Household Attitudes to Price Risk with Multiple Commodities: Evidence from Rural Ethiopia*

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Abstract

We study household attitude with respect to price risk. Whereas price risk aversion has so far been studied empirically only for single staple commodities, we expand the analytical framework so as to derive an estimable matrix of own- and cross-price risk aversion coefficients. This imposes strong restrictions on the matrix of price risk aversion coefficients, which has a complex relation to the household's Slutsky substitution matrix. Using a panel of rural Ethiopian households, we test whether the restrictions implied by the theory hold empirically, as well as whether distinct patterns of price risk aversion emerge. We ultimately find strong empirical support for the theory and widespread support for the hypothesis that households are on average risk-averse over own- and cross-price fluctuations.

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"[*T*]*he tendency of these measures to focus directly or indirectly on price, which, as stated, is the greatest source of uncertainty, has led economists to regard the management of prices as being of unique importance. And they have far more frequently related such management to the maximization of profits than to the minimization of risks. That is unfortunate, for the development of the modern business enterprise can be understood only as a comprehensive effort to reduce risk.*"

– John Kenneth Galbraith, *The Affluent Society* (1958).

1. Introduction

The effects of price risk on the behavior of competitive firms have been well-explored in the theoretical literature (Baron, 1970; Sandmo, 1971). Yet, when it comes to individuals and households, economists have largely ignored the effects of price risk, focusing instead on income risk. To be sure, the analysis of commodity price risk has been extended theoretically to individuals (Deschamps, 1973; Hanoch, 1977; Turnovsky et al., 1980; Newbery and Stiglitz, 1981) and to agricultural households both in theory (Finkelshtain and Chalfant, 1991) and in practice (Barrett, 1996), but these analyses have largely focused on one staple good. Turnovsky et al. (1980) considered price risk over multiple commodities, but they only did so theoretically. In this paper, we combine the empirical Finkelshtain-Chalfant-Barrett and the theoretical Turnovsky et al. frameworks to (i) derive an estimable matrix of coefficients of price risk aversion and relate it to the Slutsky matrix; (ii) define and estimate absolute and relative price risk aversion; and (iii) test the implications of the theory as it applies to agricultural households and the more general and realistic case of multiple commodities.

We use the general framework of agricultural household models (Singh et al., 1986) because it allows for households to be net buyers, net sellers, or autarkic with respect to goods, relaxing the stronger assumptions of the pure theories of the consumer or firm.

Because a household's indirect utility function is defined over the *vector* of prices for the commodities it produces or consumes and the household's income, it is possible to derive and estimate a matrix of price risk aversion coefficients, i.e., a matrix that reflects how price risk premia with respect to one good change with respect to changes in the price of any other good. For each of the commodities observed by the econometrician, such a matrix yields the usual (own-) price risk aversion coefficients on the diagonal, but also the cross-price risk aversion coefficients off the diagonal. To our knowledge, these off-diagonal terms have hitherto been overlooked in the empirical literature, although they have an intuitive interpretation and are important to understanding behavior with respect to price risk.

Based on the theoretical work of Turnovsky et al. (1980) and on an extension of Barrett's (1996) empirical work to the multiple commodity case, we derive the pseudo-Slutsky matrix of price risk aversion as well as its properties in section 2 and establish its formal relationship with the Slutsky matrix. In section 3, we discuss the data and present some descriptive statistics. We then develop a simple empirical framework to estimate the price risk aversion coefficients in section 4. In section 5 we estimate several marketable surplus equations, use their results to estimate own- and cross-price risk aversion coefficients, use these coefficients to construct the matrix of price risk aversion coefficients, and finally test the restrictions implied by the theory, progressively imposing more structure as a way of conducting robustness checks. We conclude by discussing the implications of our findings in section 6.

2. Theoretical Framework

This section develops a simple unitary agricultural household model (AHM) and derives the household's matrix of own- and cross-price risk aversion coefficients for the multiple commodity case. We then derive some key properties of this matrix and relate it to the Slutsky matrix., which yields the implications that we test in section 5.

2.1 Agricultural Household Model

The derivations in this section closely follow those in Barrett (1996), which built on Turnovsky et al.'s (1980) work on individual consumers and Finkelshtain and Chalfant's (1991) work on AHM.

Consider an agricultural household whose preferences are represented by a von Neumann-Morgenstern utility function $U(\cdot)$ defined over consumption of a vector $c_o = (c_{o1}, c_{o2}, ..., c_{oK})$ of all goods whose consumption and/or production is observed by the econometrician; a composite c_u of all goods whose consumption and/or production is unobserved by the econometrician;[5] and leisure ℓ. The function $U(\cdot)$ is quasiconcave but concave in each of its arguments, with the Inada condition $\frac{\partial U}{\partial x}\Big|_{x=0} = \infty$ with respect to each argument *x*.

[5] In order simplify the exposition, we refer to the vector of commodities whose consumption and/or production is unobserved by the econometrician as "the unobserved good" in what follows.

All K goods observed and the good unobserved by the econometrician can, in principle, be produced and/or consumed by the household.[6] The household has an endowment E^L of time and an endowment E^A of land. The production of each of the K observed commodities is denoted by

$$F_{oi}(L_{oi}, A_{oi}),\ i \in \{1,...,K\}, \tag{1}$$

where L_{oi} denotes the amount of labor used in producing observed commodity i and A_{oi} denotes the amount of cultivable land used in producing observed commodity i. The production of the unobserved good is denoted by

$$F_u(L_u, A_u), \tag{2}$$

where L_u and A_u denote the amount of labor and cultivable land, respectively, used in producing the unobserved commodity. Both F_{oi} and F_u are strictly increasing but weakly concave in each argument.

Agricultural labor is a function of household labor on the farm L^f and of hired labor L^h, but note that those are imperfect substitutes given that monitoring of hired workers may be imperfect, with the usual moral hazard consequence (Feder, 1985; Frisvold, 1994). The household can also sell a quantity L^m of labor on the market at parametric wage rate w, but the market for credit is missing.

[6] It is quite common in developing countries for rural household to grow a staple (e.g., maize) and nonstaples (e.g., coffee.) For a specific crop, it is also common for some households to be net buyers of it, for some households to be autarkic with respect to it, and for some households to be net sellers of it. Additionally, households often switch from one category – net buyer, autarkic, or net seller – to another from one period to the next.

The household's time constraint is such that $L^m + \ell + \sum_i L_{oi}^f + L_u^f = E^L$, where ℓ is the household's leisure time; L_{oi}^f is the amount of household labor devoted to production of observed commodity i and L_u^f is the amount of household labor devoted to production of the unobserved good. The household's land constraint is such that $A^m + A^f = E^A$, where A^m is the amount of household land leased out on the tenancy market at parametric rental rate r; and $A^f = \sum_i A_{oi}^f + A_u^f$ is the amount of household land devoted to the production of the observable and unobservable commodities, respectively. Likewise, A_{oi}^h and A_u^f are the amounts of leased in land devoted to the production of the observable and unobservable commodities, so that $A_{oi} = A_{oi}^f + A_{oi}^h$ and $A_u = A_u^f + A_u^h$ are the total amounts of land allocated to the production of the observable and unobservable commodities. Finally, let I denote the household's unearned income, i.e., income from transfers or remittances.

In what follows, we consider a two-period model. All product prices are unknown when labor allocation decisions are made, but post-harvest prices are revealed before consumption decisions are made.[7] The household's maximization problem is thus

$$\max_{\{A_{oi}^h, L_{oi}^h, A_{oi}^f, L_{oi}^f, L_u^h, L_u^f, A^m, \ell\}} E \max_{\{c_o, c_u\}} U(c_o, c_u, \ell) \tag{3}$$

subject to

$$p_o c_o + p_u c_u \leq Y^*, \tag{4}$$

[7] In what follows, we assume nothing about the shape of price uncertainty. Turnovsky (1978) noted how different theoretical results emerge depending on whether price uncertainty arises due to an additive or multiplicative error term. Our framework allows us to remain agnostic about the structure of price uncertainty, which in turn allows us to let the data speak for itself in section 5.

$$Y^* \equiv w[L^m - \sum_{oi} L_{oi}^h - L_u^h] + r[A^m - \sum_{oi} A_{oi}^h - A_u^h]$$

$$+ \sum_i p_{oi} F_{oi}(L_{oi}, A_{oi}) + p_u F_u(L_u, A_u) + I \,, \qquad (5)$$

$$L_{oi} \equiv h(L_{oi}^h) + L_{oi}^f \quad \forall i \,, \qquad (6)$$

$$L_u \equiv h(L_u^h) + L_u^f \,, \qquad (7)$$

$$L^m + \ell + \sum_i L_{oi}^f + L_u^f = E^L \,, \qquad (8)$$

$$A^f = \sum_i A_{oi}^f + A_u^f \qquad (9)$$

$$A^h = \sum_i A_{oi}^h + A_u^h \qquad (10)$$

$$A^m + A^f = E^A \qquad (11)$$

$$h(L_{oi}^h) \in [0, L_{oi}^h], \text{ and} \qquad (12)$$

$$h(L_u^h) \in [0, L_u^h]. \qquad (13)$$

Given that the household's utility function is strictly increasing, its budget constraint is binding. The household allocates labor and land conditional on its expectations regarding its *ex post* optimal choices of c_o, c_u, and ℓ.

