The Welfare Impacts of Commodity Price Volatility: Evidence from Rural Ethiopia*

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Abstract

Many governments have tried to stabilize commodity prices based on the widespread belief that households in developing countries – especially poorer ones – value price stability, defined here as the lack of fluctuations around a mean price level. We derive a measure of multivariate price risk aversion as well as an associated measure of willingness to pay for price stabilization across multiple commodities. Using data from a panel of Ethiopian households, our estimates suggest that the average household would be willing to pay 6-32 percent of its income to eliminate volatility in the prices of the seven primary food commodities. Not everyone benefits from price stabilization, however. Contrary to conventional wisdom, the welfare gains from eliminating price volatility would be concentrated in the upper 40 percent of the income distribution, making food price stabilization a distributionally regressive policy in this context.

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1. Introduction

How does commodity price volatility affect the welfare of households in developing countries, where consumption smoothing is often difficult? When governments intervene to stabilize commodity prices, who gains and who loses? Throughout history and all over the world, governments have frequently set commodity price stability – defined here as the absence of price fluctuations around a mean price level – as an important goal of economic policy. Using a host of policy instruments, from buffer stocks to administrative pricing and from variable tariffs to marketing boards, the same governments have tried to stabilize prices. These efforts have typically met with limited success, and after a period of significant policy research on the topic (Newbery and Stiglitz 1981), price stabilization had largely fallen off the policy agenda by the early 1990s.

Since the mid-1990s, however, commodity prices have been on a rollercoaster ride (Cashin and McDermott, 2002; Jacks et al., 2009; Roache, 2010). Food price ten-year volatility reached its highest level in almost 30 years in December 2010 (FAO, 2010). Likewise, food prices have reached an all-time high since 1990 in December 2010 (Treanor, 2011). Food price volatility over the past decade or so, punctuated by the food crisis of 2007-2008 and the biggest one-month jump in wheat prices in more than three decades in summer 2010, has rekindled widespread popular interest in commodity price stabilization. Several governments (e.g., Ghana, India, and Kenya) have recently reintroduced food price stabilization schemes. A simple search finds more than five times as many articles on the topic in the media over the last five years as in the preceding five years.[4] Likewise, the 2008 edition of the World Bank's flagship *World Development Report* (World Bank, 2008, pp. 121-122), which focused on agriculture, discussed various policy options for price stabilization. More recently, the Food and Agriculture Organization of the United Nations dedicated a policy brief to price volatility in agricultural markets, cautioning that "there is emerging consensus that the global food system is becoming more vulnerable and susceptible to episodes of extreme price volatility" (FAO, 2010). Similarly, the International Fund for Agricultural Development's

[4] An August 13, 2010, LexisNexis search turned up just 51 articles, 2000-2005, on "commodity price stabilization," or variants replacing "commodity" with "food" or "stabilization" with "stability" or "volatility," but 266 articles on the same search terms over the (slightly shorter) 2006-10 period.

2011 *Rural Poverty Report* (IFAD, 2011, p. 97) takes note of policy makers' "growing interest in the role public policy can play in (…) stabilizing food markets."

The impulse toward state interventions to stabilize domestic food prices commonly arises because (i) households are widely believed to value price stability; (ii) the poor are widely perceived to suffer disproportionately from food price instability; and (iii) futures and options markets for hedging against food price risk are commonly inaccessible to consumers and poor producers in developing countries (Newbery, 1989; Timmer, 1989). Although few experts would dispute claim (iii) above, convincing empirical tests of claims (i) and (ii) are noticeably absent from the published literature. Given the policy importance of the topic, however, and given that 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), it is puzzling that our theoretical and empirical toolkits for understanding the relationship between price volatility – what we will also refer to as "price uncertainty" or "price risk" in this paper – and household welfare are dated and have only rarely been applied empirically.

In this paper, we address that important gap in the literature by studying whether (i) households value price stability; and (ii) the poor suffer disproportionately from food price instability. There are ultimately empirical questions requiring household data and a clear strategy for relating a measure of household welfare to a measure of food price volatility. A regression of household welfare on food price variance is infeasible for several reasons.[5] We therefore tap the established theoretical literature on price risk to derive an estimable measure of multi-commodity price risk aversion and the associated willingness to pay for price stabilization. This lets us hold mean prices constant to focus on the welfare effects of food price variances and covariances.[6] As Sarris et al. (2011,

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[5] For an example of one of those reasons, note that there is no unique "food price." Although there exist food price indices (e.g., the FAO's Food Price Index), any index necessarily aggregates prices and suppresses variance using an arbitrary weighting scheme that almost surely does not match that of the households under study.

