The Welfare Impacts of Price Fluctuations: Evidence from Rural Ethiopia*

Marc F. Bellemare[1]
Christopher B. Barrett[2]
David R. Just[3]

January 29, 2010

Abstract

Many governments throughout history have tried to stabilize commodity prices based on the widespread belief that households – especially the poor – value price stability. We extend the existing microeconomic theory literature to derive an estimable matrix of the coefficients of price risk aversion and associated willingness to pay measures over multiple commodities. Using longitudinal household-level data from Ethiopia, we then estimate that the average household would be willing to pay almost 20 percent of its income to eliminate price fluctuations among the seven most important staples in the data. We further show that not everyone benefits from price stabilization and that the welfare gains from eliminating price fluctuations would be concentrated in the upper half of the income distribution, making price stabilization a regressive policy in this context.

JEL Classification Codes: D13, D80, E64, O12, Q12
Keywords: Price Fluctuations, Price Stabilization, Price Risk, Risk and Uncertainty.

---

* We thank Pascale Schnitzer for excellent research assistance as well as Zack Brown, Stefan Dercon, and Vivian Hoffman for help with the data. We also thank Bob Myers and Wally Thurman as well as seminar audiences at Leuven, Michigan State, Namur, NC State, Ottawa, Paris, Tennessee, Texas A&M, Toulouse, and Western Michigan and conference participants at the 2008 Southern Economics Association meetings and the 2009 SCC-76 "Economics and Management of Risk in Agriculture and Natural Resources" for useful comments and suggestions. Part of this paper was written while Bellemare was visiting the University of Namur's Centre de recherche en économie du développement, whose generous financial support he would like to acknowledge. All remaining errors are ours.

[1] Corresponding Author and Assistant Professor, Duke University, Box 90312, Durham, NC 27708-0312, (919) 613-9382, marc.bellemare@duke.edu.
[2] Steven B. and Janice G. Ashley Professor of Applied Economics and International Professor, Cornell University, Ithaca, NY 14853-7801, (607) 255-4489, cbb2@cornell.edu.
[3] Associate Professor, Cornell University, Ithaca, NY 14853-7801, (607) 255-2086, drj3@cornell.edu.

1. Introduction

Throughout history and all over the world, governments have treated commodity price stability as an important goal of economic policy. Using a host of policy objectives from buffer stocks to administrative pricing and from variable tariffs to marketing boards, these governments have had very limited success in stabilizing prices. With the food crisis of 2007-2008, however, commodity price volatility has rekindled widespread government interest in price stabilization, especially as regards food and energy commodities.

State intervention with respect to domestic staples prices commonly arises because households are widely believed to value price stability; because the poor are widely perceived to suffer disproportionately from food price instability; and because futures and options markets for hedging against price risk are commonly unavailable to consumers and poor producers given the minimum scale of transactions required by the relevant exchanges (Newbery, 1989; Timmer, 1989). Given the policy importance of the topic, and although economists have commonly questioned the net economic benefit of government price stabilization interventions (Newbery and Stiglitz, 1981; Krueger et al., 1988; Knudsen and Nash, 1990), the theoretical and empirical toolkit for understanding how household welfare is affected by price fluctuations is somewhat dated and surprisingly limited.

The effects of price fluctuations – which we also refer to as “price risk” in this paper − on producer behavior and welfare have been well-explored in the theoretical literature.

Output price uncertainty generally causes firms to employ fewer inputs, foregoing expected profits in order to hedge against price fluctuations (Baron, 1970; Sandmo, 1971). The analysis of commodity price fluctuations has been extended theoretically to individual consumers (Deschamps, 1973; Hanoch, 1977; Turnovsky et al., 1980; Newbery and Stiglitz, 1981; Newbery, 1989) and to agricultural households both theoretically (Finkelshtain and Chalfant, 1991 and 1997) and empirically (Barrett, 1996). These analyses, however, have focused largely on a single staple commodity. Although Turnovsky et al. (1980) have considered the price fluctuations of multiple commodities, they did so only theoretically. Given that indirect utility functions – the usual measure of welfare in microeconomic theory – are defined over both income and a vector of prices, the literature's heavy focus on income risk, at most extended to a single stochastic price, paints a very incomplete picture of total (i.e., income and prices) attitudes to risk.

The central contribution of this paper is to combine the theoretical framework of Turnovsky et al. (1980) with the empirical framework of Finkelshtain and Chalfant (1991, 1997) and Barrett (1996) in order to derive an estimable matrix of price risk aversion and associated willingness to pay (WTP) measures for price stabilization over multiple commodities, and then to demonstrate the empirical implications of the theory as it applies to rural Ethiopian households who both produce and consume several commodities characterized by stochastic prices.

The matrix of price risk aversion coefficients we derive and estimate reflects price risk premia with respect to the covariance matrix of prices faced by the household, yielding not only the usual own-price risk aversion coefficients on the diagonal (i.e., the effect of price variance on household welfare), but also the off-diagonal cross-price risk aversion coefficients (i.e., the effect of the price covariances on household welfare). These off-diagonal terms have so far been overlooked in the literature, although they have an intuitive interpretation and are necessary to understanding behavior and welfare with respect to multivariate price risk. Indeed, even when focusing on price risk for a single commodity, ignoring cross-price risk aversion coefficients leads to biased estimates of the effect of price risk. Staple prices rarely, if ever, fluctuate independently of one another. Based on the matrix of price risk aversion coefficients, we then derive formulae for the household's WTP to stabilize at their means both the price of individual commodities and the prices of a set of commodities.

Using panel data on rural Ethiopian households, we then estimate the matrix of price risk aversion and compute WTP estimates. These estimates show that the average household is willing to give up 19 percent of its income to stabilize the price of the seven most important staple commodities in the sample. Nonparametric analysis suggests that, contrary to conventional wisdom, the welfare gains of price stabilization in these data are only positive in upper half of the income distribution, and that they are negative in a significant fraction of the lower half of the income distribution, which would make price stabilization a regressive policy intervention in these data. Lastly, we conduct an *ex ante* analysis of an alternative to strict price stabilization policy, one in which the households who are price risk-averse receive a transfer payment to compensate them for the welfare loss they suffer due to price risk exposure, but which leaves households who benefit from price fluctuations unaffected. We demonstrate that this would be Pareto superior to fixing prices.

The rest of this paper is organized as follows. Based on the theoretical work of Turnovsky et al. (1980), we extend Barrett's (1996) empirical approach to the estimation of price risk aversion coefficients to the multiple commodity case. We derive the matrix of price risk aversion and its properties in section 2. In section 3, we present the data and briefly discuss descriptive statistics. We then develop a reduced form empirical framework to estimate the matrix of price risk aversion coefficients and discuss identification in section 4. In section 5 we estimate own- and cross-price risk aversion coefficients, construct the matrix of price risk aversion coefficients, test the restrictions of the theory, compute and analyze household WTP estimates for price stabilization, and casually explore a price risk compensation scheme as an alternative to pure price stabilization policy. We conclude in section 6 by discussing the research and policy implications of our findings.

2. Theoretical Framework

This section develops a simple two-period agricultural household model (AHM) and derives the household's matrix of own- and cross-price risk aversion coefficients for the multiple-commodity case. This is the most parsimonious model possible, as we need a framework that encompasses both consumer and producer behavior while an interest in price instability requires, at a minimum, a two period model, with at least one period in which agents make decisions subject to uncertainty with respect to prices, both in levels and in relation to incomes and other prices. We then derive some key properties of the price risk aversion matrix and relate it to the Slutsky matrix, which yields implications that we test in section 5. Lastly, and more importantly, we analytically derive measures of household willingness to pay to stabilize the prices of one or more commodities.

2.1 Agricultural Household Model

The derivations in this section closely follow those in Barrett (1996), who in turn builds on Turnovsky et al.'s (1980) work on individual consumers and Finkelshtain and Chalfant's (1991) work on price risk in the context of the AHM. In what follows in this subsection, we report the basics of the model. Readers interested in more detailed explanations and derivations of these findings should consult those prior works.

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, and leisure ℓ. Assume function $U(\cdot)$ is quasiconcave but concave in each of its arguments, and that the Inada condition $\frac{\partial U}{\partial x}|_{x=0} = \infty$ applies with respect to each argument *x*. All *K* observed goods and the unobserved good can, in principle, be produced or consumed by the household, which draws on its endowments of labor time and land and an exogenously given production technology defined over land, labor and a composite of other variable inputs. The household faces the usual cash and time budget constraints and may receive some unearned income.

The household maximizes its welfare over two periods, making production decisions in the first period, when all product prices are unknown but input prices z are known. While Turnovsky (1978) noted how different qualitative results obtain depending on whether price uncertainty arises due to an additive or multiplicative error term, our framework allows us to assume nothing about the shape of price uncertainty and let the data speak for themselves.

By Epstein's (1975) duality result, we can use the household's expected indirect utility function $V(\cdot)$, which is homogeneous of degree zero in prices and income, to solve the household's optimization problem. We thus set the price of the unobserved commodity as numéraire. Lastly, we assume that the household is (income) risk-averse, in the sense that $\frac{\partial^2 V}{\partial y^2} = V_{yy} < 0$, where y represents household total income.

Using the household's expected indirect utility function, Barrett (1996) then solves the household's maximization problem and derives an expression of household price risk aversion in the case of a single commodity. We now extend that framework to the case of multiple commodities whose prices are stochastic and derive the household's matrix of own- and cross-price risk aversion coefficients.