By Epstein's (1975) duality result, we can use the household's (variable) indirect utility function $V(\cdot)$, which is homogeneous of degree zero in prices and income, i.e., the measurement unit chosen to measure prices and income do not matter. Thus, we can set the price of the unobserved commodity p_u as numéraire, so that $p_i = p_{oi} / p_u$ and $y = Y^* / p_u$.[8] Finally, assume that the household is (income) risk-averse, in the sense that

[8] Relative risk aversion, which we rely on when computing coefficients of price risk aversion below, is not affected by the choice of numéraire (Deschamps, 1973).

$\frac{\partial^2 V}{\partial y^2} = V_{yy} < 0$.[9] Consequently, as Deschamps (1973) noted, because of our setting p_u as numéraire, the household's Arrow-Pratt coefficient of absolute risk aversion is equal to $-p_u V_{yy}/V_y$, although its Arrow-Pratt coefficient of relative risk aversion remains unchanged and equal to $-yV_{yy}/V_y$.

Using the household's (variable) indirect utility function, we can rewrite the household's maximization problem as

$$\max_{\{A_{oi}^h, L_{oi}^h, A_{oi}^f, L_{oi}^f, L_u^h, L_u^f, A^m, \ell\}} EV(\ell, p, y) \tag{14}$$

subject to

$$Y = w[E^L - \ell - \sum_i L_{oi}^f - \sum_i L_{oi}^h - L_u^f - L_u^h] + r[E^A - \sum_i A_{oi}^f - \sum_i A_{oi}^h - A_u^f - A_u^h]$$
$$+ \sum_i p_i F_{oi}(L_{oi}, A_{oi}) + F_u(L_u, A_u) + I\ . \tag{15}$$

The first-order necessary conditions (FONCs) for this problem are then:

w.r.t. L_{oi}^h: $$E\left\{V_y\left(p_i \frac{\partial F_{oi}}{\partial L_{oi}^h} - w\right)\right\} \le 0 \ \ (=0 \text{ if } L_{oi}^h > 0), \tag{16}$$

w.r.t. A_{oi}^h: $$E\left\{V_y\left(p_i \frac{\partial F_{oi}}{\partial A_{oi}^h} - r\right)\right\} \le 0 \ \ (=0 \text{ if } A_{oi}^h > 0), \tag{17}$$

w.r.t. L_{oi}^f: $$E\left\{V_y\left(p_i \frac{\partial F_{oi}}{\partial L_{oi}^f} - w\right)\right\} \le 0 \ \ (=0 \text{ if } L_{oi}^f > 0), \tag{18}$$

w.r.t. A_{oi}^f: $$E\left\{V_y\left(p_i \frac{\partial F_{oi}}{\partial A_{oi}^f} - r\right)\right\} \le 0 \ \ (=0 \text{ if } A_{oi}^f > 0), \text{ and} \tag{19}$$

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[9] In a slight abuse of notation, we use subscripts not only to denote commodities but also the partial derivatives of the function $V(\cdot)$ in what follows.

w.r.t. ℓ : $$E\{V_\ell - V_y w\} \leq 0 \ (= 0 \text{ if } \ell > 0). \qquad (20)$$

Intuitively, equations (16) to (19) mean that the household is a profit maximizer, and equation (20) means that the household will set its (expected) marginal utility of leisure equal to its marginal cost. This set of FONCs is similar what is usually derived from the basic agricultural household model (see Singh et al., 1986 and Bardhan and Udry, 1999 for introductory treatments).

This framework was used by Barrett (1996) to explain the existence of the inverse farm size–productivity relationship as a result of staple food crop price risk. We now extend this framework to the case of multiple goods with stochastic prices. As such, the next subsection shows how to derive the household's matrix of own- and cross-price risk aversion coefficients.

2.2 Price Risk Aversion over Multiple Commodities

Let $V(\underline{p}, y)$ denote the household's indirect utility function, which has the usual properties. The vector $\underline{p} = (p_1,..., p_K)$ is the vector of commodity prices faced by the household over the observed commodities, and the scalar y denotes household income. Let p_i denote the price of commodity *i* and p_j denote the price of commodity *j*, without any loss of generality. We know from Barrett (1996) that

$$sign[Cov(V_y, p_i)] = sign(V_{yp_i}), \qquad (21)$$

and that, by Roy's identity (i.e., $M_i = -\frac{\partial V / \partial p_i}{\partial V / \partial y}$, where M_i is the marketable surplus of commodity *i*), we have that

$$V_y = -\frac{V_{p_i}}{M_i} = -\frac{V_{p_j}}{M_j}, \tag{22}$$

where M_j is the marketable surplus of commodity *j*. Additionally, note that

$$V_{yp_j} = \left(\frac{V_{p_i p_j}}{M_i} - \frac{V_{p_i}}{M_i^2} \frac{\partial M_i}{\partial p_j} \right) = \frac{1}{M_i} \left\{ V_{p_i p_j} - \frac{\partial M_i}{\partial p_j} V_y \right\}. \tag{23}$$

From Barrett (1996), we also have that

$$M_i = \frac{V_{p_i}}{V_y} \Leftrightarrow V_{p_i} = M_i V_y, \tag{24}$$

which implies that

$$V_{p_i p_j} = M_i V_{yp_j} + V_y \frac{\partial M_i}{\partial p_j}, \tag{25}$$

which, in turn, implies that

$$V_{p_i y} = M_i V_{yy} + V_y \frac{\partial M_i}{\partial y} = V_{yp_i}, \tag{26}$$

where the last equation is simply the result of applying Young's theorem. Replacing V_{yp_i} by equation (26) in equation (25) yields

$$V_{p_i p_j} = M_i \left\{ M_j V_{yy} + V_y \frac{\partial M_j}{\partial y} \right\} + V_y \frac{\partial M_i}{\partial p_j}. \tag{27}$$

Then, we have that

$$V_{p_i p_j} = M_i M_j V_{yy} + M_i V_y \frac{\partial M_j}{\partial y} + V_y \frac{\partial Mi}{\partial p_j}. \tag{28}$$

Multiplying the first term by $V_y y / V_y y$ yields (29)

$$V_{p_i p_j} = -\frac{M_i M_j R V_y}{y} + M_i V_y \frac{\partial M_j}{\partial y} + V_y \frac{\partial Mi}{\partial p_j}, \tag{30}$$

where R is the household's Arrow-Pratt coefficient of relative risk aversion. Multiplying the second term by $M_j y / M_j y$ and the third term by $M_i p_j / M_i p_j$ yields

$$V_{p_i p_j} = -\frac{M_i M_j R V_y}{y} + M_i V_y \eta_j \frac{M_j}{y} + V_y \varepsilon_{ij} \frac{M_i}{p_j}, \tag{31}$$

Where η_j is the income-elasticity of the marketable surplus of commodity j and ε_{ij} is the elasticity of commodity i with respect to the price of commodity j. Equation (31) is thus equivalent to

$$V_{p_i p_j} = M_i V_y \left[-\frac{M_j R}{y} + \eta_j \frac{M_j}{y} + \varepsilon_{ij} \frac{1}{p_j} \right]. \tag{32}$$

Multiplying the first two terms in the bracketed expression by p_j / p_j yields

$$V_{p_i p_j} = \frac{M_i V_y}{p_j} \left[-R\beta_j + \eta_j \beta_j + \varepsilon_{ij} \right], \tag{33}$$

in which β_j is the budget share of commodity j. When simplified, equation (33) is such that

$$V_{p_i p_j} = \frac{M_i V_y}{p_j} \left[\beta_j (\eta_j - R) + \varepsilon_{ij} \right]. \tag{34}$$

But then, applying Young's theorem again yields the following equation:

$$V_{p_i p_j} = \frac{M_i V_y}{p_j} \left[\beta_j (\eta_j - R) + \varepsilon_{ij} \right] = \frac{M_j V_y}{p_i} \left[\beta_i (\eta_i - R) + \varepsilon_{ji} \right] = V_{p_j p_i}. \tag{35}$$