[6] Indeed, the issue of commodity price volatility is often inextricably linked with that of rising commodity prices in the public's mind. As discussed above, this paper sets aside the issue of rising commodity prices (i.e., increases in the mean of the distribution of commodity prices) to focus on the volatility of commodity prices (i.e., the variance of the distribution of commodity prices). Economists have had a good

p.48) note in their investigation of potential policy responses to food price volatility in low-income countries, “the main problem is not price or quantity variations *per se*, but rather unforeseen and undesirable departures from expectations” regarding commodity prices.

The effects of price volatility on producer behavior and profit have been well-explored in the theoretical literature. Output price uncertainty generally causes firms to employ fewer inputs, forgoing expected profits in order to hedge against price volatility (Baron, 1970; Sandmo, 1971). The analysis of commodity price risk has been extended theoretically to individual consumers (Deschamps, 1973; Hanoch, 1977; Turnovsky et al., 1980; Newbery and Stiglitz, 1981; Newbery, 1989), who are generally thought to be price risk-loving given the quasiconvexity of the indirect utility function. But because agricultural households can be both producers and consumers of the same commodities, it is entirely possible for some households to be price risk-averse, for others to be price risk-neutral, and for yet others to be price risk-loving, although prior empirical analyses have focused on just a single commodity (Finkelshtain and Chalfant 1991, 1997; Barrett, 1996). And while Turnovsky et al. (1980) considered the price volatility of multiple commodities, they only did so theoretically. But given that indirect utility functions – the usual measure of welfare in microeconomic theory – are defined over both income and a whole vector of prices, the literature’s focus on income risk, extended at most to a single stochastic price, paints an incomplete picture of complete (i.e., income *and* prices) attitudes toward risk as well as the impacts thereof. More concretely, the literature is of limited usefulness in informing the growing popular debates that surround food price volatility and food price stabilization policies.

In order to study whether households value multi-commodity price stability and whether the poor suffer disproportionately from food price instability, we combine the theoretical framework of Turnovsky et al. (1980) with the empirical framework developed by Finkelshtain and Chalfant (1991) and extended by Barrett (1996). This

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understanding of how changes in mean food prices affect welfare ever since Deaton’s (1989) seminal work on the topic.

allows us to derive a measure of multivariate price risk aversion – more precisely, an estimable matrix of price risk aversion over multiple commodities – and its associated willingness to pay (WTP) measures for price stabilization. We then apply this measure to estimate the heterogeneous welfare effects of food price volatility among rural Ethiopian households who both produce and consume several commodities characterized by stochastic prices and whose food storage capacity is extremely limited (Tadesse and Guttormsen, 2011). Prices in our data are highly variable: using the coefficient of variation (standard deviation/mean) for each commodity price retained for analysis as a rough measure of the importance of price volatility, the lowest such ratio equals 14 percent, and the highest such ratio equals 33 percent. In other words, the least volatile price in the data will on average rise or fall by about one seventh, and the most volatile price in the data will on average rise or fall by one third.

Each element of the matrix of price risk aversion coefficients we derive and estimate reflects the risk premium associated with the covariance between two of the prices faced by the household. On the diagonal, this yields familiar own-price risk aversion coefficients (Barrett 1996). These measure the *direct* impacts on welfare of the volatility in each price, i.e., the impact on welfare of the variance of each price, holding everything else constant. But because a price almost never fluctuates alone – commodities are typically substitutes for or complements to one another – the off-diagonal elements of the matrix of price risk aversion measure the *indirect* impacts on welfare of the volatility in a each price, i.e., the impacts on welfare of the covariance between a given price and the prices of all the other commodities considered, holding everything else constant. Of course, the indirect welfare impact of the covariance between the price of one commodity and the price of another must be symmetric. The symmetry implied by the theory offers a convenient test of the core microeconomic behavioral assumptions akin to that of the familiar Slutsky matrix.

To obtain the total welfare impacts of price vector volatility, one thus needs to consider both (i) the variance in each commodity price series as well as (ii) the covariances among these price series. Ignoring the covariances between prices leads in principle to a biased estimate of the *total* (i.e., direct and indirect) welfare impacts of price vector volatility unless the unrealistic assumption that commodities are neither complements nor substitutes holds, although the sign of the bias is impossible to determine *ex ante*. The off-diagonal terms (i.e., the indirect effects of price risk, or price covariance effects) of the matrix of price risk aversion have so far been ignored in the literature. Our analysis is the first to quantify their importance relative to the diagonal terms (i.e., the direct effects of price risk, or price variance effects) of the matrix of price risk aversion.

Based on the matrix of price risk aversion coefficients, we further show how to derive the household's WTP to stabilize at their means the prices of a set of commodities. Using panel data on rural Ethiopian households, we exploit the variation in prices within each household over time and between district-rounds to estimate the aforementioned measures of price risk aversion. We find that the average household is willing to give up 6 to 32 percent of its income to stabilize the price of the seven most important food commodities in the data. We also find that ignoring the covariances between prices would lead to underestimating household WTP to stabilize prices in this context.