2.2 Price Risk Aversion over Multiple Commodities

Let $V(p, y)$ denote the household's indirect utility function. The vector $p = (p_1, ..., p_K)$ is the vector of commodity prices faced by the household over the observed commodities, while 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

$$\text{sign}[Cov(V_y, p_i)] = \text{sign}(V_{yp_i}). \tag{1}$$

Moreover, let $M_i = s_i(z, p) - x_i(p, y) = M_i(z, p, y)$ be the marketable surplus of commodity *i*, where $s_i(\cdot)$ is the household supply of commodity *i*, which depends on input and commodity prices, and $x_i(\cdot)$ is its Marshallian demand for commodity *i*, which depends on commodity prices and income. By Roy's identity, i.e., $M_i = -\frac{\partial V / \partial p_i}{\partial V / \partial y}$,[4] we have that

$$V_y = -\frac{V_{p_i}}{M_i} = -\frac{V_{p_j}}{M_j}, \tag{2}$$

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

$$V_{yp_j} = \left(\frac{V_{p_i p_j}}{M_i} - \frac{V_{p_i}}{M_i^2} \frac{\partial M_i}{\partial p_j} \right) = \frac{1}{M_i} \left\{ V_{p_i p_j} - \frac{\partial M_i}{\partial p_j} V_y \right\}. \tag{3}$$

We also have that

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

[4] One can apply Roy's identity to the marketable surplus equation given that it is both additive and convex. See also Finkelshtain and Chalfant (1991).

which implies that

$$V_{p_i p_j} = M_i V_{yp_j} + V_y \frac{\partial M_i}{\partial p_j}, \tag{5}$$

which, in turn, implies that

$$V_{p_i y} = M_i V_{yy} + V_y \frac{\partial M_i}{\partial y} = V_{yp_i}, \tag{6}$$

where the last equation is the result of applying Young's theorem on the symmetry of second derivatives, which requires that (i) $V(\cdot)$ be a differentiable function over (p, y); and (ii) its cross-partials exist and be continuous at all points on some open set.

Replacing V_{yp_i} by equation 6 in equation 5 yields

$$V_{p_i p_j} = M_i \left\{ M_j V_{yy} + V_y \frac{\partial M_j}{\partial y} \right\} + V_y \frac{\partial M_i}{\partial p_j}. \tag{7}$$

Then, we have that

$$V_{p_i p_j} = M_i M_j V_{yy} + M_i V_y \frac{\partial M_j}{\partial y} + V_y \frac{\partial Mi}{\partial p_j}. \tag{8}$$

Multiplying the first term by $V_y y / V_y y$ yields (9)

$$V_{p_i p_j} = -\frac{M_i M_j R V_y}{y} + M_i V_y \frac{\partial M_j}{\partial y} + V_y \frac{\partial Mi}{\partial p_j}, \tag{10}$$

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

$$V_{p_i p_j} = -\frac{M_i M_j R V_y}{y} + M_i V_y \eta_j \frac{M_j}{y} + V_y \varepsilon_{ij} \frac{M_i}{p_j}, \tag{11}$$

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 11 is thus equivalent to

$$V_{p_i p_j} = M_i V_y \left[-\frac{M_j R}{y} + \eta_j \frac{M_j}{y} + \varepsilon_{ij} \frac{1}{p_j} \right]. \tag{12}$$

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

$$V_{p_i p_j} = \frac{M_i V_y}{p_j} \left[-R\beta_j + \eta_j \beta_j + \varepsilon_{ij} \right], \tag{13}$$

where β_j is the budget share of commodity *j*. When simplified, equation 13 is such that

$$V_{p_i p_j} = \frac{M_i V_y}{p_j} \left[\beta_j (\eta_j - R) + \varepsilon_{ij} \right]. \tag{14}$$

Consequently, if $M_i = 0$, the household is indifferent to fluctuations in the price of good *i* and to cofluctuations in the prices of goods *i* and *j* since its autarky from the market leaves it unaffected at the margin by price volatility. Applying Young's theorem again yields the following equation:

$$V_{p_i p_j} = \frac{M_i V_y}{p_j} \left[\beta_j (\eta_j - R) + \varepsilon_{ij} \right] = \frac{M_j V_y}{p_i} \left[\beta_i (\eta_i - R) + \varepsilon_{ji} \right] = V_{p_j p_i}. \tag{15}$$

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

$$V_{pp} = \begin{bmatrix} V_{p_1 p_1} & V_{p_1 p_2} & \cdots & V_{p_1 p_K} \\ V_{p_2 p_1} & V_{p_2 p_2} & \cdots & V_{p_2 p_K} \\ \vdots & \vdots & \ddots & \vdots \\ V_{p_K p_1} & V_{p_K p_2} & \cdots & V_{p_K p_K} \end{bmatrix}, \tag{16}$$

is symmetric. From the V_{pp} matrix, we can derive matrix A of price risk aversion coefficients, which is as follows:

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

$$= \begin{bmatrix} A_{11} & A_{12} & \cdots & A_{1K} \\ A_{21} & A_{22} & \cdots & A_{2K} \\ \vdots & \vdots & \ddots & \vdots \\ A_{K1} & A_{K2} & \cdots & A_{KK} \end{bmatrix}, \tag{17}$$

where

$$A_{ij} = -\frac{M_i}{p_j}\left[\beta_j(\eta_j - R) + \varepsilon_{ij}\right]. \tag{18}$$

Matrix A has a straightforward interpretation, as developed in the scalar stochastic price case (Barrett 1996). The diagonal elements are analogous to Pratt's (1964) coefficient of absolute (income) risk aversion with respect to income variability, but here with respect to prices. Thus, $A_{ii} > 0$ implies that welfare is decreasing in the volatility of the price of *i*, i.e., that the household is price risk-averse (or a hedger) over *i*; $A_{ii} = 0$ implies that welfare is unaffected by the volatility of the price of *i*, i.e., that the household is price risk-neutral; and $A_{ii} < 0$ implies that welfare is increasing in the volatility of the price of *i*, i.e., that the household is price risk-loving (or a speculator) over *i*.[5] Price risk-aversion is the classic concern of the literature on commodity price stabilization (Deschamps, 1973; Hanoch, 1974, Turnovsky, 1978; Turnovsky et al., 1980; 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.

[5] The hedger-speculator terminology is from Hirshleifer and Riley (1992), who apply it to the Keynes-Hicks theory of futures markets.

Consequently, A_{ij} > (<) 0 implies that greater volatility in price *j* reduces (increases) welfare associated with the net consumption of good *i*, or that the household stands to gain from hedging against (speculating over) covariance in the prices of goods *i* and *j*. The price risk aversion coefficient matrix thus speaks directly to the welfare effects of and household preferences with respect to multivariate price risk. Intuitively, the diagonal terms can be interpreted as the (direct) effect on household welfare of the variance in the price of a single good, and the off-diagonal terms can be interpreted as the (indirect) effect on household welfare of the covariance between the prices of two goods.

Perhaps more importantly, there is no theoretical restriction on the sign of any element of *A*. The theory, however, implies a testable symmetry restriction on the estimated price risk aversion coefficients. With adequate data, one can test the following null hypothesis:

$$H_0 : A_{ij} = A_{ji} \text{ for all } i \neq j, \quad (19)$$

which represents $K(K-1)/2$ testable restrictions. The next section characterizes the relationship between the price risk aversion matrix A and the Slutsky matrix and shows how a test of the symmetry of A is thus a test of the rationality of the household.

2.3 Relationship between the Price Risk Aversion and Slutsky Matrices

The derivations above raise a natural question: What is the relationship between the price risk aversion matrix and the Slutsky matrix? Let $M_i(z, p, y)$ 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)

$$S_{ij}(p,y) = \frac{\partial M_i}{\partial p_j} + \frac{\partial M_i}{\partial y} M_j = B_{ij} + C_{ij}, \tag{20}$$

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

$$A_{ij} = M_i \left[\frac{1}{M_j} C_{jj} - \frac{R}{y} + B_{ij} \right]. \tag{21}$$

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 (i.e., $V_{p_i p_j}$), and this change 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 to both S_{jj} and S_{ij}, there does not exist a one-to-one correspondence between the elements of matrices *A* and *S*. This can be seen by rewriting the last expression as

$$A = \begin{pmatrix} M_1 & 0 & 0 \\ 0 & \ddots & 0 \\ 0 & 0 & M_K \end{pmatrix} \left[\begin{pmatrix} \frac{\partial M_1}{\partial y} & \cdots & \frac{\partial M_K}{\partial y} \\ \vdots & \ddots & \vdots \\ \frac{\partial M_1}{\partial y} & \cdots & \frac{\partial M_K}{\partial y} \end{pmatrix} + \begin{pmatrix} \frac{\partial M_1}{\partial p_1} & \cdots & \frac{\partial M_1}{\partial p_K} \\ \vdots & \ddots & \vdots \\ \frac{\partial M_K}{\partial p_1} & \cdots & \frac{\partial M_K}{\partial p_K} \end{pmatrix} \right.$$

$$\left. - \begin{pmatrix} \frac{R}{y} & \cdots & \frac{R}{y} \\ \vdots & \ddots & \vdots \\ \frac{R}{y} & \cdots & \frac{R}{y} \end{pmatrix} \right]. \tag{22}$$

In other words, one cannot recover the Slutsky matrix from the matrix of price risk aversion coefficients. The two, however, are related, and the derivations above lead to the following result.

Proposition 1: Under the preceding assumptions and if the cross-partials of the household's indirect utility function exist and are continuous at all points on some open set, symmetry of the matrix of price risk aversion coefficients is equivalent to symmetry of the Slutsky matrix.