In other words, the V_{pp} matrix, which is such that

$$V_{pp} = \begin{bmatrix} V_{p_1 p_1} & V_{p_1 p_2} & \cdots & V_{p_1 p_K} \\ V_{p_2 p_1} & V_{p_2 p_2} & \cdots & V_{p_2 p_K} \\ \vdots & \vdots & \ddots & \vdots \\ V_{p_K p_1} & V_{p_K p_2} & \cdots & V_{p_K p_K} \end{bmatrix} \tag{36}$$

is symmetric. More importantly, from the V_{pp} matrix, we can derive matrix A of price risk aversion coefficients:

$$\mathrm{A} = -\frac{1}{V_y} \cdot V_{pp} = -\frac{1}{V_y} \cdot \begin{bmatrix} V_{p_1 p_1} & V_{p_1 p_2} & \cdots & V_{p_1 p_K} \\ V_{p_2 p_1} & V_{p_2 p_2} & \cdots & V_{p_2 p_K} \\ \vdots & \vdots & \ddots & \vdots \\ V_{p_K p_1} & V_{p_K p_2} & \cdots & V_{p_K p_K} \end{bmatrix}$$

$$= \begin{bmatrix} A_{11} & A_{12} & \cdots & A_{1K} \\ A_{21} & A_{22} & \cdots & A_{2K} \\ \vdots & \vdots & \ddots & \vdots \\ A_{K1} & A_{K2} & \cdots & A_{KK} \end{bmatrix} \tag{37}$$

Matrix A has a straightforward interpretation. The sign of the diagonal elements is analogous to Pratt's (1964) coefficient of absolute income risk aversion (Appendix A derives true price analogs of the Arrow-Pratt measure, i.e., measures of absolute and relative price risk aversion, both theoretically and empirically), but for prices; $A_{ii} > 0$ implies that welfare is decreasing in the volatility of the price of *i*. This is the classic concern of the literature on commodity price stabilization (Deschamps, 1973; Hanoch, 1974, Turnovsky, 1978; Turnovsky et al., 1980; and Newbery and Stiglitz, 1981).

The off-diagonals, meanwhile, reflect how variation in one good's price affects the household's marginal utility with respect to variation in the other good's price. Consequently, $A_{ij} > 0$ implies that greater volatility in price *j* reduces welfare associated with the net consumption of good *i*. The price risk aversion coefficient matrix thus speaks directly to the welfare effects of and household preferences with respect to price risk. Intuitively, these coefficients provide information on how the marginal utility of a price change in good *i* is impacted by a price change in good *j*. In other words, whereas the diagonal terms can be interpreted as the effect on household welfare of the variance in the price of a single good, the off-diagonal terms can be interpreted as the effect on household welfare of the *covariance* between the prices of two goods.

The theory clearly implies a strong and testable symmetry restriction on the estimated price risk aversion coefficients. With adequate data, one can test the following null hypothesis:

$$H_0 : A_{ij} = A_{ji} \text{ for all } i \neq j , \tag{38}$$

which consists of $\frac{K(K-1)}{2}$ testable restrictions. A test of the symmetry of A is a test of the rationality of the household and, consequently, a test of unitary preferences. Among other things, we pursue this test in the empirical application below.

2.3 Relationship between *A* and the Slutsky Matrix

The derivations above and their result culminating in the price risk aversion coefficient matrix A raise the natural question: What is the relationship between matrix A and the Slutsky matrix? That is, what can knowledge of a household's marketable surplus equations tell us about its attitude with respect to price risk?

Let $M_i(p, y)$ be the household's marketable surplus of commodity *i* as a function of the prices the household faces and its income. We know from first principles that the Slutsky matrix *S* is such that (Mas-Colell et al., 1995)

$$S_{ij}(p, y) = \frac{\partial M_i}{\partial p_j} + \frac{\partial M_i}{\partial y} x_j = B_{ij} + C_{ij} . \tag{39}$$

Where $B_{ij} \equiv \frac{\partial M_i}{\partial p_j}$ and $C_{ij} \equiv \frac{\partial M_i}{\partial y} x_j$. Based on the derivations of the previous section, we can show that

$$A_{ij} = M_i \left[\frac{1}{M_j} C_{jj} - \frac{R}{y} + B_{ij} \right]. \tag{40}$$

That is, a household's marginal utility with respect to a change in the price of good i varies as a result of a change in the price of good j, and that variation is a function of the commodity's own-income effect as well as the cross-price effect between goods i and j. In this sense, since the cross-price risk aversion between goods i and j is linked both to S_{jj} and to S_{ij}, there does not exist a one-to-one correspondence between the elements of matrices *A* and *S*.[10] The sign of any cross-price risk aversion coefficient does not depend on whether two goods are complements or substitutes, and one cannot recover the Slutsky matrix from the matrix of price risk aversion coefficients.

Even though there does not exist a one-to-one correspondence between the Slutsky matrix and the matrix of price risk aversion coefficients, the following results further characterize the relationship between the two matrices.

Proposition 1 Symmetry of the matrix of price risk aversion coefficients is equivalent to symmetry of the Slutsky matrix.

Proof See Appendix B.

[10] Intuitively, this is because the Slutsky matrix is derived from an ordinal function, and the matrix of price risk aversion is derived from a cardinal function (Deschamps, 1973).

Corollary Symmetry of the Slutsky matrix imposes more structure on the data than symmetry of the price risk aversion matrix.

Proof See Appendix B.

Symmetry of the Slutsky matrix and symmetry of the matrix of price risk aversion coefficients have the same empirical content: they both embody rationality of the household. As such, failing to reject the symmetry of either matrix corresponds to a failure to reject that the household has unitary preferences.

2.4 Certainty Equivalent for Price Stabilization

Suppose that a policymaker wanted to eliminate the income equivalent of all price uncertainty. In order to do so, one would first need to compute the total certainty equivalent (CE), i.e., the certainty equivalent for the entire price risk aversion matrix. Then,

$$CE = V\left(E\left(p\right), y\right) - E\left(V\left(p, y\right)\right), \tag{41}$$

so that

$$CE = V\left(E\left(p\right), y\right) - E\left(V\left(p, y\right)\right) = E\left[V\left(E\left(p\right), y\right) - V\left(p, y\right)\right]. \tag{42}$$

A Taylor series approximation around $V\left(E\left(p\right), y\right)$ yields

$$CE \approx E\left[-V_p\left(E\left(p\right), y\right)\left(p - E\left(p\right)\right) - \frac{1}{2}\left(p - E\left(p\right)\right)' V_{pp}\left(E\left(p\right), y\right)\left(p - E\left(p\right)\right)\right], \tag{43}$$

in other words,

$$CE \approx -\frac{1}{2}E\left[\left(p-E(p)\right)'V_{pp}\left(E(p),y\right)\left(p-E(p)\right)\right]$$

$$= -\frac{1}{2}E\left[\begin{bmatrix} p_1-\mu_1 & p_2-\mu_2 & \cdots & p_k-\mu_k \end{bmatrix}\begin{bmatrix} V_{p_1p_1} & V_{p_1p_2} & \cdots & V_{p_1p_K} \\ V_{p_2p_1} & V_{p_2p_2} & \cdots & V_{p_2p_K} \\ \vdots & \vdots & \ddots & \vdots \\ V_{p_Kp_1} & V_{p_Kp_2} & \cdots & V_{p_Kp_K} \end{bmatrix}\begin{bmatrix} p_1-\mu_1 \\ p_2-\mu_2 \\ \vdots \\ p_k-\mu_k \end{bmatrix}\right]$$

$$= -\frac{1}{2}E\left[(p_1-\mu_1)\sum_{i=1}^{k}\left[(p_i-\mu_i)V_{p_ip_1}\right] \quad (p_2-\mu_2)\sum_{i=1}^{k}(p_i-\mu_i)V_{p_ip_2} \quad \cdots \quad (p_k-\mu_k)\sum_{i=1}^{k}(p_i-\mu_i)V_{p_ip_k}\right]$$

$$= -\frac{1}{2}\left[\sum_{i=1}^{k}\sigma_{p_1p_i}V_{p_ip_1} \quad \sum_{i=1}^{k}\sigma_{p_2p_i}V_{p_ip_2} \quad \cdots \quad \sum_{i=1}^{k}\sigma_{p_kp_i}V_{p_ip_k}\right]$$

If instead one wishes to stabilize only one price, the above derivations become such that

$$CE = V\left(E(p_i), p_{\sim i}, y\right) - E\left(V(p_i, p_{\sim i}, y)\right), \tag{44}$$

so that

$$CE = E\left(V\left(E(p_i), p_{\sim i}, y\right)\right) - E\left(V(p_i, p_{\sim i}, y)\right) = E\left[V\left(E(p_i), p_{\sim i}, y\right) - V(p_i, p_{\sim i}, y)\right]. \tag{45}$$