Nonparametric analysis further suggests that in the rural Ethiopian context, contrary to conventional wisdom, the welfare gains of price stabilization are concentrated among the upper 40 percent of the income distribution, while more than 30 percent of the (poorer) population would suffer statistically significant welfare losses from price stabilization, although the magnitude of per capita losses among the latter subpopulation is much smaller than the magnitude of estimated gains among the wealthier winners. Hence the average net gains, as wealthier households who are largely net sellers of these food commodities benefit at the expense of poorer, largely net buyers.

The intuition behind this counterintuitive result is as follows. Stochastic food prices reflect temporal risk; price uncertainty is resolved as the moment of sale or purchase approaches. Temporal risk therefore hurts growers more than consumers because growers must commit resources to cultivation well before sales prices are realized. Indeed, for buyers with sufficient flexibility, price variability affords them options to time purchases for when prices have temporarily fluctuated below their mean. And because net sellers are systematically wealthier than net buyers in low-income agrarian economies (Barrett 2008), price stabilization therefore becomes regressive, but with aggregate net gains because growers lose so much more from price volatility than buyers gain. This result is not an artifact of the assumptions we make in the theoretical framework. Rather, it emerges from the estimation results to provide empirical support for the theoretical framework.

Given the strong political pressures to tackle food price volatility, we conclude with a simple thought experiment. We show that 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 loss they suffer due to price risk exposure but which leaves households who benefit from price volatility unaffected, may 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 in section 2. In section 3, we present the data and discuss some descriptive statistics. We then develop a reduced form empirical framework to estimate the matrix of price risk aversion coefficients and discuss empirical identification in section 4. In section 5 we estimate own- and cross-price risk aversion coefficients, test the symmetry 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.

2. Theoretical Framework

This section explores the welfare implications of multicommodity price volatility by relying on a two-period agricultural household model (see Appendix A for the basic model) and by deriving the household's matrix of price risk aversion coefficients.[7] Our interest in price instability requires at a minimum a two-period model,[8] with at least one period in which agents make decisions subject to temporal uncertainty with respect to prices, both in levels and in relation to incomes and other prices.

A similar version of 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. In what follows, we extend Barrett's 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.

In what follows, we abstract from credit market, storage, and informal transfer considerations. While incorporating the credit and informal transfer aspects of household behavior would undoubtedly make for a more realistic model of household behavior, we opt for a simpler specification so as to focus on the behavior of households in the face of temporal price risk. As regards storage, Tadesse and Guttormsen (2011, pp. 88-89) note that in Ethiopia, "smallholder farmers sell the bulk of their produce right after harvest to pay taxes and loans and to meet their cash requirements for social services, (…) few farmers store grain for long periods in order to benefit from temporal arbitrage," and how "storage cost is generally very high in Ethiopia." Enhancements to our admittedly parsimonious framework, which will have to be combined with more detailed empirical data, are thus left for future research.

[7] We opt for a unitary version of the agricultural household model as it is the most parsimonious model possible and because we need a framework that encompasses both the consumer and producer sides of household behavior.

[8] We caution the reader against interpreting our model as dynamic. The careful reader will notice the absence of subscripts denoting time periods in our theoretical framework. This is because the "dynamic" aspect is with respect to the resolution of uncertainty, with expectations denoting "first-period" (i.e., *ex ante*) variables. Inversely, the absence of expectations denotes "second-period" (i.e., *ex post*) variables.

2.1. 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

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

where V_y and V_{yp} are first and second derivatives, respectively.

Moreover, let M_i be the marketable surplus of commodity *i* (i.e., the quantity supplied minus the quantity demanded by the household of commodity *i*). By Roy's identity, i.e., $M_i = \frac{\partial V / \partial p_i}{\partial V / \partial y}$,[9] 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

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

[9] One can apply Roy's identity to the marketable surplus equation given that it is both additive and convex. See Barrett (1996) and 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

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

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 31 is thus equivalent to

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

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 becomes 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 volatility in the price of good *i* (i.e., the variance in the price of good *i*) and to covolatility in the prices of goods *i* and *j* (i.e., the covariance between the prices of good *i* and *j*) since its autarky from the market leaves it unaffected at the margin by price volatility.