Proof: Symmetry of the Slutsky matrix implies that

$$\frac{\partial M_i}{\partial p_j} + \frac{\partial M_i}{\partial y} M_j = \frac{\partial M_j}{\partial p_i} + \frac{\partial M_j}{\partial y} M_i \,. \tag{23}$$

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

$$\frac{\partial}{\partial p_j}\left(-\frac{V_{p_i}}{V_y}\right) + \frac{\partial}{\partial y}\left(-\frac{V_{p_i}}{V_y}\right)\cdot\left[-\frac{V_{p_j}}{V_y}\right] = \frac{\partial}{\partial p_i}\left(-\frac{V_{p_j}}{V_y}\right) + \frac{\partial}{\partial y}\left(-\frac{V_{p_j}}{V_y}\right)\cdot\left[-\frac{V_{p_i}}{V_y}\right], \tag{24}$$

which, once the second-order partials are written explicitly, is equivalent to

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

$$-\left(\frac{V_{p_j p_i} V_y - V_{y p_i} V_{p_j}}{V_y^2}\right) + \left(\frac{V_{p_j y} V_y - V_{yy} V_{p_j}}{V_y^2}\right)\cdot\left[\frac{V_{p_i}}{V_y}\right]. \tag{25}$$

This last equation can then be arranged to show that

$$\left(V_{p_i p_j} - V_{p_j p_i}\right) V_y = V_{y p_j} V_{p_i} - V_{p_j y} V_{p_i} - V_{y p_i} V_{p_j} + V_{p_i y} V_{p_j} \,. \tag{26}$$

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.■

The symmetry of the Slutsky matrix and the symmetry of the matrix of price risk aversion coefficients have the same empirical content in that they both embody the rationality of the household. But symmetry of the Slutsky matrix should be easier to reject than symmetry of the matrix of price risk aversion given that it imposes much more structure on the data than symmetry of the matrix of price risk aversion. Indeed, 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 $V_{p_j y}$. As a result, it should be easier to reject symmetry of the Slutsky matrix than it is to reject symmetry of the matrix of price risk aversion coefficients, simply because the former imposes more restriction on the data.

2.4 Willingness to Pay for Price Stabilization

Policymakers routinely try to stabilize one or more staple good prices, but what are the welfare effects of such efforts? This subsection derives the appropriate WTP measures necessary to establish the welfare gains from partial price stabilization, i.e., from stabilizing one or more commodity prices.[6]

In order to tackle this question with respect to K prices, one first needs to compute the total WTP, i.e., the WTP for K commodities. Then,

$$WTP = \frac{V(E(p), y) - E(V(p, y))}{V_y} = \frac{E[V(E(p), y) - V(p, y)]}{V_y}. \tag{27}$$

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

$$WTP \approx \frac{E\left[-V_p(E(p), y)(p - E(p)) - \frac{1}{2}(p - E(p))' V_{pp}(E(p), y)(p - E(p))\right]}{V_y}. \tag{28}$$

In other words,

$$WTP \approx -\frac{1}{2} \frac{E[(p - E(p))(p - E(p))' V_{pp}(E(p), y)(p - E(p))]}{V_y} \tag{29}$$

and so

$$WTP \approx -\frac{1}{2} \sum_{i=1}^{K} \sum_{j=1}^{K} \sigma_{ij} \frac{V_{p_i p_j}}{V_y} = \frac{1}{2} \sum_{i=1}^{K} \sum_{j=1}^{K} \sigma_{ij} A_{ij} , \tag{30}$$

where σ_{ij} is the covariance between prices i and j and A_{ij} is the coefficient of price risk aversion, as defined above. By symmetry of matrix A, the above is equivalent to

$$WTP \approx \frac{1}{2} \sum_{i=1}^{K} \sum_{j=1}^{K} \sigma_{ji} A_{ji} . \tag{31}$$

[6] The measures derived in this section are partial in the sense that they only stabilize prices for a subset of the (potentially infinite) set of commodities consumed and produced by the household, as it is essentially impossible to stabilize prices completely since the costs of stabilization increase exponentially with the degree of stabilization pursued (Knudsen and Nash, 1990).

These derivations provide the transfer payment a policymaker would need to make to the household in order to compensate it for the uncertainty over $(p_1,...,p_K)$. If instead one wishes to stabilize only one price *i*, the above derivations reduce to

$$WTP_i \approx \frac{1}{2}\left[\sigma_{ii}A_{ii} + \sum_{j\neq i}^{K}\sigma_{ij}A_{ij}\right], \tag{32}$$

and, by symmetry of matrix A and of the price covariance matrix, the above is equivalent to

$$WTP_i \approx \frac{1}{2}\left[\sigma_{ii}A_{ii} + \sum_{j\neq i}^{K}\sigma_{ji}A_{ji}\right]. \tag{33}$$

Because equations 32 and 33 are equivalent, the WTP for commodity *i* can be computed in two ways, i.e., via either the rows or the columns of matrix A. This provides the transfer payment a policymaker would need to make to the household in order to compensate it for the uncertainty over p_i. Finkelshtain and Chalfant (1997) introduced a similar measure, but their framework considered only one stochastic price, *de facto* ignoring the covariances between prices. Realistically, however, even the WTP for a single commodity *i* depends on the covariance between the price *i* and the prices of other commodities *j*. In other words, a price stabilization policy focusing solely on the price of commodity *i* would bias the estimated WTP for commodity *i*, unless $\sigma_{ij} = 0$ or $A_{ij} = 0$ for all $i \neq j$.

3. Data and Descriptive Statistics

We empirically demonstrate the core theory developed in the preceding section by estimating the price risk aversion coefficient matrix and household WTP for price stabilization. In order to do so, we use data from the publicly-available Ethiopian Rural Household Survey (ERHS),[7] which includes results from four rounds: 1994a, 1994b, 1995, and 1997. The ERHS data record both household consumption and production decisions over multiple years. The ERHS has a low attrition rate as well as a standardized survey instrument across the rounds we retain for analysis. The sample includes a total of 1494 households across 16 districts (*woredas*) with an attrition rate of only 2 percent across the four rounds selected for analysis (Dercon and Krishnan, 1998).[8] The average household in the data was observed 5.7 times over four rounds and three seasons;[9] only 7 households appear only once in the data. The estimations in this paper thus rely on a sample of 8556 observations.[10]

Given that many of the households in our data were autarkic with respect to several commodities (i.e., they neither bought or sold those commodities), for every season (i.e.,

---

[7] 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, but 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.

[8] Ethiopia is subdivided into eleven zones subdivided into *woredas*, which are roughly equivalent to US districts.

[9] Within-round variation in seasons occurred only in 1994a and 1997. Because the season was not specified for the 1994b and 1995 rounds, we cannot control for seasonality in the empirical analysis of section 5.

[10] The original data included several outliers when considering the marketable surpluses of the seven commodities retained for analysis, and these outliers caused certain percentage values (e.g., the WTP measures below) to lie far outside the 0 to 100 percent interval. As a remedy, for each of the seven marketable surpluses used below, we kept only the 99 percent confidence interval (i.e., ± 2.576 standard deviations) around the median, the mean being too sensitive to outliers. We thus dropped 188 observations.

the time period we consider in these data) in which a household is neither a net buyer nor 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 coffee, maize, horse beans, barley, wheat, teff, and sorghum, i.e., the top seven staple commodities when considering the proportion of observations with a nonzero marketable surplus.

Table 1 presents descriptive statistics for these seven commodities. A positive (negative) mean marketable surplus indicates that the average household is a net seller (buyer) of a commodity. The average household is a net buyer of every staple, and table 2 further characterizes the dependent variables by focusing on the nonzero marketable surplus observations and by comparing descriptive statistics between net buyers and net sellers. Except for maize and wheat, the purchases of the average net buyer household exceed the sales of the average net seller household. For every commodity, there are many households in both the net buyer, autarkic, and net seller categories, reflecting potentially heterogeneous welfare effects with respect to commodity prices in rural Ethiopia.

Table 3 lists the mean real (i.e., corrected for the consumer price index) price in Ethiopian birr for each of the seven commodities we study,[11] the average seasonal household income, and the average seasonal nonzero household income in the full sample. The income measure used in this paper is the sum of proceeds from crop revenues, off-farm income, and livestock sales. That said, average income from the aforementioned sources is different from zero in about 82 percent of cases, which

[11] As of writing, US$1 ≈ Birr 9.43.

explains why the average seasonal income of about $94 may seem low. When focusing on nonzero income, the average seasonal income increases to about $106.

Table 3 also presents the budget share of each staple commodity retained for analysis. Purchases of teff and coffee represent the largest budget shares in magnitude, with 21 and 15 percent of the average household budget, respectively. The purchases of staple crops such as maize, barley, and wheat come close, however, with 13, 12, and 11 percent, respectively, of the average household budget. The purchases of beans and sorghum come last, with 7 and 6 percent, respectively, of the average household budget.

Finally, because price covariances play an important role in computing household WTP for price stabilization, table 4 reports the variance-covariance matrix for the prices of the seven staple commodities retained for analysis.

4. Empirical Framework

As defined previously, a household's marketable surplus of a given commodity *i*, $M_i(z, p, y)$, is the quantity harvested of that commodity net of the quantity purchased and the household's consumption of its own harvest, a reduced form function of input and output prices and household income. Our data include commodity prices and allow us to compute household income, but include only village-level average wage as an input price. Given that all households in an area face common market prices at the same time, however, we use *woreda*-round fixed effects to control for the vector of input prices faced by each household in each location in each period. Time invariant household fixed effects provide further control for household-specific transactions costs related to distance from the main *woreda* market, social relationships that may confer preferential pricing, etc.