Again, a Taylor series expansion yields

$$CE \approx E\begin{bmatrix} V\left(E(p),y\right)+V_{p_{\sim i}}\left(E(p),y\right)\left(p_{\sim i}-E(p_{\sim i})\right)+\frac{1}{2}\left(p_{\sim i}-E(p_{\sim i})\right)'V_{p_{\sim i}p_{\sim i}}\left(E(p),y\right)\left(p_{\sim i}-E(p_{\sim i})\right) \\ -V\left(E(p),y\right)-V_{p_{\sim i}}\left(E(p),y\right)\left(p_{\sim i}-E(p_{\sim i})\right)-V_{p_i}\left(E(p),y\right)\left(p_i-E(p_i)\right) \\ -\left(p_{\sim i}-E(p_{\sim i})\right)'\frac{1}{2}V_{p_{\sim i}p_{\sim i}}\left(E(p),y\right)\left(p_{\sim i}-E(p_{\sim i})\right) \\ -\left(p_i-E(p_i)\right)'\frac{1}{2}V_{p_ip_i}\left(E(p),y\right)\left(p_i-E(p_i)\right) \\ -\left(p_i-E(p_i)\right)'\frac{1}{2}V_{p_ip_{\sim i}}\left(E(p),y\right)\left(p_{\sim i}-E(p_{\sim i})\right) \\ -\left(p_{\sim i}-E(p_{\sim i})\right)'\frac{1}{2}V_{p_{\sim i}p_i}\left(E(p),y\right)\left(p_i-E(p_i)\right) \end{bmatrix}$$

$$CE \approx -\frac{1}{2}\sigma_i^2 V_{p_i p_i} + E\begin{bmatrix} -\left(p_i - E\left(p_i\right)\right)'\frac{1}{2}V_{p_i p_{\sim i}}\left(E\left(p\right), y\right)\left(p_{\sim i} - E\left(p_{\sim i}\right)\right) \\ -\left(p_{\sim i} - E\left(p_{\sim i}\right)\right)'\frac{1}{2}V_{p_{\sim i} p_i}\left(E\left(p\right), y\right)\left(p_i - E\left(p_i\right)\right) \end{bmatrix}$$

$$CE \approx -\frac{1}{2}\sigma_i^2 V_{p_i p_i} - \sum_{j \neq i}\sigma_{ij} V_{p_i p_j} \ . \tag{46}$$

This last equation thus provides the transfer payment a policymaker would need to make to the household in order to compensate them for the uncertainty over p_i, and the previous derivations provide the transfer payment a policymaker would need to make to the household in order to compensate them for the uncertainty over *all* prices.

3. Data and Descriptive Statistics

In the remainder of the paper we demonstrate the estimation of the price risk aversion coefficient matrix and test the symmetry restriction implied by the theory. We use data from the Ethiopian Rural Household Survey (ERHS).[11] The publicly available ERHS data include results from five rounds: 1989, 1994a, 1994b, 1995, and 1997, although the latter round is still being processed. The survey instrument and sampling strategy changed significantly between 1989 and 1994a, with the addition of nine peasant associations (PAs) to the original six surveyed in 1989.[12]

[11] These data are made available by the Department of Economics at Addis Ababa University (AAU), the Centre for the Study of African Economies (CSAE) at Oxford University, and the International Food Policy Research Institute (IFPRI). Funding for data collection was provided by the Economic and Social Research Council (ESRC), the Swedish International Development Agency (SIDA) and the US Agency for International Development (USAID). The preparation of the public release version of the ERHS data was supported, in part, by the World Bank. Evidently, AAU, CSAE, IFPRI, ESRC, SIDA, USAID, and the World Bank are not responsible for any errors in these data or for their use or interpretation.

[12] Ethiopia is subdivided into eleven zones subdivided into *woredas*, which are roughly equivalent to US counties. Within each *woredas* are PAs. Random sampling occurred at the village level within these PAs. Other differences between 1989 and subsequent surveys included (i) a longer list of consumption items starting in the 1994a round; (ii) a lack of price data in the 1989 round which were collected in future rounds; and (iii) a war that ended between the 1989 and 1994a rounds (Dercon 1998).

We chose this dataset because it records household consumption and production decisions over multiple years, and because there is low attrition and a standardized survey instrument across the rounds we retain for analysis. The sample includes a total of 1471 households with an attrition rate of around 2 percent of households across the four rounds selected for analysis (Dercon and Krishnan, 1998).

Given that many of the households in our data were autarkic with respect to several commodities due to the presence of transactions costs that prevents them from participating in the market either as net buyers or as net sellers (de Janvry et al., 1991), for every time period in which a household is neither a net buyer or a net seller of a given commodity, this household has a marketable surplus of zero for that particular commodity. In what follows, we focus on maize, coffee, barley, cooking oil, wheat, beans, and sorghum, i.e., the top seven commodities when considering the proportion of observations with a nonzero marketable surplus. We would have liked to also include teff, which is used to make *injera* – a flatbread used to pick food out of a common pot – and is perhaps the most well-known Ethiopian staple, but since its price was dropped from every marketable surplus equation due to collinearity, we had to omit it from our empirical derivations.

Table 1 introduces descriptive statistics for these seven top commodities. A positive (negative) mean marketable surplus indicates that the average household is a net seller (buyer) of a commodity. On the one hand, the average household is a net buyer of cooking oil. On the other hand, the average household is a net seller of maize, coffee, barley, wheat, beans, and sorghum.[13] This is consistent with maize, barley, wheat, and sorghum being staple crops produced by the households in our data set and cooking oil being a non-staple. Also note that no household in our data is a net seller of cooking oil.

Table 2 lists the mean price for each of the seven commodities under study in Ethiopian birr,[14] along with the mean budget share for each commodity. Note that maize, sorghum, and barley represent the highest budget shares, at 160, 108, and 80 percent of the average household's budget, and that wheat and cooking oil – which both have negative budget shares because the average household is a net buyer of these commodities – represent the lowest budget shares, at -0.2 percent of the average household's budget.

The income measure used in this paper is the sum of proceeds from off-farm labor, off-farm business activity and land rentals, gifts, crop sales, and other sources of incomes.[15] That said, average income from the aforementioned sources is different from zero in about 65 percent of cases, which explains why the average annual income of $57 may seem low.

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[13] It may be surprising that the number of net buyers of coffee far exceeds the number of net sellers of coffee in these data. As it turns out, the average buyer buys about 17 kg of coffee per year, and the average seller sells 87 kg of coffee per year (both statistics are significant at the 1 percent significance level). These numbers are consistent with the net buyers buying coffee purely for household consumption, and the net sellers selling much larger quantities at market. Although the 17 kg figure for net buyers may seem high, note that a can of Illy coffee contains 0.25 kg of (roasted) whole coffee beans. This means that the average net buyer household would have to consume about 64 such cans per year (i.e., a little more than one per week), which is far from unlikely considering the average household size as well as the frequency at which coffee is consumed in Ethiopia.

[14] As of writing, US$1 ≈ Birr 9.43.

[15] We omit own-crop revenue from income in each marketable surplus equation estimated below so as to avoid biasing our estimates below due to the endogeneity of income to marketable surplus.

4. Empirical Framework

We define a household's marketable surplus of a given commodity as the quantity harvested of that commodity net of the quantity purchased and the household's consumption of its own harvest. Structurally speaking then, a household's marketable surplus M_i is such that

$$M_i = s_i(z, p) - x_i(p, y)\,, \tag{47}$$

where $s_i(z, p)$ is the household's supply of commodity *i*, which depends on a vector *z* of input prices and a vector *p* of commodity prices; and $x_i(p, w)$ is the household's Walrasian demand of commodity *i*, which depends on a vector *p* of commodity prices and household income *w*. This means that, at the very least, our marketable surplus equations must include (i) commodity prices; (ii) input prices; and (iii) household income. Our data includes commodity prices and allows computing household income, but includes only wage as an input price. And even when wage is included, it is recorded at the village level. Thus, to control for input prices, which are common for every household in a given *woreda* at any given point in time, we include *woreda*-round fixed effects in each marketable surplus equation. Additionally, household fixed effects allow us to partially control for input shadow prices. Although the ideal data set would include precise measures for all commodity and input prices, this is the best we can do with the data at hand.[16]

[16] It could be argued, however, that input prices do not need to be included in our marketable surplus equations because the production of each commodity is predetermined, in which case the *woreda*-round fixed effects would only serve as additional controls.

We estimate the following marketable surplus functions for the seven commodities *i* discussed in the previous section:

$$M_i = \beta_{i0} + \beta_{i1} \ln W_i + \beta_{i2} \ln P_i + \beta_{i3} \ln P_{-i} + \lambda + \tau + \varepsilon_i, \qquad (48)$$

where *i* denotes the commodity rather than the observation; W_i denotes household income net of income from commodity *i*; P_i is a household-specific measure of the price of commodity *i*; P_{-i} is a household-specific measure of the prices of all (observed) commodities other than *i* (including *j*); λ is a region-*woreda*-peasant association-household fixed effect; τ is a woreda-round fixed effect which allows controlling for the price of the unobservable composite good as well as for the input prices; and ε_i is an error term with mean zero.