Applying Young's theorem once again, we obtain 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, we obtain the V_{pp} matrix:

$$V_{pp} = \begin{bmatrix} V_{p_1p_1} & V_{p_1p_2} & \cdots & V_{p_1p_K} \\ V_{p_2p_1} & V_{p_2p_2} & \cdots & V_{p_2p_K} \\ \vdots & \vdots & \ddots & \vdots \\ V_{p_Kp_1} & V_{p_Kp_2} & \cdots & V_{p_Kp_K} \end{bmatrix}, \tag{16}$$

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

$$\begin{aligned} \mathrm{A} = -\frac{1}{V_y}\cdot V_{pp} &= -\frac{1}{V_y}\cdot \begin{bmatrix} V_{p_1p_1} & V_{p_1p_2} & \cdots & V_{p_1p_K} \\ V_{p_2p_1} & V_{p_2p_2} & \cdots & V_{p_2p_K} \\ \vdots & \vdots & \ddots & \vdots \\ V_{p_Kp_1} & V_{p_Kp_2} & \cdots & V_{p_Kp_K} \end{bmatrix} \\ &= \begin{bmatrix} A_{11} & A_{12} & \cdots & A_{1K} \\ A_{21} & A_{22} & \cdots & A_{2K} \\ \vdots & \vdots & \ddots & \vdots \\ A_{K1} & A_{K2} & \cdots & A_{KK} \end{bmatrix}, \end{aligned} \tag{17}$$

where

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

Matrix A has a relatively straightforward interpretation. The diagonal elements are analogous to Pratt's (1964) coefficient of absolute (income) risk aversion, but with respect to individual prices instead of income or wealth. 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 (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 (a speculator) over *i*.[10] 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 interpretation of the off-diagonal terms is a bit trickier in that those reflect how variations in the price of one good due to variations in the price of another good affect household welfare. Put simply, if A_{ii} captures the welfare impact of the variance of the price of commodity *i* holding other prices constant, the off-diagonal elements capture the impacts of price covariances. Consequently, $A_{ij} > (<) \ 0$ implies that for an indirect increase in the volatility of price *i* attributable to an increase in the volatility of price *j*, household welfare decreases (increases), i.e., the household stands to gain from hedging against (speculating over) covariance in the prices of goods *i* and *j*.

Taken as a whole, the price risk aversion coefficient matrix thus speaks directly to the total welfare effects of and household preferences with respect to multivariate price risk. Intuitively, each diagonal term can be interpreted as the (direct) effect on household welfare of the variance in the price of a single good, *ceteris paribus*. Similarly, each off-diagonal term can be interpreted as the (indirect) effect on household welfare of the covariance between the prices of two goods, *ceteris paribus*.

Perhaps more importantly, there is no theoretical restriction on the sign of any element of *A*. As per equation 18, the sign of A_{ij} depends on (i) whether the household is a net buyer or a net seller of commodity *i*, i.e., on the sign of M_i; (ii) the sign of the budget share of the marketable surplus of commodity *j*, i.e., β_j; (iii) whether the household's coefficient of relative risk aversion *R* is less or greater than the income

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

elasticity of the marketable surplus of commodity *j*, i.e., η_j; and (iv) the sign and magnitude of the elasticity of the marketable surplus of commodity *i* with respect to price *j*, i.e. ε_{ij}. The theory, however, implies a testable symmetry restriction on the estimated price risk aversion coefficients. With adequate data, one can test the null hypothesis

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

which, for a matrix of price risk aversion defined over K commodities, represents $K(K-1)/2$ testable restrictions. Intuitively, the empirical content of equation 19 is simply that the impact on household welfare of the covariance between prices *i* and *j* should be the same as the impact on household welfare of the covariance between prices *j* and *i*, which is analogous to symmetry of the Slutsky matrix. The next section characterizes analytically the relationship between the price risk aversion matrix A and the Slutsky matrix and shows how a test of the symmetry of A is a test of household rationality.

2.2. Relationship between the Price Risk Aversion and Slutsky Matrices

What is the relationship between the price risk aversion matrix and the Slutsky matrix? Let M_i be the household's marketable surplus of commodity i, which is a function of the vector of commodity prices *p* the household faces as well as of its income *y*. We know the Slutsky matrix *S* is

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

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 A_{ij} 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 matrices, 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

$$\begin{aligned}-\left(\frac{V_{p_ip_j}V_y-V_{yp_j}V_{p_i}}{V_y^2}\right)+\left(\frac{V_{p_iy}V_y-V_{yy}V_{p_i}}{V_y^2}\right)\cdot\left[\frac{V_{p_j}}{V_y}\right]&=\\ -\left(\frac{V_{p_jp_i}V_y-V_{yp_i}V_{p_j}}{V_y^2}\right)+\left(\frac{V_{p_jy}V_y-V_{yy}V_{p_j}}{V_y^2}\right)&\cdot\left[\frac{V_{p_i}}{V_y}\right].\end{aligned} \tag{25}$$

This last equation can then be arranged to show that

$$\left(V_{p_ip_j}-V_{p_jp_i}\right)V_y=V_{yp_j}V_{p_i}-V_{p_jy}V_{p_i}-V_{yp_i}V_{p_j}+V_{p_iy}V_{p_j}\,. \tag{26}$$

By Young's Theorem, we know that $V_{p_ip_j}=V_{p_jp_i}$, that $V_{yp_i}V_{p_j}=V_{p_iy}V_{p_j}$, and that $V_{yp_j}=V_{p_jy}$, 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 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 restrictions on the data.