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

$$M_{ik\ell t} = \alpha_i + \delta_i \ln y_{ik\ell t} + \phi_i \ln p_{ik\ell t} + \varphi_i \ln p_{jkt} + \lambda_{ik\ell} + \tau_{i\ell t} + \nu_{ik\ell t}, \qquad (34)$$

where *i* denotes a specific commodity,[12] *k* denotes the household, ℓ denotes the *woreda*, and *t* denotes the round; y denotes household income net of income from commodity *i*; p_i is a measure of the price of commodity *i*; p_j is a vector of measures of the prices of all (observed) commodities other than *i*; λ is a household-*woreda* fixed effect; τ is a *woreda*-round fixed effect that controls for the price of the unobservable composite consumer good as well as for input prices, among other things; and ν is a mean zero, iid error term.[13,14]

We estimate equation 34 over 1,494 households across four rounds and three seasons, clustering the standard errors at the *woreda* level. No household was observed over all four rounds and three seasons; the number of observations per household ranged from

[12] Subscripts on coefficients thus denote coefficients from specific commodity equations.
[13] 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 (MaCurdy and Pencavel, 1986). Robustness checks were conducted during preliminary empirical work in which 0.1 and 0.000001 were added instead of 0.001, with no significant change to the empirical results.
[14] We do not estimate the marketable surplus equations using seemingly unrelated regression (SUR) since SUR estimation brings no efficiency gain over estimating the various equations in the system separately when the dependent variables are all regressed on the same set of regressors.

one to six.[15] We also include as explanatory variables all commodity prices available in the data (i.e., coffee, maize, beans, barley, wheat, teff, sorghum, potatoes, onions, cabbage, milk, *tella*, sugar, salt, and cooking oil.)[16]

Computation of own- and cross-price elasticities, of the income-elasticity, and of the budget share of marketable surplus follows directly from equation 34. For example, to derive the estimated cross-price risk aversion coefficient $\hat{A}_{ij}$, one first computes budget share $\hat{\beta}_j = M_j p_j / y$ from the data; income elasticity $\hat{\eta}_j = \hat{\delta}_j / M_j$ from the data and the marketable surplus equation parameter estimate for commodity *j*; and cross-price elasticity $\hat{\varepsilon}_{ij} = \hat{\varphi}_i / M_i$ from the data and the marketable surplus equation parameter estimate for commodity *i*. One then combines these estimates to obtain the point estimate

$$\hat{A}_{ij} = \frac{M_i}{p_j}[\hat{\beta}_j(\hat{\eta}_j - R) + \hat{\varepsilon}_{ij}]. \quad (35)$$

Given that marketable surplus is often zero, we use the mean of M_j and M_i so as to compute elasticities, and later compute WTP, *at means*. Although it might be preferable to use mean elasticities, it is simply not possible to do so in these data.[17] The coefficient of relative risk aversion *R* can either be directly estimated, if the data allow, or assumed 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

[15] 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 (Verbeek and Nijman, 1992).
[16] *Tella* is a traditional Ethiopian beer made from teff and maize.
[17] Likewise, given that we use the household's income from non-agricultural sources as a proxy for total income *y* so as to avoid endogeneity problems, many households have a residual income of zero. In this case, we compute the estimated budget share by dividing by *y* + 0.001 (MaCurdy and Pencavel, 1986).

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 additional robustness checks on our empirical results.

4.1 Identification

Identification of ϕ and φ comes from the variation in own-price both within each household over time, since each household retained in the estimation is observed more than once, and between *woreda*-round, since prices are common to all households in the same region within the same *woreda* in the same round. Identification of δ comes from the intertemporal variation in income both within each household and between households within a round and *woreda*.

Since households are price takers for all commodities, all prices are exogenous in equation 34. Income, however, is likely endogenous, if only because a positive marketable surplus implies an additional source of revenue for the household. Unfortunately, the data do not include a credible instrument for income. Including both household and *woreda*-round fixed effects should purge the error term of a great deal of its prospective correlation with income, however, since a household's status as a net seller is primarily driven by preferences and by the transactions costs it faces (de Janvry et al., 1991; Goetz, 1992; Bellemare and Barrett, 2006), which are accounted for by the household fixed effect, as well as by climatic and other environmental fluctuations that effect production (Sherlund et al., 2002), which are largely accounted for by the *woreda*-round fixed effect. Finally, as discussed above, the potential endogeneity problem caused by the absence of input prices from the data is accounted for by our inclusion of *woreda*-round fixed effects, which control for local market conditions.[18]

Because many households have a marketable surplus of zero for several commodities, we test several estimates of the matrix of price risk aversion coefficients. We first test the A sub-matrix for the top three commodities consumed and produced by the households in our data (i.e., coffee, maize, and beans), and then test the sub-matrices defined by the top four, five, six, and seven commodities. With three different assumptions on relative risk aversion *R* and six different sub-matrices in each case, we conduct a total of 15 tests of the null hypothesis of symmetry of the matrix of price risk aversion. The consistency of results – which is mirrored by associated qualitative consistency in estimated WTP for price risk stabilization – provides some assurance of the robustness of the empirical findings.

5. Estimation Results and Hypothesis Tests

This section first presents estimation results for the marketable surplus equations. Given that these results are ancillary, we only briefly discuss them so as to devote the bulk of our analysis to the estimated matrix of price risk aversion and, more importantly, to the estimates of household willingness to pay to stabilize prices.

Table 5 presents estimation results for the seven marketable surplus equations. Intuitively, one would expect the ϕ_i (i.e., own-price) coefficients to be positive. That is,

[18] Alternatively, input prices are predetermined in the theoretical model of section 2.

as the price of commodity *i* increases, the household buys less or sells more of the same commodity. Indeed, own price has a positive and statistically significant effect on the marketable surplus of all commodities except wheat, for which the point estimate is statistically insignificantly different from zero. The results that are consistent indicate that some goods are substitutes for one another (e.g., coffee and barley; maize and sorghum; beans and sorghum; and wheat and sorghum) while others are complements (e.g., coffee and wheat; teff and coffee; beans and barley; barley and teff; wheat and teff).

5.1 Price Risk Aversion Matrix

We use the results of table 5 to compute coefficients of own- and cross-price risk aversion and use these coefficients to construct sub-matrices A_3 to A_7 of price risk aversion.[19] The ERHS households are significantly own-price risk-averse over all commodities (Table 6a). In addition, the average household is most significantly own-price risk-averse over barley, maize and teff – the staples for which net buyers' net purchase volumes are greatest (table 2) – and least price risk-averse over coffee and beans.[20]

The statistical significance and magnitude of the off-diagonal elements of the estimated A matrix underscore the importance of estimating price risk aversion in a multivariate context. All 42 off-diagonal point estimates are statistically significantly

[19] 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.

[20] The coefficients in table 6a are directly comparable between one another given that the marketable surpluses are all expressed in kilograms, and prices are all expressed in Ethiopian birr. Should either measurement unit differ between commodities, these coefficients would no longer be comparable.

different from zero, all of them positive, indicating aversion to positive cofluctuations in commodity prices. Single price approaches to estimating price risk aversion would neglect these effects, leading to biased estimates of own-price risk aversion. In particular, they would fail to capture how they routinely dislike covariation in multiple prices that limits households' capacity to substitute among crops in response to price shocks.

Recall that the standard theory employed in section 2 implies symmetry of the A matrix. We reject the null hypothesis of symmetry for sub-matrices A_3 to A_7, as shown in table 6b under the assumption that $R = 2$. This result is robust to alternative assumptions about the coefficient of income risk aversion R (see Appendix A). Rejecting the hypothesis of symmetry of the matrix of price risk aversion is equivalent to rejecting the symmetry of the Slutsky matrix, as per proposition 1 above. Deaton and Muellbauer (1980) and Browning and Chiappori (1998), among others, likewise rejected the null hypothesis of Slutsky symmetry.

Our results so far thus indicate that the households in our data (i) are averse to fluctuations in the prices of specific commodities; (ii) are risk-averse to cofluctuations in the prices of specific pairs of commodities; but (iii) do not behave in a manner fully consistent with the canonical model that implies Slutsky and price risk aversion symmetry. We now turn to how price fluctuations affect welfare.

5.2 Willingness to Pay Estimates for Price Stabilization

Recall from section 2.4 that the WTP for stabilization of a single commodity price can be estimated by considering either the rows or columns of matrix A of price risk aversion, but that both values coincide by construction for total WTP. For our three relative risk aversion assumptions (i.e., $R \in \{1,2,3\}$), tables 7a and 7b show the estimated average household WTP, expressed as a proportion of household income, to stabilize the prices of individual commodities as well as to stabilize the prices of all seven commodities considered in this paper. In what follows, we only discuss the results for $R = 2$, but the interpretation of the results for $R = 1$ or $R = 3$ is straightforward.

Estimating WTP with the rows of A in table 7a, the average WTP estimates are all statistically significantly different from zero. The commodity for which the average household would be willing to pay the highest proportion of its budget to stabilize the price is coffee (14.2 percent). Barley and teff come far behind as second and third (1.6 percent), and wheat is fourth (1 percent). Alternatively, estimating WTP with the columns of A in table 7b, the average WTP estimates are all statistically significant, but while coffee remains the commodity for which the average household would be willing to pay the highest proportion of its budget to stabilize the price (13.4 percent), sorghum and teff now come second and third (1.9 and 1.8 percent), and barley comes fourth (1 percent).

The average household's WTP estimate to stabilize the prices of these seven commodities is about 19 percent of its income (6 percent for $R = 1$; 32 percent for $R = 3$), a proportion that is statistically significant at the one percent level. By way of comparison, we compute the WTP measures derived by Finkelshtain and Chalfant (1997) in the case of a single stochastic commodity price, ignoring the covariances between prices (Table 7c). We reject the null hypothesis that either of our total WTP measures equals the analog measure ignoring the covariance between prices, with a *p*-value of 0.00. In these data, covariances between prices matter; ignoring them underestimates the average welfare gain by more than one-fifth, dropping estimated WTP to stabilize the prices of the seven most important commodities from 19 percent of income to 15 percent.

