)'$. Because $\ln(\mathrm{Exp}_t)$ and Exp_t^{-1} are highly correlated within-units, the density of D is substantial in the neighborhood of zero. The extreme 'irregularity' of the augmented model suggests that its estimation will require a substantial amount of trimming.

Table 4 reports average elasticity estimates based on the extended model. The average elasticities associated with the OLS and FE-OLS parameter estimates of the nonlinear model are virtually identical to their linear model

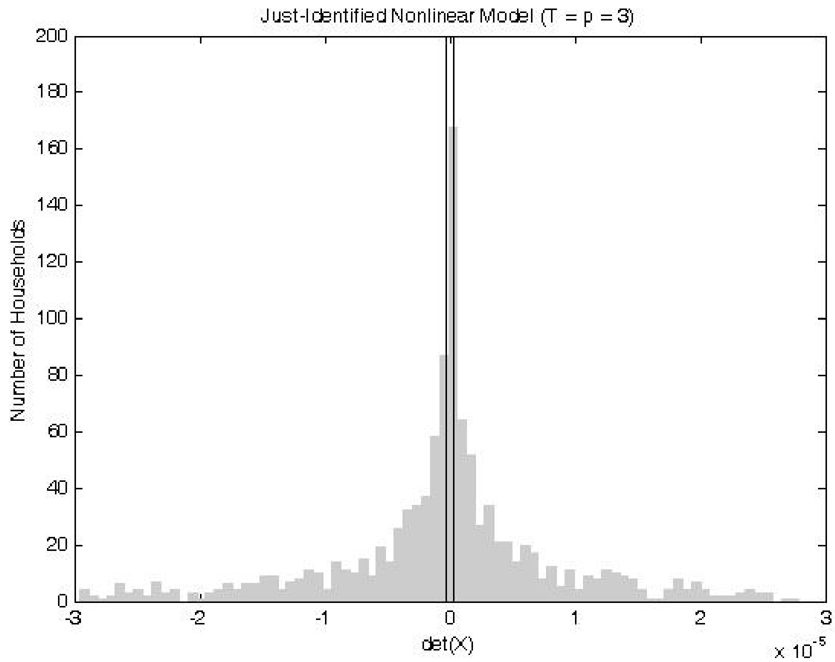

Figure 3: Histogram of the distribution of D ($T = p = 3$)
NOTES: The smallest and largest 10 percent of the D_i's are excluded from the histogram. The two vertical blue lines correspond to the portion of the sample that is trimmed in our preferred estimates (Table 4, Column 3).

	Calorie Demand Elasticities				
	(1)	(2)	(3)	(4)	(5)
	OLS	FE	I-CRC	I-CRC	I-CRC
2000 Elasticity	0.6226	0.6841	0.3720	0.2463	0.2442
	(0.0173)	(0.0329)	(0.2918)	(0.3984)	(0.2737)
2001 Elasticity	0.6269	0.6872	0.6882	1.4738	0.7057
	(0.0181)	(0.0340)	(0.2825)	(0.3777)	(0.2780)
2002 Elasticity	0.6274	0.6875	0.4641	0.2010	0.5689
	(0.0182)	(0.0341)	(0.3093)	(0.4340)	(0.2938)
Percent trimmed	–	–	12.3	5	15
Intercept shifts	Yes	Yes	Yes	Yes	Yes
Slope shifts	No	No	No	No	No

Table 4: Estimates of the calorie Engel curve: nonlinear case
NOTES: Estimates based on the balanced panel of 1,358 households described in the main text. "OLS" denotes least squares applied to the pooled 2000, 2001, and 2002 samples, "FE-OLS" least squares with household-specific intercepts, and "I-CRC" our irregular correlated random coefficients estimator (now using all three waves). All models include common intercept, but not slope, shifts across periods. The standard errors are computed in a way that allows for arbitrary within-village correlation in disturbances across households and time. The average elasticity estimates in the OLS and FE-OLS columns are computed using the delta method. Those in the I-CRC columns as described in Section 3.

Table 3 counterparts. Although the coefficient on Exp_t^{-1} is significant in both models (not reported), the effect of its inclusion on the average elasticity estimates is negligible. Column 3 reports I-CRC estimates with $h_N = \frac{c_D}{2} N^{-1/3}$ (which, given the large density in the neighborhood of zero, results in the trimming of 12 percent of the sample). In contrast to the linear case, the I-CRC estimates are imprecisely determined; they are also more sensitive to variations in the bandwidth (Columns 3 and 4). We conclude that we are unable to reliably fit the nonlinear CRC model with the data available.

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