2.3. Willingness to Pay for Price Stabilization

As we discussed in the introduction, policy makers routinely try to stabilize one or more staple good prices, but what are the welfare effects of such efforts if they are successful? This subsection derives the WTP measures necessary to establish the welfare gains from partial price stabilization, i.e., from stabilizing one or more commodity prices.[11]

In order to tackle this question with respect to the prices of K observable commodities, one first needs to compute the total WTP for those K commodities, which is obtained by computing the difference between (i) the consumer's utility if prices were held fixed at their respective expectations (i.e., the first term in the numerator below); and (ii) the expected utility of the consumer in the face of stochastic prices (i.e., the second term in the numerator below), such that

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

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

[11] 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).

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

In other words,

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

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

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

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 right-hand side of equation 32 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.

Equations 32 and 33 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 $i \neq j$. It is impossible to determine *a priori* the sign of the bias, which depends on the sign of the covariances and on the sign of the off-diagonal terms of the matrix of price risk aversion.

3. Data and Descriptive Statistics

We empirically demonstrate the theory developed in the previous section by estimating the price risk aversion coefficient matrix and household WTP for price stabilization using the 1994a, 1994b, 1995, and 1997 rounds of the Ethiopian Rural Household Survey (ERHS) data.[12] Tadesse and Guttormsen (2011, p. 88) note that in Ethiopia,

> "[a] rise or decline in price trend is not as bad as its variability. (…) [P]rice volatility and, more recently, food price inflation remain the overriding national

[12] 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.

> concerns. Post-reform grain prices are subject to significant and continuing interannual price volatility that ranks among the highest in the developing world."

The ERHS recorded both household consumption and production decisions using a standardized survey instrument across the rounds we retain for analysis. The sample includes a total of 1494 households across 16 districts (*woreda* in Amharic, the official language of Ethiopia) with an attrition rate of only 2 percent across the four rounds selected for analysis (Dercon and Krishnan, 1998).[13] The average household in the data was observed 5.7 times over four rounds and three seasons (i.e., three-month periods),[14] with only 7 households appearing only once in the data. The estimations in this paper thus rely on a sample of 8556 observations.[15]

In what follows, we focus on coffee, maize, beans, barley, wheat, teff, and sorghum, which are the most important seven commodities in the data when considering the fraction of households producing or consuming them. Table 1 presents descriptive statistics: a positive mean marketable surplus indicates that the average household is a net seller of a commodity, and a negative mean marketable surplus indicates that the average household is a net buyer of a commodity, so the average household is a net buyer of every commodity. For each commodity, a significant number of households have a marketable surplus of zero, however, because they neither bought nor sold that commodity.[16] Per equations 15 to 17, a household is price risk neutral for any commodity for which its net marketable surplus equals zero. In other words, it is unaffected by fluctuations in the price of that commodity.

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[13] Ethiopia is subdivided into eleven zones subdivided into districts, which are roughly equivalent to counties in the United Kingdom or United States.

[14] 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.

[15] The original data included several outliers when considering the marketable surpluses of the seven commodities we study. 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.

[16] Indeed, there were no cases where a household bought and sold a commodity in the exact same quantities.

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 coffee 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 price volatility 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,[17] 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 sales, off-farm income, and livestock sales per period. That said, average income from the aforementioned sources is different from zero in only about 82 percent of cases, which explains why the average seasonal income of about $94 ($376 annually) may seem low. When focusing on nonzero income, the average seasonal income increases to about $106 ($424 annually). These figures, while seemingly low, encompass the sources of income available in the data and reflect the extreme poverty prevalent in rural Ethiopia.

Table 3 also presents the budget share of each staple commodity. Food represents the overwhelming majority of rural Ethiopian household expenditures, at least 85 percent. This falls on the upper end of global estimates of such budget shares, reflecting the extreme poverty of this population, the conspicuous absence of much other than food to purchase in rural Ethiopia, and our inability to impute the value of home rental income and expenditure in the ERHS data. Purchases of teff and coffee represent the largest budget shares, with 21 and 15 percent of the average household budget, respectively.

Finally, because price variances and 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. Note that coffee exhibits

[17] As of writing, US$1 ≈ Birr 9.43.

by far the most price volatility. Since coffee is also one of only two crops (along with wheat) where net sellers' mean net sales volumes exceed net buyers' mean net purchase volumes – recall that net sellers are always price risk averse in the single stochastic price setting (Finkelshtain and Chalfant 1991, Barrett 1996) – these descriptive statistics suggest that stabilization of coffee prices is more likely to generate welfare gains than would stabilization of other commodity prices. The estimates we report in Section 5 corroborate this simple insight.