In order to be more specific about the distribution of the welfare gains from price stabilization, figure 1 plots the results of a fractional polynomial regression of the household-specific WTP to stabilize the prices of all seven commodities on household income, along with the associated 95 percent confidence band. Three different features immediately jump out from this regression. First, a significant share (31%) of households are price risk-loving (i.e., the households whose WTP for price stabilization is statistically significantly negative) while a somewhat larger share (39%) are price risk-averse (i.e., the households whose WTP for price stabilization is statistically significantly positive). Thus the population is roughly equally divided among those who favor, oppose or are indifferent about price stabilization.

Second, the significantly price risk-loving households are markedly poorer than the significantly price risk-averse ones. Table 8 shows the income percentile ranges for which households are, on average, price risk-loving, price risk-neutral, and price risk-averse. Households in the top 39 percent of the income distribution (i.e., the households whose seasonal income lies between 442 and 10,000 birr) are expected to gain from price stabilization, while the poorest 62 percent of the income distribution lose out from price stabilization, on average. This suggests that price stabilization would be a regressive policy in Ethiopia, benefiting the better off at the expense of poorer households. Turnovsky (1978) discussed various theoretical predictions regarding the winners and losers from price stabilization between consumers and producers. His results depended on whether price uncertainty stems from random fluctuations in supply or in demand; on whether price uncertainty is additive or multiplicative; and on whether supply and demand functions are linear. Our approach is free from such assumptions and lets the data speak for themselves.

Third, the magnitude of price risk preferences is far higher among the price risk-averse than among the price risk-loving, hence the sizable average WTP for price risk stabilization although the population is roughly evenly divided among the price risk-averse, price risk-loving and price risk-neutral subpopulations. Given the generally greater political influence of wealthier subpopulations in determining food price policy (Lipton 1977, Bates 1981) and the greater incentives for political mobilization among subgroups with a larger personal stake in the outcome (Olson 1965), together these three observations may help partly explain some of the political economy of food price stabilization in spite of heterogeneous preference for food price stability. These observations correspond relatively well with the "developmental paradox", i.e., the empirical regularity that the more developed the country, the more its government subsidizes agriculture and favors stabilizing crop prices (Lindert, 1991; Barrett, 1999).

5.3. *Ex Ante* Changes in Social Welfare under Three Policy Scenarios

As is well known, pure price stabilization through price fixing regulations or buffer stock management would introduce considerable distortions in the economy. Therefore, in this subsection we briefly consider a price risk compensation scheme, i.e., a policy that fully compensates those households who incur a welfare loss from price fluctuations but offers nothing to households who gain from price fluctuations. Although we have just shown that WTP for price risk reduction is greatest among those in the upper half of the income distribution, thus such a policy would be explicitly regressive, it may merit consideration as an alternative to full-blown price stabilization where political pressures, perhaps from economic elites, effectively compel the state to act in some fashion to reduce price fluctuations.

We begin by considering full price stabilization, i.e., a policy in which the households who gain from price fluctuations are fully taxed for their gains and in which the households who lose out from price fluctuations are fully compensated for their losses. This represents the naïve benchmark of pure price stabilization without general equilibrium effects.

Table 9a characterizes the in-sample winners and losers from such a policy. Under an assumed relative risk aversion $R = 2$, in this Ethiopian sample, those who would lose out from price stabilization vastly outnumber those who would gain (5216 versus 3060 households). But those who would lose out would incur a welfare loss from price stabilization that is on average much smaller than magnitude than the welfare gain of those who would benefit from nonstochastic prices (53 birr versus 660 birr). This echoes the point made in the previous section.

Table 9b then characterizes the social welfare changes for three policy options. The first is a *laissez-faire* policy (column 5) where nothing is done about commodity price volatility. The second is the pure price stabilization policy discussed above (column 6). The last column reflects a price risk compensation scheme in which the households who are price risk-averse are compensated for their exposure to price fluctuations but in which the households who are price risk-neutral and price risk-loving are unaffected. Under *laissez-faire*, the change in social welfare is equal to zero by construction. The change in social welfare is highest under a price risk compensation scheme, with the pure price stabilization policy falling between *laissez-faire* and price risk compensation.

Ignoring fiscal costs and general equilibrium effects, both of which cannot be quantified with the data at hand, it thus appears that only price risk compensation is Pareto-improving given that it improves the welfare of the subpopulation of price risk-averse households while leaving price risk-neutral and price risk-loving households unaffected. By contrast, pure price stabilization would make a majority of households worse off, even though average welfare gains are positive because the average gains to the price risk-averse subpopulation are an order of magnitude greater than the average losses to the price risk-loving subpopulation.

6. Conclusion

This paper has modestly extended microeconomic theory so as to enable the empirical study of price risk aversion over multiple commodities. Specifically, we first derived a matrix measuring the curvature of the indirect utility function in the hyperspace defined by the prices faced by agricultural households. The elements of this matrix describe own- and cross-price risk aversion. We also show how testing for the symmetry of the matrix of price risk aversion coefficients is equivalent to, but imposes less structure than, testing the symmetry of the Slutsky matrix. In the empirical portion of the paper, consistent with a vast literature that tests the symmetry of estimated Slutsky matrices, we reject the hypothesis that household behavior manifests this symmetry in the matrix of price risk aversion.

We estimated this matrix of price risk aversion coefficients using well-known survey data on a panel of rural Ethiopian households. We find these households are, on average, significantly price risk-averse over the prices of specific commodities and over cofluctuations in the prices of the same commodities, with an average willingness to pay to fully stabilize prices of 6 to 32 percent, depending on one's assumption about Arrow-Pratt relative income risk aversion. Hence governments' interest in price stabilization; on average, households stand to benefit from it. Nonparametric analysis of household-specific WTP estimates, however, suggests that the welfare gains from stabilizing prices at their means would accrue to households in the upper half of the income distribution and that a significant proportion of the households in the bottom half of the income distribution would actually be hurt by price stabilization, suggesting regressive benefit incidence from price stabilization policy as well as an explanation for the “developmental paradox. Finally, when the political economy of price stabilization effectively compels some government activity, we suggest a price risk compensation alternative to outright price stabilization and demonstrate that this might prove a Pareto superior, albeit distributionally regressive, policy response to the welfare costs associated with commodity price volatility.

References

Baron, D.P. (1970), "Price Uncertainty, Utility, and Industry Equilibrium in Pure Competition," *International Economic Review* 11(3): 463-480.

Barrett, C.B. (1996), "On Price Risk and the Inverse Farm Size–Productivity Relationship," *Journal of Development Economics* 51(2): 193-215.

Barrett, C.B. (1999), "The Microeconomics of the Developmental Paradox: On the Political Economy of Food Price Policy," *Agricultural Economics* 20(2): 159-172.

Bates, R. (1981), *Markets and States in Tropical Africa,* Berkeley: University of California Press.

Bellemare, M.F., and C.B. Barrett (2006), "An Ordered Tobit Model of Market Participation: Evidence from Kenya and Ethiopia," *American Journal of Agricultural Economics* 88(2): 324-337.

Browning, M., and P.-A. Chiappori (1998), "Efficient Intrahousehold Allocations: A General Characterization and Empirical Tests," *Econometrica* 66(6): 1241-1278.

Chavas, J.-P., and M.T. Holt (1993), "Acreage Decisions under Risk: The Case of Corn and Soybeans," *American Journal of Agricultural Economics* 72(3): 529-538.

Deaton, A., and J. Muellbauer (1980), "An Almost Ideal Demand System," *American Economic Review*, 70(3): 312-326.

de Janvry, A., M. Fafchamps, and É. Sadoulet (1991), "Peasant Household Behavior with Missing Markets: Some Paradoxes Explained," *Economic Journal* 101(409): 1400-1417.

Dercon, S., and P. Krishnan (1998), "Changes in Poverty in Rural Ethiopia 1989-1995: Measurement, Robustness Tests and Decomposition," Working Paper, Centre for the Study of African Economies, Oxford University.

Deschamps, R. (1973), "Risk Aversion and Demand Functions," *Econometrica* 41(3): 455-465.

Epstein, L. (1975), "A Disaggregate Analysis of Consumer Choice under Uncertainty," *Econometrica* 43(5-6): 877-892.

Finkelshtain, I. and J. Chalfant (1991), "Marketed Surplus under Risk: Do Peasants Agree with Sandmo?," *American Journal of Agricultural Economics* 73(3): 557-567.

Finkelshtain, I. and J. Chalfant (1997), "Commodity Price Stabilization in a Peasant Economy," *American Journal of Agricultural Economics* 79(4): 1208-1217.

Friend, I. and M.E. Blume (1975), "The Demand for Risky Assets," *American Economic Review* 65(5): 900-922.

Goetz, S.J. (1992) "A Selectivity Model of Household Food Marketing Behavior in Sub-Saharan Africa," *American Journal of Agricultural Economics* 74(2): 444-452.

Hanoch, G. (1977), "Risk Aversion and Consumer Preferences," *Econometrica* 45(2): 413-426.

Hansen, L.P. and K.J. Singleton (1982), "Stochastic Consumption, Risk Aversion, and the Temporal Behavior of Asset Returns," *Journal of Political Economy* 91(2): 249-265.

Hirshleifer, J. and J.G. Riley (1992), *The Analytics of Uncertainty and Information*, Cambridge: Cambridge University Press.

Knudsen, O. and J. Nash (1990), "Domestic Price Stabilization Schemes in Developing Countries," *Economic Development and Cultural Change* 38(3): 539-558.

Krueger, A.O., M. Schiff and A. Valdes (1988), "Agricultural Incentives in Developing Countries: Measuring the Effect of Sectoral and Economywide Policies," *World Bank Economic Review* 2(3): 255-271.

Lindert, P.H. (1991), "Historical Patterns of Agricultural Policy," in C.P. Timmer, ed., *Agriculture and the State*, Ithaca: Cornell University Press.

Lipton, M. (1977), *Why poor people stay poor: urban bias in world development* Cambridge, MA: Harvard University Press.