Since the household is a price-taker for all commodities, then all prices are fully exogenous to the dependent variable in equation 48, and household income is exogenous by virtue of net of the household's income from commodity *i*.

Given the panel nature of our data, we estimate equation over a total of 1,471 units of observation spread out over four rounds and three seasons. No household was observed over all four rounds and three seasons, as the number of observations per household ranged from one to eight with an average of 5.9 observations per household.[17] We also include all commodity prices available in our data (i.e., barley, wheat, maize, sorghum,

[17] By controlling for household unobservables, the use of fixed effects controls for the possible selection problem posed by households for which we only have one observation through time, which are dropped from the fixed effects regressions we estimate.

beans, coffee, potatoes, onions, cabbage, cooking oil, soap, and three commodities found in the data but for which we could not find corresponding crop codes).

Computation of own- and cross-price elasticities, of the income-elasticity, and of the budget share of marketable surplus follows from equation 48. For example, given the functional form of equation 48, to derive the estimated cross-price risk aversion coefficient $\hat{A}_{ij}$, one would first need to compute $\hat{\beta}_j = \frac{M_j p_j}{W}$, $\hat{\eta}_j = \frac{\hat{\beta}_{j1}}{M_j}$, and $\hat{\varepsilon}_{ij} = \frac{\hat{\beta}_{i3}}{M_i}$, where $\hat{\beta}_{i3}$ is the estimated coefficient for the price of commodity *j* in marketable surplus equation for commodity *i*. Then, one could combine these estimates to obtain the point estimate

$$\hat{A}_{ij} = \frac{M_i}{p_j}[\hat{\beta}_j(\hat{\eta}_j - R) + \hat{\varepsilon}_{ij}], \tag{48}$$

where $\hat{\beta}_j = \frac{M_j p_j}{W}$, $\hat{\eta}_j = \frac{\hat{\beta}_{j3}}{M_j}$, and $\hat{\varepsilon}_{ij} = \frac{\hat{\beta}_{i5}}{M_i}$. Given that marketable surplus is often zero, we use the mean of M_j and M_i so as to compute elasticities at means. Although it might be preferable to use mean elasticities, it is simply not possible to do so in these data.[18] The standard errors can then be computed using either the delta method or through

[18] Likewise, given that we use the household's income from non-agricultural sources as a proxy for total income *W*, so as to avoid endogeneity problems, many households have an income of zero. Table 2 shows that income was different from zero in 2969 observations out of 5667. In this case, we compute the estimated budget share by dividing by *W* + 0.001 (MaCurdy and Pencavel, 1986). We also add 0.001 to each observation for the variables for which logarithms are taken so as to not drop observations in a nonrandom fashion and introduce selection bias.

bootstrapping, given that $\hat{A}_{ij}$ is a nonlinear combination of estimates. The coefficient of relative risk aversion *R* can either be directly estimated – if the data allows it – or assumed to be equal to a certain value. Given that our data do not allow direct estimation of *R*, we estimate the A_{ij} coefficients for $R=1$, $R=2$, and $R=3$, which covers the range of credible values found in the literature (Friend and Blume, 1975; Hansen and Singleton, 1982; Chavas and Holt, 1993; and Saha et al., 1994). This provides an additional robustness check.

As shown in table 1, many households have a marketable surplus of zero for several commodities, so we test several version of the matrix of price risk aversion coefficients. We first estimate the A matrix for the top three commodities consumed and produced by the households in our data (i.e., maize, coffee, and barley), and then estimate it for the top four, and so on for the five, six, and seven top commodities (i.e., maize, coffee, barley, cooking oil, wheat, beans, and sorghum). We stop at estimating the matrix of price risk aversion coefficients for the top seven commodities because the percentage of nonzero marketable surplus observations does not exceed five percent for the remaining commodities. Finally, note that we divide each coefficient of price risk aversion by its standard error so as to standardize them and make them comparable with one another, given that it would be otherwise impossible to compare different crops.[19]

[19] For more on the impossibility to compare different crops without standardizing price risk aversion coefficients, see Bellemare (2005).

5. Estimation Results and Hypothesis Tests

This section first presents estimation results for the marketable surplus equations. Given that these results are ancillary in the sense that they represent an intermediate step in computing own- and cross-price risk aversion coefficients, we only discuss them briefly and devote the bulk of the analysis to the estimated matrices of price risk aversion as well as the tests of the hypothesis that these matrices are symmetric.

Table 3 presents estimation results for the seven marketable surplus equations considered in this paper. The first thing to note is that although own- and cross-price elasticities are used along with income elasticities of marketable surpluses in computing coefficients of own- and cross-price risk aversion, what matters for the estimated marketable surplus equations are the estimated coefficients themselves. In other words, one should normally expect the estimated coefficient on p_i to be positive in the marketable surplus equation for M_i, even though the elasticity will flip signs depending on whether the mean of M_i is positive or negative since we are relying on elasticities *at means* rather than mean elasticities when computing coefficients of price risk aversion. Consequently, one should focus on β_{i2} in equation 47 in order to determine whether p_i has the expected effect on M_i.

Intuitively, one would expect β_{i2} to be positive, i.e., as the price of commodity *i* increases, the household buys less and less or sells more and more of the same commodity. But given that we are studying agricultural households, and not pure producers or pure consumers, however, it is entirely possible that β_{i2} be negative. Recall that within an agricultural household, there exists a profit effect on top of the usual income and substitution effects (Singh et al., 1986). It is thus entirely possible that, as p_i increases, the household increases its production of commodity *i* (which would increase M_i), but that it also decides to increase its consumption of it via an increased profit due to the increase in M_i. If the latter effect dominates, then β_{i2} will be negative. Obviously, one would expect β_{i2} to be more likely to be negative for goods that are actually produced by the households in the data, i.e., for goods for which there is indeed a profit effect. Such counterintuitive results have been reported in the market participation studies of Goetz (1992), Bellemare and Barrett (2006), and Stephens and Barrett (2006).

Turning to the estimation results, own price has a positive and significant effect on the marketable surplus of almost all commodities in table 3. The price of coffee, although it has a positive effect on the marketable surplus of coffee in every specification, is not statistically significant.[20]

As discussed above, we use the results of table 3 to compute standardized coefficients of own- and cross-price risk aversion, and we use these coefficients to construct sub-matrices A_3 to A_7 of price risk aversion.[21] Looking at table 4, the first thing to note is that the households in our data are price risk-averse over most commodities except maize, whose own-price risk aversion coefficients is not significant at any of the conventional

[20] Note that the low overall R^2 measures comes from the fact that the R^2 measure reported by Stata when using the `xtreg` command with fixed effects is not directly comparable to its analog in an ordinary least squares regression, and should not be interpreted as such (StataCorp, 2005).
[21] We use the term "sub-matrix" given that the number of commodities produced and consumed by the household in theory goes to infinity. This is similar to Turnovsky et al. (1980), who only consider a subset of commodities in their theoretical analysis.

levels. In addition, the average household is most price risk-averse over cooking oil and sorghum, and least price risk-averse over maize and beans. Similarly, the average household is risk-averse over co-fluctuations in the prices of coffee and all the other commodities and over co-fluctuations in the prices of cooking oil all the other commodities except wheat, as witnessed by the off-diagonal elements in the matrix in table 4.

Turning to our main testable hypothesis, i.e., the symmetry of the matrix of price risk aversion coefficients, for sub-matrix A_3 (i.e., the sub-matrix of price risk aversion coefficients for maize, coffee, and barley) the null hypothesis of symmetry cannot be rejected, with a *p*-value of 0.65, as shown in table 5. Thus, even though the test of symmetry has low power because most of the probability mass rests on non-rejection of the null hypothesis, the *p*-value gives us some confidence in the outcome of the test. Likewise, symmetry cannot be rejected for sub-matrices A_4 to A_7 in table 5, with *p*-values of 0.86, 0.69, 0.86, and 0.92, respectively. Thus, as we expand sub-matrix A to cover more and more commodities, our core finding remains: symmetry of the matrix of price risk aversion coefficients cannot be rejected, and even though the test does have low power, the increasing number of commodity and the high *p*-values offer strong empirical support for the theoretical framework. Note that these results are not due to the fact that the coefficients included in matrices A_3 to A_7 are not statistically significant: the null hypothesis that all coefficients are equal to zero is rejected with a *p*-value of 0.00, and well over half the coefficients in matrix A_7 are significantly different from zero.