4. Empirical Framework

As defined previously, a household's marketable surplus of a given commodity *i*, $M_i(p, y)$, is the quantity harvested of that commodity net of the quantity purchased and the household's consumption of its own harvest. For each commodity, we thus estimate a reduced form regression of the marketable surplus of that commodity as a function of output prices and household income with controls for a range of observables and unobservables.

4.1. Estimation Strategy

The ERHS data include commodity prices and allow us to compute a measure of household income. We use district-round fixed effects to control for the input prices faced by each household in each district in each round as well as for macroeconomic factors such as inflation, interest rates, the international price of commodities, etc. Time-invariant household fixed effects provide further control for household-specific transactions costs related to distance from the main district market, social relationships that may confer preferential pricing, and other household-specific transaction costs that determine whether a household is a net buyer of a commodity, autarkic with respect to it, or a net seller of the same commodity (de Janvry et al., 1991; Bellemare and Barrett, 2006). The use of household and district-round fixed effects also controls for access to storage, so that our estimates should largely account for what little commodity storage there is in rural Ethiopia (Tadesse and Guttormsen, 2011).

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

$$M_{ik\ell t} = \alpha_i + \delta_i \ln y_{k\ell t} + \phi_i \ln p_{i\ell t} + \varphi_i \ln p_{j\ell t} + \lambda_i d_{k\ell} + \tau_i d_{\ell t} + \nu_{ik\ell t}, \tag{34}$$

where i denotes a specific commodity,[18] k denotes the household, ℓ denotes the district, 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; $d_{k\ell}$ is a vector of household dummies; $d_{\ell t}$ is a vector of district-round dummies 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.[19,20]

We estimate equation 34 over 1,494 households across four rounds and three seasons, clustering standard errors at the district level. No household was observed over all four rounds and three seasons; the number of observations per household ranged from one to six.[21] 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).[22]

Computation of own- and cross-price elasticities, of the income elasticity, and of the budget share of marketable surplus follows directly from equation 34. As an example, to obtain the estimated cross-price risk aversion coefficient $\hat{A}_{ij}$, one first computes budget

[18] Subscripts on coefficients thus denote coefficients from specific commodity equations.
[19] 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.
[20] 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.
[21] 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).
[22] *Tella* is a traditional Ethiopian beer made from teff and maize.

share $\hat{\beta}_j = M_j p_j / y$ and income elasticity $\hat{\eta}_j = \hat{\delta}_j / M_j$ using the estimates of equation for commodity *j*; and cross-price elasticity $\hat{\varepsilon}_{ij} = \hat{\varphi}_i / M_i$ using estimates of the equation for commodity *i*. The variables M_j, p_j, and M_i are available in the data. One then combines these estimates to obtain the point estimate

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

whose standard error is obtained by the delta method. 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 sample means.[23]

Given that our data do not allow directly estimating *R*, the coefficient of relative risk aversion, 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; Saha et al., 1994).

4.2. Identification Strategy

What would the ideal data set to estimate equation 34 look like? Ideally, one would want to randomize over all prices and income so as to obtain causal estimates of the δ, ϕ and φ parameters. Randomizing over macroeconomic factors such as prices is infeasible, however, while randomizing over incomes would be both costly and ethically suspect. Alternatively, one could rely on instrumental variables (IV) techniques to exogenize the variables on the right-hand side of equation 34 that are endogenous. In these data – as in every other household data set of which we are aware – there are no valid instruments for each of the prices plus household income, so IV estimation is similarly infeasible.

[23] 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).

The best feasible option is therefore panel data analysis that allows controlling for unobservable household, district and period characteristics. Household fixed effects should control for the systematic way in which each household forms its price expectations, and district-round fixed effects should control for departures from the systematic way in which each household forms its price expectations by accounting for the price information available to each household in a given district in a given time period. Likewise, if a household's status as a net buyer, autarkic, or a net seller with respect to a given commodity is primarily driven by its preferences for producing and consuming that specific commodity or by the household-specific transactions costs it faces (de Janvry et al., 1991; Goetz, 1992; Bellemare and Barrett, 2006), these factors are accounted for by the household fixed effect. While this panel data approach does not purge the error term of all its correlation with the explanatory variables in equation 34, it surely purges much prospective endogeneity and is ultimately the best one can do in terms of empirical identification on this important empirical question. Still, we caution the reader against interpreting our estimates for the coefficients in equation 34 as strictly causal, but we equally caution the reader against ignoring crucial policy questions for which ironclad identification is inherently elusive.