MaCurdy, T.E., and J.H. Pencavel (1986), "Testing between Competing Models of Wage and Employment Determination in Unionized Markets," *Journal of Political Economy* 94(3): S3-S39.

Mas-Colell, A., M.D. Whinston, and J.R. Green (1995), *Microeconomic Theory*, Oxford, UK: Oxford University Press.

Newbery, D.M.G. (1989), "The Theory of Food Price Stabilization," *Economic Journal* 99(398): 1065-1082.

Newbery, D.M.G, and J.E. Stiglitz (1981), *The Theory of Commodity Price Stabilization*, Oxford, UK: Oxford University Press.

Olson, Mancur (1965), *The Logic of Collective Action: Public Goods and the Theory of Groups*, Cambridge, MA: Harvard University Press.

Pratt, J.W. (1964), "Risk Aversion in the Small and in the Large," *Econometrica* 32(1-2): 122-136.

Saha, A., C.R. Shumway, and H. Talpaz (1994), "Joint Estimation of Risk Preference Structure and Technology Using Expo-Power Utility," *American Journal of Agricultural Economics*, 76(2): 173-184.

Sandmo, A. (1971), "On the Theory of the Competitive Firm under Price Uncertainty," *American Economic Review* 61(1): 65-73.

Sherlund, S., C.B. Barrett, and A.A. Adesina (2002), "Smallholder Technical Efficiency Controlling for Environmental Production Conditions," *Journal of Development Economics*, 69(1): 85-101.

Singh, I., L. Squire, and J. Strauss (1986), *Agricultural Household Models*, Baltimore, MD: Johns Hopkins University Press.

Timmer, C.P. (1989), *Food Price Stabilization: Rationale, Design, and Implementation.* Cambridge, MA: Harvard Institute for International Development.

Turnovsky, S.J. (1978), "The Distribution of Welfare Gains from Price Stabilization: A Survey of Some Theoretical Issues," in F.G. Adams and S.A. Klein, eds., *Stabilizing World Commodity Markets*, Lexington, MA: Lexington Books.

Turnovsky, S.J., H. Shalit, and A. Schmitz (1980), "Consumer's Surplus, Price Instability, and Consumer Welfare," *Econometrica* 48(1): 135-152.

Verbeek, M., and T. Nijman (1992), "Testing for Selectivity Bias in Panel Data Models," *International Economic Review* 33(3): 681-703.

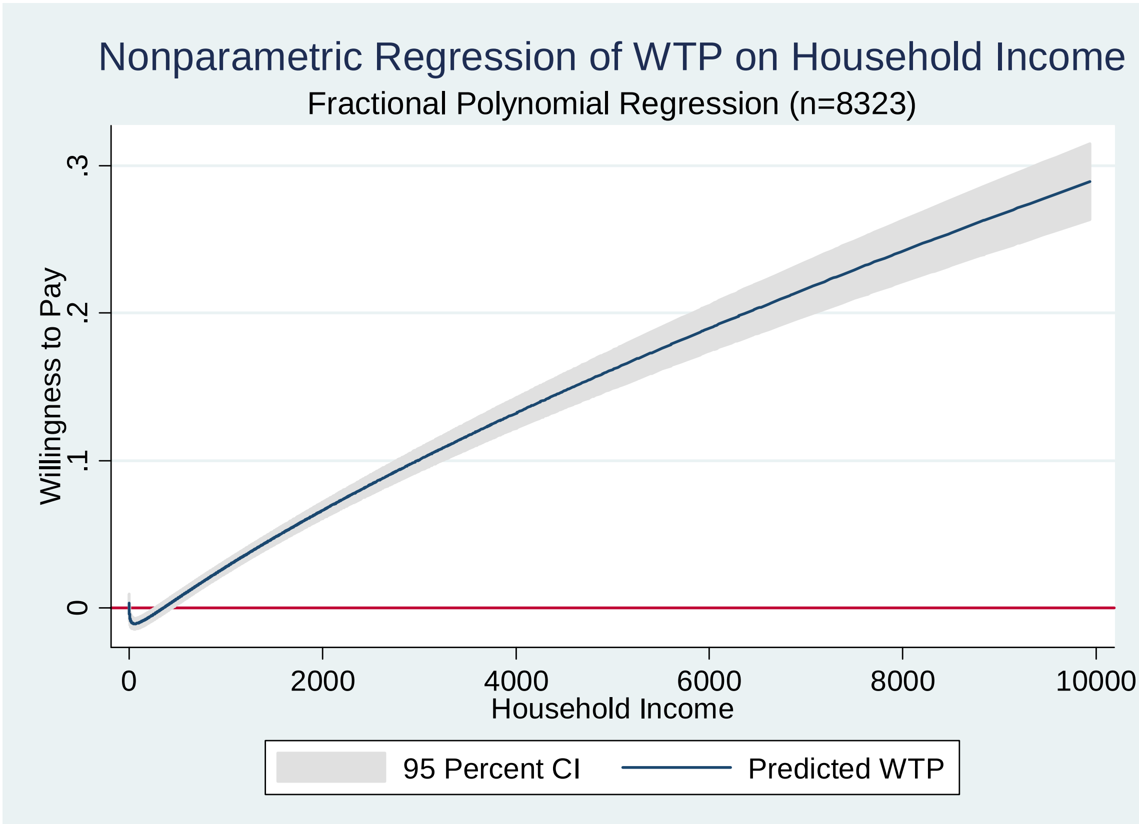

Figure 1: Fractional polynomial regression of household WTP to eliminate price fluctuations among seven staple commodities on household income for households whose seasonal income does not exceed 10,000 birr.

Table 1: Seasonal Descriptive Statistics for the Dependent Variables (Full Sample)

Crop	Mean	(Std. Dev.)	Observations	Nonzero Observations
Coffee Marketable Surplus (Kg)	-13.36	(87.37)	8556	6744
Maize Marketable Surplus (Kg)	-121.57	(364.54)	8556	3966
Beans Marketable Surplus (Kg)	-40.39	(95.63)	8556	3030
Barley Marketable Surplus (Kg)	-88.76	(367.04)	8556	2825
Wheat Marketable Surplus (Kg)	-64.82	(279.28)	8556	2796
Teff Marketable Surplus (Kg)	-100.92	(335.37)	8556	2666
Sorghum Marketable Surplus (Kg)	-38.82	(204.00)	8556	1712

Table 2: Seasonal Descriptive Statistics for the Dependent Variables (Nonzero Observations)

Crop	Net Buyer Mean Marketable Surplus	(Std. Dev.)	Net Buyer Observations	Net Seller Mean Marketable Surplus	(Std. Dev.)	Net Seller Observations
Coffee (Kg)	-23.44	(95.64)	6206	57.92	(95.02)	538
Maize (Kg)	-397.18	(438.32)	3115	231.55	(388.10)	851
Beans (Kg)	-127.14	(122.91)	2848	90.70	(95.32)	182
Barley (Kg)	-459.27	(553.31)	2097	279.81	(329.47)	728
Wheat (Kg)	-296.70	(337.00)	2420	434.74	(620.52)	376
Teff (Kg)	-471.03	(453.10)	2136	269.06	(432.08)	530
Sorghum (Kg)	-349.56	(320.29)	1313	317.96	(290.27)	399

Table 3: Seasonal Descriptive Statistics for the Independent Variables (n=8556)

Crop	**Mean**	**(Std. Dev.)**
Commodity Real Prices		
Coffee (Birr/Kg)	13.32	(5.20)
Maize (Birr/Kg)	1.29	(0.38)
Beans (Birr/Kg)	1.88	(0.43)
Barley (Birr/Kg)	1.50	(0.41)
Wheat (Birr/Kg)	1.74	(0.33)
Teff (Birr/Kg)	2.28	(0.40)
Sorghum (Birr/Kg)	1.52	(0.42)
Potatoes (Birr/Kg)	1.52	(0.74)
Onions (Birr/Kg)	1.97	(0.78)
Cabbage (Birr/Kg)	0.92	(0.68)
Milk (Birr/Liter)	2.09	(0.88)
Tella (Birr/Liter)	0.69	(0.25)
Sugar (Birr/Kg)	5.85	(2.08)
Salt (Birr/Kg)	1.70	(1.02)
Cooking Oil (Birr/Liter)	9.14	(2.60)
Income		
Income (Birr)	886.17	(9869.70)
Nonzero Income (Birr)	1087.35	(10922.88)
Budget Shares		
Budget Share of Coffee	-0.15	(1.05)
Budget Share of Maize	-0.13	(0.40)
Budget Share of Beans	-0.07	(0.16)
Budget Share of Barley	-0.12	(0.52)
Budget Share of Wheat	-0.11	(0.43)
Budget Share of Teff	-0.21	(0.69)
Budget Share of Sorghum	-0.06	(0.33)

Note: Income (i.e., the sum of off-farm income, all crop revenues, and livestock sales) was different from zero for only 6973 observations, so budget shares are computed for that sub-sample. Because of the presence of nonzero incomes, budget shares were obtained by dividing marketable surpluses by mean nonzero income.