Furthermore, table 6 shows that the symmetry results are consistent whether one assumes that relative risk aversion is such that $R=1$, $R=2$, or $R=3$, which provides an additional robustness check.[22] Our results thus seem to indicate that (i) the households in our data are significantly risk-averse over the price of specific commodities; (ii) the households in our data are significantly risk-averse over several of the co-movements in the prices of specific pairs of commodities; and (iii) the households in our data behave as theory predicts, in the sense that their risk preferences over co-movements in the prices of specific pairs of commodities are symmetric.

A *caveat* is in order, however. Recall that the proportion of non-zero observations for marketable surplus was at most 25 percent for the commodities considered, and that only six commodities exhibit 10 percent or more of non-zero observations. This high proportion of zero-marketable surplus observations may introduce a significant amount of noise in the data. Evidently, this would entail large standard errors around our estimates for the coefficients of own- and cross-price risk aversion, which would make non-rejection of the symmetry of the A matrix more likely. We take comfort, however, in the high *p*-values obtained when testing for symmetry as well as in the fact that given the high transactions costs faced by households in developing countries (Key et al., 2000; Renkow et al., 2004; Bellemare and Barrett, 2006), one would be hard-pressed to find a data set without such a large number of zero-valued observations for the marketable surplus of many commodities.

[22] Note that because these coefficients are treated as constants in the price risk aversion coefficients, they do not introduce any noise in these coefficients, and so changing risk aversion should have an effect on our point estimates, but not on our standard errors.

Having discussed above the link between the matrix of price risk aversion coefficients and the Slutsky substitution matrix, it is only natural to wish to compare the two. Table 7 shows the estimated Slutsky matrix for the households in our data based on the marketable surplus equations in table 3. All of the own-price Slutsky substitution terms are significant, and all but one (i.e., cooking oil) own-price substitution terms are negative, as microeconomic theory suggests for pure consumers. Does this mean that cooking oil is a Giffen good? Most likely not, given that we are not estimating demand functions for pure consumers, but marketable surplus equations for agents that are both consumers *and* producers.

The evidence gets more interesting in table 8, where we present tests of symmetry of the various Slutsky sub-matrices. In every case, symmetry of the Slutsky matrix is overwhelmingly rejected, and even though several off-diagonal equality constraints were dropped due to collinearity, each individual restriction was found not to hold in the data. This means that rationality of the household is rejected when looking at the Slutsky matrix, but not when looking at the price risk aversion matrix. Deaton and Muellbauer (1980) had also rejected the symmetry of the Slutsky matrix.

Our result could be due to several reasons. First off, households are not pure consumers, and their production behavior could contaminate estimates of traditional consumer-theoretic parameters. Second, it is entirely possible that the theory of the consumer needs to be adapted to the case where there is price risk, and that taking price risk into consideration could lead to a non-symmetric Slutsky matrix. Our theoretical framework, however, shows that it is easier to reject symmetry of the Slutsky matrix than it is to reject symmetry of the matrix of price risk aversion, so our rejection of the former and our non-rejection of the latter are consistent with our theoretical framework.

6. Conclusion

Based on the observation that the indirect utility function of a unitary household allows computing coefficients of risk aversion over more than just income or wealth, this paper has modestly extended microeconomic theory so as to allow studying price risk aversion over multiple commodities. Specifically, we have derived a pseudo-Slutsky matrix that measures the curvature of the indirect utility function in the hyperspace defined by the vector of all commodity prices faced by the household. Then, based on the method developed by Newbery and Stiglitz (1981) and extended in turn by Finkelshtain and Chalfant (1991) and by Barrett (1996), we have estimated the aforementioned matrix of price risk aversion coefficients using well-known survey data on a panel of rural Ethiopian households.

We then tested the paper's main testable implication (i.e., symmetry of the matrix of price risk aversion coefficients) on five nested subsets of commodities. In no case could we reject symmetry of the matrix of price risk aversion coefficients, and the *p*-values of the tests were always high enough so as to give us faith in the results and obviate concerns regarding the low power of our test.

Above and beyond the technical results presented in this paper, our results also have important policy implications. First off, the average household in these data is not only adversely affected by the price fluctuation of most commodities in the data, but also by co-movements in the price of the commodities we chose to include in our analysis. Based on this finding, we have shown that it is possible to compute certainty equivalents for each commodity or for all commodities, depending on whether policymakers want to compensate households for the price fluctuations of a few select staple crops or of a number of commodities taken as a whole.[23] As such, not only does our method allow computing average transfer payments, it also allows computing such transfer payments conditional on certain household covariates, such as income and the prices faced by each household.

That said, a *caveat* is in order. Many of the households in our data were autarkic with respect to all the commodities considered in our analysis, and even for the households whose marketable surplus was non-zero for one commodity, their marketable surplus of other commodities was often equal to zero. This means that we had to rely on elasticities at means rather than on mean elasticities when computing coefficients of price risk aversion, and so the variation exploited in computing the typical coefficient of price risk aversion came from marketable surplus, price, and budget share, rather than from the same variables augmented with price elasticity, and income elasticity. As such, the standard errors around our estimated coefficients of price risk aversion are smaller than they would be under ideal circumstances, i.e., they are smaller than they would be if we

[23] Giving such a transfer payment to American taxpayers to compensate them for fluctuations in the price of gasoline in 2008 resulted in the Bush administration economic stimulus checks.

had data for which *all* households have a nonzero marketable surplus of *all* commodities. This could in turn lead to our not rejecting symmetry of the matrix of price risk aversion coefficients. It would thus be of interest to both theoretical and applied microeconomists to estimate matrices of price risk aversion coefficients using data from industrialized countries, testing for symmetry for both single- and multiple-person households, much like Browning and Chiappori (1998) did in their seminal study of efficient intrahousehold decisions.

Finally, we have estimated the average household's Slutsky matrix, which relied on the same estimates we used to compute the pseudo-Slutsky matrix of price risk aversion. Interestingly enough, whereas symmetry of the former matrix could not be rejected, symmetry of the latter matrix was overwhelmingly rejected. We speculate that this could either be due to the fact that we are dealing with agents that both produce and consume or to the fact that symmetry of the Slutsky matrix may not necessarily obtain in the presence of price risk. For now, such an investigation is left for future research.

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Table 1: Descriptive Statistics for the Dependent Variables

Crop	Mean	(Std. Dev.)	Observations	Nonzero Observations	Net Buyers	Net Sellers
Maize (Kg)	63.83	(895.00)	8722	1763	895	868
Coffee (Kg)	1.08	(44.70)	8722	1534	1194	340
Barley (Kg)	87.33	(553.69)	8722	1504	518	986
Cooking Oil (Kg)	-4.41	(21.77)	8722	1333	1333	0
Wheat (Kg)	33.57	(325.15)	8722	993	664	329
Beans (Kg)	4.37	(178.43)	8722	733	576	157
Sorghum (Kg)	59.15	(503.34)	8722	625	236	389

Note: Although teff figures prominently in the average household's buying and selling behavior and 1089 households had a nonzero marketable surplus of teff, we omit this commodity from our analysis given that its price was always dropped from each marketable surplus equation due to collinearity.

Table 2: Descriptive Statistics for the Independent Variables

Crop	Mean	(Std. Dev.)	Observations
Prices			
Maize (Birr/Kg)	1.22	(0.34)	8722
Coffee (Birr/Kg)	12.21	(4.95)	8722
Barley (Birr/Kg)	1.43	(0.37)	8722
Cooking Oil (Birr/Kg)	1.65	(1.00)	8722
Wheat (Birr/Kg)	1.66	(0.31)	8722
Beans (Birr/Kg)	1.80	(0.42)	8722
Sorghum (Birr/Kg)	1.46	(0.40)	8722
Budget Shares			
Budget Share of Maize	1.600	(44.62)	5632
Budget Share of Coffee	0.148	(2.69)	5632
Budget Share of Barley	0.797	(13.91)	5632
Budget Share of Cooking Oil	-0.002	(0.07)	5632
Budget Share of Wheat	-0.019	(11.58)	5632
Budget Share of Beans	0.211	(9.02)	5632
Budget Share of Sorghum	1.075	(19.39)	5632
Income (Birr)	542.438	(7344.269)	8722

Note: Income (i.e., the sum of off-farm income and all crop revenues) was different from zero for 5632 observations only.