In the empirical work below, ϕ and φ – the marginal impacts on marketable surplus of the logarithms of own- and cross-prices – are identified by (i) the variation in prices within each household over time (given our use of household fixed effects); and (ii) the between-district variation within a given round and over time for each district (given our use of district-round fixed effects). For example, the price of maize is common to all the households in a given district in a given round, so controlling for the unobserved heterogeneity between households and the unobserved heterogeneity between district-round, ϕ and φ are identified because prices vary over time for each household and because prices also vary between each district-round both across space and over time. The identification of δ is more straightforward given that income varies both within households over time and between households in a given district within a given round.

Conditional on a household's status as a net buyer, autarkic, or a net seller, its purchase or sales of a given commodity is also driven by its preferences and by the household-specific transactions costs it faces but also by climatic and other environmental fluctuations that affect production (Sherlund et al., 2002), which are largely accounted for by the district-round fixed effect, and by prices and income, which we control for.

Because many households have a marketable surplus of zero for several commodities, we estimate several sub-matrices of price risk aversion coefficients.[24] We first test the A sub-matrix for the top three commodities consumed and produced by the sample households (i.e., coffee, maize, and beans; we label that sub-matrix A_3), and then test the sub-matrices defined by the top four, five, six, and seven commodities (we label these sub-matrices A_4 to A_7). With three different assumptions on relative risk aversion *R* and five different sub-matrices in each case, we generate a range of estimated WTP for incomplete commodity price stabilization and conduct a total of 15 tests of the null hypothesis of symmetry of the matrix of price risk aversion. The consistency of results provides some assurance as to the robustness of the empirical findings.

5. Estimation Results and Hypothesis Tests

This section first presents estimation results for the marketable surplus equation in equation 34 for all seven commodities retained for analysis. Given that these results are ancillary, we only briefly discuss them so as to devote the bulk of our discussion to the estimated matrix of price risk aversion and, more importantly, to our estimates of household willingness to pay for price stabilization.

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, 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

[24] 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.

marketable surplus of all commodities except wheat, for which the point estimate is statistically insignificantly different from zero.

The cross-price coefficients whose signs are consistent (i.e., of the same sign) among equations – for example, the coefficient estimate for the price of barley is negative in the coffee equation, and the coefficient estimate for the price of coffee is also negative in the barley equation – 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). Since these estimates reveal statistically significant complementarities and substitution effects, ignoring covariance effects would necessarily bias estimates of price risk aversion, as discussed previously.

5.1. Price Risk Aversion Matrix

We use the estimation results reported in 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. The ERHS households are significantly own-price risk-averse, on average, over all commodities. Table 6a reports estimates under the intermediate assumption of R=2; Appendix B tables B1a and B2a report similar estimates for R=1 and R=3, respectively. The average household appears most significantly own-price risk-averse over barley, maize and teff – the commodities with the greatest net purchase volumes – and least price risk-averse over coffee and beans, which have the lowest mean net sales volumes among net sellers and the lowest mean net purchases volumes among net buyers, as reflected in Table 2.[25] Most rural Ethiopians' price risk exposure to these latter commodities is quite modest, hence the relatively low price risk aversion coefficient estimates.

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

[25] 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.

multivariate context. All 42 off-diagonal point estimates are statistically significantly different from zero, and all of them are positive, indicating aversion to positive co-fluctuations in commodity prices, which limits substitution possibilities for both net buyers and net sellers. Looking at either the upper or lower triangle of matrix A in table 6a, households are most price risk-averse over co-fluctuations in the prices of (i) teff and maize; (ii) teff and wheat; (ii) wheat and barley; and (iv) teff and barley. Given that barley, maize, and teff are staples, it is not surprising that households get hurt the most by co-movements between their respective prices. Similarly, households are least price risk-averse over co-fluctuations in the prices of (i) sorghum and coffee; (ii) beans and coffee; and (iii) wheat and coffee. Given that coffee is a nonstaple, it is not surprising that households get hurt the least by co-movements between its price and the prices of other commodities.

We illustrate the necessity of our approach with the example of teff. First, note that in table 6a, households are, on average, risk-averse over the price of teff. This is the *direct* effect of fluctuations in the price of teff. Recall, however, that the covariances between price of teff and the prices of other commodities were all positive in table 4, so that an increase in the volatility of the price of teff is correlated with variation in other food prices, over which households are also risk averse. This generates an *indirect* welfare effect of volatility in the price of teff through its covariance with other food prices. To obtain the *total* welfare effect in the price of teff, one needs to consider the coefficient estimates in the "teff" row or the coefficient estimates in the "teff" column of matrix A, as in equations 32 and 33, as we discuss in the next section.