Table 4: Seasonal Variance-Covariance Matrix of Commodity Prices

	Coffee	Maize	Beans	Barley	Wheat	Teff	Sorghum
Coffee	**27.05**						
Maize	0.46	**0.15**					
Beans	0.25	0.05	**0.19**				
Barley	0.29	0.03	-0.04	**0.17**			
Wheat	0.13	0.04	0.05	0.05	**0.11**		
Teff	0.00	0.06	0.06	0.06	0.06	**0.16**	
Sorghum	0.18	0.05	0.00	0.06	0.06	0.03	**0.17**

Table 5: Seasonal Marketable Surplus Equations

	(1)		(2)		(3)		(4)	
Dependent Variable:	**Coffee Marketable Surplus**		**Maize Marketable Surplus**		**Beans Marketable Surplus**		**Barley Marketable Surplus**	
Coefficients	**Coefficient**	**(Std. Err.)**	**Coefficient**	**(Std. Err.)**	**Coefficient**	**(Std. Err.)**	**Coefficient**	**(Std. Err.)**
Coffee Price	**40.273*** **	**(0.092)**	96.297***	(0.688)	15.471***	(0.051)	-178.545***	(0.500)
Maize Price	-6.344	(3.703)	**389.529*** **	**(27.526)**	36.811***	(2.043)	118.277***	(20.027)
Beans Price	-9.567*	(5.359)	-115.378**	(39.841)	**39.952*** **	**(2.957)**	257.049***	(28.988)
Barley Price	-27.499***	(3.585)	-53.697*	(26.648)	7.396***	(1.978)	**305.961*** **	**(19.389)**
Wheat Price	30.689***	(5.884)	137.883***	(43.738)	146.613***	(3.246)	254.879***	(31.823)
Teff Price	104.537***	(7.905)	-66.515	(58.761)	-94.449***	(4.361)	-326.964***	(42.754)
Sorghum Price	-68.434***	(2.637)	-73.385***	(19.603)	-95.320***	(1.455)	-445.863***	(14.263)
Potatoes Price	12.659***	(1.007)	7.845	(7.488)	26.723***	(0.556)	37.081***	(5.448)
Onions Price	-24.624***	(3.258)	59.407**	(24.220)	-51.082***	(1.798)	-275.690***	(17.622)
Cabbage Price	-10.344***	(0.563)	52.844***	(4.182)	5.462***	(0.310)	73.095***	(3.043)
Milk Price	-13.161***	(1.182)	290.977***	(8.790)	-26.748***	(0.652)	18.605**	(6.395)
Tella Price	28.556***	(4.217)	131.295***	(31.345)	75.795***	(2.326)	307.160***	(22.806)
Sugar Price	11.445***	(3.310)	-151.995***	(24.602)	5.659***	(1.826)	21.166	(17.901)
Salt Price	6.754***	(1.944)	121.660***	(14.449)	-36.798***	(1.072)	-264.330***	(10.513)
Cooking Oil Price	-5.634**	(2.350)	-15.989	(17.470)	-91.614***	(1.297)	-362.422***	(12.711)
Income	0.721	(0.499)	7.242*	(3.708)	0.302	(0.275)	6.861**	(2.698)
Intercept	-151.799***	(15.180)	-304.647**	(112.841)	187.358***	(8.374)	1448.533***	(82.102)
N	8556		8556		8556		8556	
p-value (All Coefficients)	0.00		0.00		0.00		0.00	
R^2	0.33		0.44		0.44		0.37	
Household FEs	Yes		Yes		Yes		Yes	
Woreda-Round FEs	Yes		Yes		Yes		Yes	

Note: *, **, and *** denote significance at the 90, 95, and 99 percent levels. Bolded coefficients and standard errors are for own-price effects.

Table 5 (continued): Seasonal Marketable Surplus Equations

	(5)		(6)		(7)	
Dependent Variable:	**Wheat Marketable Surplus**		**Teff Marketable Surplus**		**Sorghum Marketable Surplus**	
Variable	**Coefficient**	**(Std. Err.)**	**Coefficient**	**(Std. Err.)**	**Coefficient**	**(Std. Err.)**
Coffee Price	2.688***	(0.290)	113.406***	(0.632)	22.879***	(0.369)
Maize Price	-85.971***	(11.603)	48.316*	(25.306)	-36.271**	(14.774)
Beans Price	-35.692**	(16.795)	-63.054	(36.628)	-68.184***	(21.383)
Barley Price	-46.119***	(11.234)	-58.085**	(24.499)	50.331***	(14.303)
Wheat Price	**17.469**	**(18.438)**	38.188**	(40.211)	-144.397***	(23.475)
Teff Price	235.372***	(24.771)	**123.266****	**(54.022)**	84.194**	(31.538)
Sorghum Price	-45.547***	(8.263)	3.172	(18.022)	**39.693*****	**(10.521)**
Potatoes Price	31.551***	(3.157)	10.700	(6.884)	-34.192***	(4.019)
Onions Price	-64.140***	(10.210)	103.915***	(22.267)	61.417***	(12.999)
Cabbage Price	21.528***	(1.763)	29.668***	(3.845)	8.517***	(2.244)
Milk Price	-134.989***	(3.705)	111.745***	(8.081)	25.431***	(4.718)
Tella Price	106.587***	(13.213)	-0.722	(28.817)	-80.962***	(16.823)
Sugar Price	43.907***	(10.371)	-175.316***	(22.619)	-22.747	(13.205)
Salt Price	-3.396	(6.091)	140.199***	(13.284)	8.750	(7.755)
Cooking Oil Price	8.577	(7.364)	12.020	(16.061)	75.466***	(9.376)
Income	0.626	(1.563)	4.950	(3.409)	2.861	(1.990)
Intercept	-130.943**	(47.568)	-396.577***	(103.741)	-300.241***	(60.564)
N	8556		8556		8556	
p-value (All Coefficients)	0.00		0.00		0.02	
R^2	0.39		0.45		0.37	
Household FEs	Yes		Yes		Yes	
Woreda-Round FEs	Yes		Yes		Yes	

Note: *, **, and *** denote significance at the 90, 95, and 99 percent levels. Bolded coefficients and standard errors are for own-price effects.

Table 6a: Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 2

	Coffee	Maize	Beans	Barley	Wheat	Teff	Sorghum
Coffee	**18.148*** (5.229)**	10.091*** (1.983)	3.427*** (0.663)	17.293*** (2.758)	6.894*** (0.997)	11.056*** (1.879)	2.510*** (0.783)
Maize	10.063*** (1.978)	**620.421*** (72.300)**	15.567*** (2.035)	58.287*** (12.732)	45.306*** (11.961)	134.083*** (24.918)	22.676*** (6.237)
Beans	3.507*** (0.679)	15.969*** (2.088)	**51.661*** (4.387)**	96.571*** (10.278)	42.830*** (6.383)	57.995*** (6.444)	10.821*** (1.952)
Barley	17.098*** (2.727)	57.788*** (12.624)	93.324*** (9.933)	**893.913*** (101.013)**	125.214*** (17.033)	112.650*** (19.065)	28.062*** (8.739)
Wheat	7.046*** (1.019)	46.433*** (12.258)	42.785*** (6.376)	129.431*** (17.606)	**275.618*** (60.152)**	136.169*** (26.187)	16.764*** (4.529)
Teff	11.083*** (1.883)	134.770*** (25.046)	56.819*** (6.314)	114.203*** (19.327)	133.552*** (25.684)	**514.857*** (58.887)**	28.438*** (4.839)
Sorghum	2.486*** (0.776)	22.521*** (6.194)	10.476*** (1.889)	28.108*** (8.754)	16.246*** (4.389)	28.099*** (4.781)	**94.009*** (13.201)**

Note: All coefficients are divided by 100,000. Standard errors are in parentheses, and *, **, and *** denote significance at the 90, 95, and 99 percent levels. Bolded coefficients are own-price risk aversion coefficients.

Table 6b: Tests of Symmetry of the Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 2

Sub-Matrix	Test Statistic	*p*-value
Symmetry of Sub-Matrix A_3 (Coffee, …, Beans)	$F(3, 8553) = 24.48$	0.00
Symmetry of Sub-Matrix A_4 (Coffee, …, Barley)	$F(6, 8550) = 26.53$	0.00
Symmetry of Sub-Matrix A_5 (Coffee, …, Wheat)	$F(9, 8547) = 22.33$	0.00
Symmetry of Sub-Matrix A_6 (Coffee, …, Teff)	$F(14, 8542) = 16.59$	0.00
Symmetry of Sub-Matrix A_7 (Coffee, …, Sorghum)	$F(27, 8529) = 13.61$	0.00
Joint Significance (All Coefficients)	$F(29, 8527) = 18.62$	0.00
Joint Significance (Diagonal Coefficients)	$F(7, 8549) = 47.52$	0.00
Joint Significance (Off-Diagonal Coefficients)	$F(22, 8534) = 22.85$	0.00

Note: Constraints were dropped due to collinearity in every test.

Table 7a: WTP as Proportion of Household Income (Rows)

	R = 1		*R* = 2		*R* = 3	
Commodity	**WTP**	**(Std. Err)**	**WTP**	**(Std. Err)**	**WTP**	**(Std. Err)**
Coffee	0.052***	(0.019)	0.142***	(0.039)	0.231***	(0.060)
Maize	-0.013***	(0.001)	-0.003***	(0.001)	0.007***	(0.001)
Beans	-0.002***	(0.000)	-0.001***	(0.000)	0.000	(0.000)
Barley	0.005***	(0.001)	0.016***	(0.001)	0.027***	(0.002)
Wheat	0.005***	(0.000)	0.010***	(0.001)	0.016***	(0.001)
Teff	0.007***	(0.000)	0.016***	(0.001)	0.026***	(0.001)
Sorghum	0.003***	(0.000)	0.007***	(0.000)	0.010***	(0.001)
All Commodities	0.056***	(0.019)	0.187***	(0.040)	0.318***	(0.060)

Note: Standard errors are in parentheses, and *, **, and *** denote significance at the 90, 95, and 99 percent levels.