Table 3: Marketable Surplus Equations for Seven Commodities Over Four Rounds

	Maize	Coffee	Barley	Cooking Oil	Wheat	Beans	Sorghum
Barley Price	76.679***	7.560***	**82.156*** **	8.596***	-28.155***	14.565***	333.429***
	(18.479)	(1.021)	**(8.748)**	(1.390)	(9.209)	(1.409)	(9.877)
Wheat Price	1081.445***	1.857	-138.941***	15.722***	**54.445*** **	-6.517***	-663.934***
	(26.591)	(1.530)	(12.776)	(1.973)	**(13.243)**	(2.005)	(14.603)
Maize Price	**172.357*** **	-8.560***	-68.527***	-10.347***	37.494***	-22.373***	-143.943***
	(4.731)	(0.041)	(0.641)	(0.111)	(0.792)	(0.108)	(0.133)
Sorghum Price	-405.314***	0.218	44.929***	-2.980***	14.894***	-10.506***	**366.884*** **
	(10.602)	(0.529)	(6.521)	(0.753)	(5.013)	(0.760)	**(6.554)**
Beans Price	421.756***	4.939***	52.811***	-1.755***	-68.444***	**9.136*** **	568.921***
	(6.172)	(0.050)	(1.090)	(0.027)	(0.030)	**(0.036)**	(1.195)
Coffee Price	97.197***	**0.951**	103.669***	21.483***	12.808	10.305***	178.029***
	(14.139)	**(0.879)**	(9.918)	(1.141)	(7.904)	(1.185)	(6.800)
Potatoes Price	328.824***	2.422***	-2.274***	0.378***	17.089***	-6.677***	123.830***
	(1.622)	(0.017)	(0.481)	(0.002)	(0.000)	(0.003)	(0.944)
Onions Price	-499.872***	1.684***	12.099***	3.320***	8.611***	11.279***	27.784***
	(4.436)	(0.187)	(1.758)	(0.264)	(1.895)	(0.270)	(0.937)
Cabbage Price	-24.654***	2.411***	2.487	2.066***	-25.936***	6.593***	87.673***
	(4.586)	(0.330)	(3.200)	(0.424)	(2.871)	(0.436)	(1.911)
Unknown Price 1	183.293***	0.104	65.225***	18.378***	17.397	-2.206	98.553***
	(17.405)	(1.121)	(12.070)	(1.458)	(10.030)	(1.507)	(8.435)
Unknown Price 2	381.839***	-4.171***	-104.729***	-23.207***	3.123	-16.159***	80.474***
	(16.501)	(0.941)	(9.839)	(1.241)	(8.535)	(1.282)	(5.918)
Unknown Price 3	-102.821***	-4.838***	-146.075***	-3.517***	89.826***	-46.658***	-192.020***
	(8.701)	(0.466)	(4.596)	(0.629)	(4.262)	(0.645)	(5.150)
Cooking Oil Price	-287.472***	-8.637***	-49.383***	**8.300*** **	22.433***	-2.429***	64.487***
	(10.633)	(0.603)	(6.767)	**(0.816)**	(5.320)	(0.820)	(3.816)
Soap Price	-349.384***	-5.429***	11.862*	-8.420***	-24.892***	11.094***	-71.504***
	(8.514)	(0.590)	(6.293)	(0.802)	(5.384)	(0.823)	(4.691)
Income	7.019**	0.511**	3.658**	0.788***	2.501	0.456*	2.505*
	(2.877)	(0.185)	(1.672)	(0.233)	(1.593)	(0.240)	(1.241)
Intercept	411.359***	12.883***	38.015	-70.607***	-84.467***	17.984***	-224.914***
	(45.637)	(2.402)	(27.232)	(3.142)	(21.804)	(3.270)	(21.972)
N	8722	8722	8722	8722	8722	8722	8722
p-value	0.00	0.82	0.01	0.00	0.00	1.00	0.00
Overall R^2	0.00	0.04	0.00	0.06	0.01	0.01	0.05
Household FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Woreda-Round FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: *, **, and *** denote significance at the 90, 95, and 99 percent levels.

Table 4: Price Risk Aversion Matrix for Seven Commodities Over Four Rounds (*N* = 8722)

$\mathbf{A}_7 =$

	Maize	Coffee	Barley	CookingOil	Wheat	Beans	Sorghum
Maize	**0.016**	0.019*	0.008	0.035***	0.014	0.012	0.022**
Coffee	0.018*	**0.036***	0.019*	0.040***	0.020*	0.023**	0.025**
Barley	– 0.011	0.019*	**0.025**	0.020*	0.010	0.011	0.010
Cooking Oil	0.036***	0.052***	0.013	**0.054***	0.023**	0.024**	0.025**
Wheat	– 0.016	0.025**	0.022***	0.017	**0.037***	0.007	0.013
Beans	– 0.010	0.031***	0.011	0.029***	– 0.006	**0.019***	– 0.009
Sorghum	0.015	0.028***	0.014	0.015	– 0.006	– 0.010	**0.052***

Note: Coefficients in bold are own-price coefficients of price risk aversion, and *, **, and *** denote significance at the 90, 95, and 99 percent levels.

Table 5: Symmetry of the Matrix of Price Risk Aversion Coefficients and Significance of Diagonal and Off-Diagonal Coefficients Test Results

Matrix of Price Risk Aversion Coefficients	**Test Statistic**	***p*-value**
$\mathbf{A}_3$	$F(3, 8719) = 0.56$	0.65
$\mathbf{A}_4$	$F(6, 8716) = 0.43$	0.86
$\mathbf{A}_5$	$F(10, 8712) = 0.74$	0.69
$\mathbf{A}_6$	$F(15, 8707) = 0.62$	0.86
$\mathbf{A}_7$	$F(21, 8701) = 0.61$	0.92
Joint Significance (All Coefficients)	$F(49, 8721) = 3.29$	0.00
Joint Significance (Diagonal Coefficients)	$F(7, 8721) = 12.03$	0.00
Joint Significance (Off-Diagonal Coefficients)	$F(42, 8721) = 2.60$	0.00

Note: The results of the symmetry tests were the same when assuming $R = 1$ or $R = 3$.

Table 6: Robustness Checks for Symmetry and Significance Under Different Assumptions for Relative Risk Aversion *R*

	***R* = 1**		***R* = 3**	
Matrix of Price Risk Aversion Coefficients	**Test Statistic**	***p*-value**	**Test Statistic**	***p*-value**
$\mathbf{A}_3$	$F(3, 8719) = 0.56$	0.65	$F(3, 8719) = 0.56$	0.65
$\mathbf{A}_4$	$F(6, 8716) = 0.43$	0.86	$F(6, 8716) = 0.43$	0.86
$\mathbf{A}_5$	$F(10, 8712) = 0.74$	0.69	$F(10, 8712) = 0.74$	0.69
$\mathbf{A}_6$	$F(15, 8707) = 0.62$	0.86	$F(15, 8707) = 0.62$	0.86
$\mathbf{A}_7$	$F(21, 8701) = 0.61$	0.92	$F(21, 8701) = 0.61$	0.92
Joint Significance (All Coefficients)	$F(49, 8721) = 3.28$	0.00	$F(49, 8721) = 3.29$	0.00
Joint Significance (Diagonal Coefficients)	$F(7, 8721) = 12.03$	0.00	$F(7, 8721) = 12.03$	0.00
Joint Significance (Off-Diagonal Coefficients)	$F(42, 8721) = 2.59$	0.00	$F(42, 8721) = 2.60$	0.00

Note: The results of the symmetry tests were the same when assuming $R = 1$ or $R = 3$.

Table 7: Slutsky Matrix of the Household for Seven Commodities Over Four Rounds (*N* = 8722)

$S_7(p,w) =$

	Maize	Coffee	Barley	CookingOil	Wheat	Beans	Sorghum
Maize	**- 620.37*****	−104.80***	−689.70***	318.44***	−1317.05***	−454.46***	−9.91***
Coffee	−24.03***	**- 1.51*****	−52.15***	10.89***	−19.00***	−7.17***	−30.43***
Barley	−164.97***	−107.63***	**- 401.65*****	65.52***	16.15***	−68.82***	−261.33***
Cooking Oil	−39.93***	−22.34***	−77.39***	**- 4.82*****	−42.16***	−1.69	2.13***
Wheat	−197.11***	−15.52***	−190.25***	−11.40***	**- 138.39*****	57.50***	−162.83***
Beans	−6.75	−10.80***	−54.42***	4.44***	−8.80***	**- 11.13*****	−16.49***
Sorghum	−15.97	−180.74***	−552.24***	−53.43***	579.84***	−579.88***	**- 515.09*****

Note: Coefficients in bold are own-price Slutsky substitution terms, and *, **, and *** denote significance at the 90, 95, and 99 percent levels.

Table 8: Symmetry of the Slutsky Matrix and Significance of Diagonal and Off-Diagonal Coefficients Test Results

Slutsky Matrix	Test Statistic	*p*-value
$S_3(p,w)$	$F(2, 8720) = 97.00$	0.00
$S_4(p,w)$	$F(3, 8719) = 4145.96$	0.00
$S_5(p,w)$	$F(4, 8718) = 4590.82$	0.00
$S_6(p,w)$	$F(5, 8717) = 3905.92$	0.00
$S_7(p,w)$	$F(7, 8715) = 1.23e^4$	0.00
Joint Significance (All Coefficients)	$F(7, 8715) = 6478.15$	0.00
Joint Significance (Diagonal Coefficients)	$F(7, 8715) = 366.82$	0.00
Joint Significance (Off-Diagonal Coefficients)	$F(7, 8715) = 6468.04$	0.00

Note: Several constraints were dropped in each test of joint significance due to collinearity.