Table 7b: WTP as Proportion of Household Income (Columns)

	R = 1		*R* = 2		*R* = 3	
Commodity	***R* = 1**	**(Std. Err)**	***R* = 2**	**(Std. Err)**	***R* = 3**	**(Std. Err)**
Coffee	0.045**	(0.019)	0.134***	(0.039)	0.224***	(0.059)
Maize	-0.014***	(0.001)	-0.004***	(0.001)	0.007***	(0.001)
Beans	0.005***	(0.000)	0.006***	(0.000)	0.007***	(0.000)
Barley	-0.002***	(0.001)	0.010***	(0.001)	0.021***	(0.002)
Wheat	-0.001***	(0.000)	0.004***	(0.001)	0.009***	(0.001)
Teff	0.008***	(0.000)	0.018***	(0.001)	0.028***	(0.001)
Sorghum	0.015***	(0.000)	0.019***	(0.001)	0.022***	(0.001)
All Commodities	0.056***	(0.019)	0.187***	(0.040)	0.318***	(0.060)

Note: Standard errors are in parentheses, and *, **, and *** denote significance at the 90, 95, and 99 percent levels.

Table 7c: WTP as Proportion of Household Income Ignoring Covariances

	R = 1		*R* = 2		*R* = 3	
Commodity	***R* = 1**	**(Std. Err)**	***R* = 2**	**(Std. Err)**	***R* = 3**	**(Std. Err)**
Coffee	0.045**	(0.019)	0.133***	(0.039)	0.222***	(0.059)
Maize	-0.015***	(0.001)	-0.006***	(0.001)	0.003**	(0.001)
Beans	-0.001***	(0.000)	0.000***	(0.000)	0.001***	(0.000)
Barley	-0.007***	(0.001)	0.003***	(0.001)	0.013***	(0.001)
Wheat	0.003***	(0.000)	0.007***	(0.001)	0.011***	(0.001)
Teff	0.005***	(0.000)	0.013***	(0.001)	0.021***	(0.001)
Sorghum	0.001***	(0.000)	0.004***	(0.000)	0.008***	(0.000)
All Commodities	0.030	(0.019)	0.154***	(0.039)	0.278***	(0.059)

Note: These measures are derived following Finkelshtain and Chalfant (1997). Standard errors are in parentheses, and *, **, and *** denote significance at the 90, 95, and 99 percent levels.

Table 8: WTP for Price Risk Stabilization Across the Income Distribution

Income Range (Birr)	Income Percentile Range	Sign of Fitted WTP
0.00 - 3.24	0.00 - 18.84	0
3.24 - 267.66	18.84 - 49.82	-
267.66 - 441.92	49.82 - 61.49	0
441.92 - 10,000.00	61.49 - 100.00	+

Note: These numbers reflect the regression plotted in figure 1. A negative sign in the third column means that households in this interval are statistically significantly price risk-loving; a 0 means that households in this interval have no statistically significant preference for or against price variability; and a positive sign means that households in this interval are statistically significantly price risk-averse.

Table 9a: Welfare Gains and Losses from Eliminating Price Fluctuations

Coefficient of Relative Risk Aversion	(1) Average Welfare Gain (Birr)	(2) Number of Households Who Would Benefit	(3) Average Welfare Loss (Birr)	(4) Number of Households Who Would Lose Out
R = 1	439.37	2134	67.23	6142
R = 2	660.41	3060	52.96	5216
R = 3	836.28	3778	43.35	4498

Note: The average welfare gains and losses are derived from the "All Commodities" estimates in tables 7a and 7b and reflect the effect on household welfare of completely eliminating price fluctuations, i.e., keeping the prices of coffee, maize, beans, barley, wheat, teff, and sorghum fixed at their means.

Table 9b: *Ex Ante* Marginal Changes in Social Welfare under Three Policy Scenarios

Coefficient of Relative Risk Aversion	(5) Change in Social Welfare under *Laissez-Faire*	(6) Change in Social Welfare under Price Stabilization (1) x (2) - (3) x (4)	(7) Change in Social Welfare under Compensation (1) x (2)
R = 1	0	524,689	937,616
R = 2	0	1,744,615	2,020,855
R = 3	0	2,964,478	3,159,466

Note: Values in columns 5 to 7 are expressed in Ethiopian birr. Column 5 describes the change in social welfare under no policy. Column 6 describes the change in social welfare under a price stabilization policy, i.e., a policy in which prices are kept equal to their means and do not fluctuate, i.e., the product of columns 1 and 2 minus the product of columns 3 and 4. Column 7 describes the change in social welfare under a price risk compensation policy, i.e., a policy in which prices fluctuate but transfers are made to compensate those who suffer from price risk, i.e., the product of columns 1 and 2.

Appendix A

Table A1a: Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 1

	Coffee	Maize	Beans	Barley	Wheat	Teff	Sorghum
Coffee	**4.657*** (1.342)**	2.596*** (0.510)	0.860*** (0.166)	4.484*** (0.715)	1.732*** (0.250)	2.830*** (0.481)	0.650*** (0.203)
Maize	2.582*** (0.507)	**159.592*** (18.598)**	3.907*** (0.511)	15.115*** (3.302)	11.381*** (3.005)	34.324*** (6.379)	5.871*** (1.615)
Beans	0.900*** (0.174)	4.108*** (0.537)	**12.964*** (1.101)**	25.043*** (2.665)	10.759*** (1.603)	14.847*** (1.650)	2.802*** (0.505)
Barley	4.387*** (0700)	14.865*** (3.247)	23.418*** (2.492)	**231.808*** (26.193)**	31.455*** (4.279)	28.840*** (4.880)	7.267*** (2.263)
Wheat	1.808*** (0.261)	11.945*** (3.153)	10.737*** (1.600)	33.564*** (4.565)	**69.239*** (15.110)**	34.858*** (6.703)	4.340*** (1.173)
Teff	2.843*** (0.483)	34.667*** (6.443)	14.259*** (1.584)	29.616*** (5.012)	33.549*** (6.452)	**131.799*** (15.074)**	7.362*** (1.253)
Sorghum	0.638*** (0.199)	5.794*** (1.593)	2.629*** (0.474)	7.288*** (2.270)	4.082*** (1.103)	7.193*** (1.224)	**24.339*** (3.418)**

Note: Standard errors are in parentheses, and *, **, and *** denote significance at the 90, 95, and 99 percent levels. Bolded coefficients are own-price risk aversion coefficients.

Table A1b: Tests of Symmetry of the Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 1

Sub-Matrix	Test Statistic	*p*-value
Symmetry of Sub-Matrix A_3 (Coffee, …, Beans)	$F(3, 8553) = 24.44$	0.00
Symmetry of Sub-Matrix A_4 (Coffee, …, Barley)	$F(6, 8550) = 26.55$	0.00
Symmetry of Sub-Matrix A_5 (Coffee, …, Wheat)	$F(9, 8547) = 22.24$	0.00
Symmetry of Sub-Matrix A_6 (Coffee, …, Teff)	$F(14, 8542) = 16.58$	0.00
Symmetry of Sub-Matrix A_7 (Coffee, …, Sorghum)	$F(20, 8536) = 13.65$	0.00
Joint Significance (All Coefficients)	$F(33, 8523) = 88.54$	0.00
Joint Significance (Diagonal Coefficients)	$F(7, 8549) = 47.52$	0.00
Joint Significance (Off-Diagonal Coefficients)	$F(26, 8530) = 110.23$	0.00

Note: Constraints were dropped due to collinearity in every test.

Table A2a: Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 3

	Coffee	Maize	Beans	Barley	Wheat	Teff	Sorghum
Coffee	**13.493*** (3.887)**	7.496*** (1.473)	2.567*** (0.497)	12.809*** (2.043)	5.162*** (0.746)	8.226*** (1.398)	1.860*** (0.581)
Maize	7.482*** (1.470)	**460.839*** (53.702)**	11.661*** (1.524)	43.174*** (9.431)	33.926*** (8.956)	99.761*** (18.539)	16.806*** (4.622)
Beans	2.607*** (0.505)	11.862*** (1.551)	**38.699*** (3.286)**	71.534*** (7.613)	32.072*** (4.779)	43.151*** (4.795)	8.020*** (1.446)
Barley	12.712*** (2.027)	42.924*** (9.376)	69.909*** (7.440)	**662.159*** (74.820)**	93.766*** (12.754)	83.822*** (14.184)	20.798*** (6.477)
Wheat	5.239*** (0.757)	34.490*** (9.105)	32.050*** (4.776)	95.875*** (13.041)	**206.392*** (45.042)**	101.315*** (19.484)	12.424*** (3.357)
Teff	8.240*** (1.400)	100.104*** (18.603)	42.563*** (4.729)	84.598*** (14.316)	100.007*** (19.232)	**383.074*** (43.813)**	21.077*** (3.586)
Sorghum	1.848*** (0.577)	16.729*** (4.601)	7.847*** (1.415)	20.819*** (6.484)	12.166*** (3.287)	20.907*** (3.557)	**69.675*** (9.783)**

Note: Standard errors are in parentheses, and *, **, and *** denote significance at the 90, 95, and 99 percent levels. Bolded coefficients are own-price risk aversion coefficients.

Table A2b: Tests of Symmetry of the Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 3

Sub-Matrix	**Test Statistic**	***p*-value**
Symmetry of Sub-Matrix A_3 (Coffee, …, Beans)	$F(3, 8553) = 24.44$	0.00
Symmetry of Sub-Matrix A_4 (Coffee, …, Barley)	$F(6, 8550) = 26.55$	0.00
Symmetry of Sub-Matrix A_5 (Coffee, …, Wheat)	$F(9, 8547) = 22.24$	0.00
Symmetry of Sub-Matrix A_6 (Coffee, …, Teff)	$F(14, 8542) = 16.58$	0.00
Symmetry of Sub-Matrix A_7 (Coffee, …, Sorghum)	$F(20, 8536) = 13.65$	0.00
Joint Significance (All Coefficients)	$F(29, 8527) = 18.61$	0.00
Joint Significance (Diagonal Coefficients)	$F(7, 8549) = 47.52$	0.00
Joint Significance (Off-Diagonal Coefficients)	$F(22, 8534) = 18.38$	0.00

Note: Constraints were dropped due to collinearity in every test.