Appendix A – Arrow-Pratt Measures of Price Risk Aversion

In section 2, we have derived the matrix A of coefficients of price risk aversion over multiple commodities, which is such that

$$A_{ij} = -\frac{V_{p_i p_j}}{V_y} = -\frac{M_i}{p_j}\left[\beta_j(\eta_j - R) + \varepsilon_{ij}\right]. \tag{A.1}$$

Although the sign of A_{ij} has an interpretation that is analogous to the usual Arrow-Pratt coefficients of absolute and relative risk aversion (i.e., $A_{ij} > 0$ indicates the household is price risk-averse; $A_{ij} = 0$ indicates it is price risk-neutral; and $A_{ij} < 0$ indicates it is price risk-loving), the coefficient A_{ij} itself cannot be interpreted as analogous to the Arrow-Pratt measures. In this section, we derive (i) the matrix of coefficients of absolute price risk aversion (APRA); and (ii) the matrix of coefficients of relative price risk aversion (RPRA).

To construct a measure of price risk aversion that is analogous to the Arrow-Pratt measures of absolute or relative (income) risk aversion, one should first divide V_{pp} by V_p. For the diagonal terms (i.e., $V_{p_i p_i}$), this poses no problem, since APRA is given by $\frac{V_{p_i p_i}}{V_{p_i}} p_i$ and RPRA is given by $\frac{V_{p_i p_i}}{V_{p_i}}$. For the off-diagonal terms (i.e., $V_{p_i p_j}$), however, the choice of denominator matters a great deal, since $\frac{V_{p_i p_j}}{V_{p_i}} \neq \frac{V_{p_i p_j}}{V_{p_j}}$. Likewise, for RPRA, $\frac{V_{p_i p_j}}{V_{p_i}} p_i \neq \frac{V_{p_i p_j}}{V_{p_j}} p_j$.

The solution for APRA, then, is to post-multiply V_{pp} by a vector $a = [1/V_{p_1}, ..., 1/V_{p_K}]'$ such that, with two commodities:

$$\text{APRA} \equiv V_{pp} \cdot a = \begin{bmatrix} V_{p_1 p_1} & V_{p_1 p_2} \\ V_{p_2 p_1} & V_{p_2 p_2} \end{bmatrix} \begin{bmatrix} 1/V_{p_1} \\ 1/V_{p_2} \end{bmatrix} = \begin{bmatrix} \frac{V_{p_1 p_1}}{V_{p_1}} & \frac{V_{p_1 p_2}}{V_{p_2}} \\ \frac{V_{p_2 p_1}}{V_{p_1}} & \frac{V_{p_2 p_2}}{V_{p_2}} \end{bmatrix}. \qquad \text{(A.2)}$$

Likewise, the solution for RPRA is to post-multiply V_{pp} by a vector $b = [p_1/V_{p_1}, ..., p_K/V_{p_K}]'$ such that, with two commodities:

$$\text{RPRA} \equiv V_{pp} \cdot b = \begin{bmatrix} V_{p_1 p_1} & V_{p_1 p_2} \\ V_{p_2 p_1} & V_{p_2 p_2} \end{bmatrix} \begin{bmatrix} p_1/V_{p_1} \\ p_2/V_{p_2} \end{bmatrix} = \begin{bmatrix} \frac{V_{p_1 p_1}}{V_{p_1}} p_1 & \frac{V_{p_1 p_2}}{V_{p_2}} p_2 \\ \frac{V_{p_2 p_1}}{V_{p_1}} p_1 & \frac{V_{p_2 p_2}}{V_{p_2}} p_2 \end{bmatrix}.$$

At this point, it should be obvious that the matrices of APRA and RPRA will generally not be symmetric.

Empirically, however, how can the matrices of APRA and RPRA be recovered from knowledge of the parameters estimated above? In other words, how to recover V_{p_i} ?

The short answer is that it is not possible to do so given that V_{p_i} is unobservable. But recall that

$$A_{ij} = -\frac{1}{V_y} \cdot V_{p_i p_j} \Leftrightarrow V_{p_i p_j} = -V_y \cdot A_{ij} \tag{A.3}$$

By Roy's Identity,

$$-\frac{V_{p_i}}{V_y} = M_i \Leftrightarrow V_{p_i} = -V_y M_i , \tag{A.4}$$

which implies that

$$\frac{V_{p_i p_j}}{V_{p_i}} = \frac{-V_y \cdot A_{ij}}{-V_y M_i} = \frac{A_{ij}}{M_i} \tag{A.5}$$

so that

$$APRA_{ij} = \frac{V_{p_i p_j}}{V_{p_i}} = \frac{V_y}{p_j}\left[\beta_j(\eta_j - R) + \varepsilon_{ij}\right] \tag{A.6}$$

and

$$RPRA_{ij} = \frac{V_{p_i p_j}}{V_{p_i}} \cdot p_i = \frac{V_y}{p_j}\left[\beta_j(\eta_j - R) + \varepsilon_{ij}\right]p_i \qquad \text{(A.7)}$$

are the true price analogs to the usual Arrow-Pratt coefficients of absolute and relative (income) risk aversion. Our framework thus allows deriving three distinct matrices of price risk aversion: (i) A; (ii) APRA; and (iii) RPRA.

Appendix B – Proofs of Propositions

Proof of Proposition 1 We know from first principles that symmetry of the Slutsky matrix implies that

$$\frac{\partial x_i}{\partial p_j} + \frac{\partial x_i}{\partial y} x_j = \frac{\partial x_j}{\partial p_i} + \frac{\partial x_j}{\partial y} x_i . \qquad \text{(B.1)}$$

By Roy's Identity, the above statement can be rewritten as

$$\frac{\partial}{\partial p_j}\left(-\frac{V_{p_i}}{V_y}\right) + \frac{\partial}{\partial y}\left(-\frac{V_{p_i}}{V_y}\right) \cdot \left[-\frac{V_{p_j}}{V_y}\right] = \frac{\partial}{\partial p_i}\left(-\frac{V_{p_j}}{V_y}\right) + \frac{\partial}{\partial y}\left(-\frac{V_{p_j}}{V_y}\right) \cdot \left[-\frac{V_{p_i}}{V_y}\right]. \qquad \text{(B.2)}$$

Which, when writing out the second-order partials explicitly, is equivalent to

$$-\left(\frac{V_{p_i p_j} V_y - V_{y p_j} V_{p_i}}{V_y^2}\right) + \left(\frac{V_{p_i y} V_y - V_{yy} V_{p_i}}{V_y^2}\right) \cdot \left[\frac{V_{p_j}}{V_y}\right] =$$

$$-\left(\frac{V_{p_j p_i} V_y - V_{y p_i} V_{p_j}}{V_y^2}\right) + \left(\frac{V_{p_j y} V_y - V_{yy} V_{p_j}}{V_y^2}\right) \cdot \left[\frac{V_{p_i}}{V_y}\right]. \qquad \text{(B.3)}$$

This last equation can then be arranged to show that

$$\left(V_{p_i p_j} - V_{p_j p_i}\right) V_y = V_{y p_j} V_{p_i} - V_{p_j y} V_{p_i} - V_{y p_i} V_{p_j} + V_{p_i y} V_{p_j} . \qquad \text{(B.4)}$$

By Young's Theorem, we know that $V_{p_i p_j} = V_{p_j p_i}$, that $V_{yp_i} V_{p_j} = V_{p_i y} V_{p_j}$, and that $V_{yp_j} = V_{p_j y}$, so both sides of the previous equation are identically equal to zero. In other words, symmetry of the Slutsky matrix implies and is implied by symmetry of the matrix A of price risk aversion coefficients.◄

Proof of Corollary Symmetry of the matrix A of price risk aversion coefficients only requires that $V_{p_i p_j}$ not be statistically significantly different from $V_{p_j p_i}$. Symmetry of the Slutsky matrix, however, requires (i) that $V_{p_i p_j}$ not be statistically significantly different from $V_{p_j p_i}$; (ii) that $V_{yp_i} V_{p_j}$ not be statistically significantly different from $V_{p_i y} V_{p_j}$; and (iii) that V_{yp_j} not be statistically significantly different from $V_{p_j y}$. Clearly then, symmetry of the Slutsky matrix imposes more structure on the data than symmetry of the matrix A of price risk aversion coefficient does. As a result, it is easier to reject symmetry of the Slutsky matrix than it is to reject symmetry of the matrix of price risk aversion coefficients.◄