Before discussing welfare effects, however, recall that the theoretical framework in section 2 implied symmetry of the A matrix. Although 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,[26] each $\hat{A}_{ij}$ is extremely close to its associated $\hat{A}_{ji}$. In fact, computing $\min\{\hat{A}_{ij}, \hat{A}_{ji}\}/\max\{\hat{A}_{ij}, \hat{A}_{ji}\}$ for all $i \neq j$ in order to measure the discrepancies between

[26] This result is robust to alternative assumptions about the coefficient of income risk aversion R (see Appendix B).

matched off-diagonal terms, the minimum such measure equals just 0.966. In other words, two matched off-diagonal coefficients differ by, at most, by three percent, hardly a substantive deviation from symmetry. Thus, even though the formal statistical test rejects the symmetry hypothesis for the estimated A matrix, households certainly seem to behave remarkably similarly to how the theory developed in section 2 would predict.

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 similar.

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). Although the estimated coefficients of own-price risk aversion are greatest for the main staple crops (teff, maize and barley) for which net marketed surplus exposure is greatest – for both net buyers and net sellers – because coffee price volatility is more than two orders of magnitude greater than for any other commodity, WTP for price stabilization is nearly an order of magnitude greater for coffee than for any grain. This underscores that WTP for price stabilization is a function of both (i) the magnitude of a commodity’s price volatility, and (ii) a household’s market exposure, and thus price risk aversion coefficients.

In other words, the discrepancy between the coefficients in matrix A and the WTP measures is due to the fact that while the WTP measures in equations 32 and 33 include prices variances and covariances, the coefficients of price risk aversion A in equation 18 do not include these variances and covariances. Thus, while households are *a priori* relatively less risk-averse with respect to the price of coffee than they are for other commodities, the fact that their WTP to stabilize the price of coffee dominates their WTP to stabilize the prices of other commodities is due to the very high volatility in the price of coffee.

The average household's WTP estimate to stabilize the prices of these seven commodities is between 6 and 32 percent of its income, depending on one's assumed relative income risk aversion (6 percent for $R = 1$; 19 percent for $R = 2$; 32 percent for $R = 3$). That proportion is statistically significant at the one percent level, clearly indicating aggregate willingness to pay to stabilize food commodity prices in rural Ethiopia.

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. Consequently, in these data, covariances between prices matter. Ignoring them significantly underestimates the average welfare loss due to price risk.

In order to be more specific about the distribution of the welfare gains from price stabilization, figure 1 plots the results of a second-degree fractional polynomial regression of the estimated household-specific WTP to stabilize the prices of all seven commodities on household income, along with the associated 95 percent confidence band.[27] Three important features appear in figure 1 and the associated table 8.

First, a significant share (31 percent) 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 percent) are price risk-averse (i.e., the households whose WTP for price stabilization is statistically significantly positive). Thus the population is

[27] We refer readers interested in using fractional polynomial regressions to Royston and Altman (1997), who prove a good discussion of both the method as well as of its usefulness. See Henley and Peirson (1997) for an economic application.

roughly equally divided among those who favor, oppose or are indifferent about price stabilization (table 8).

Second, the significantly price risk-loving households are markedly poorer than the significantly price risk-averse ones. Poorer households are more likely to be net consumers of all goods because they own fewer productive assets on average, and pure consumer theory holds that consumers are generally price risk-loving due to the quasiconvexity of the indirect utility function (Turnovsky et al., 1980). Higher price variance means that consumers can regularly take advantage of greater price discounts relative to the mean.

Conversely, better-off households are more likely to be producers of some or all food commodities because they own more productive assets. Their price risk preferences are therefore more consistent with the predictions of pure producer theory, which holds that firms are generally price risk-averse (Baron, 1970; Sandmo, 1971). This reflects growers' relative disadvantage due to production lags and temporal price risk.

Table 8 shows the income percentile ranges for which households are statistically significantly 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 61 percent of the income distribution lose out from price stabilization, on average. This suggests that price stabilization would be a distributionally 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, however, depended on whether (i) price volatility stems from random volatility in supply or in demand; (ii) price volatility is the result of an additive or multiplicative shock; and (iii) supply and demand functions are linear. Our empirical 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 even though the population is roughly evenly divided among the price risk-averse, price risk-loving and price risk-neutral subpopulations. Combined with the previous points, this underscores how simple averages may mask essential heterogeneity that is important in both equity and political economy terms.

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 stake in the outcome (Olson 1965), the three preceding observations may help partly explain some of the political economy of food price stabilization in spite of heterogeneous preference for food price stability. Indeed, because economic policy is often subject to élite capture (i.e., the wealthy often have more of a say than the poor in the political process), these observations correspond relatively well with the "developmental paradox," i.e., the empirical regularity according to which the more developed a 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 Stylized Policy Scenarios

As is well known, pure price stabilization through price fixing regulations or buffer stock management introduces considerable distortions in the economy (Krueger et al., 1988; Williams and Wright 1991). In this subsection we therefore briefly consider an alternative to the laissez-faire and government-imposed price stabilization counterfactuals, a stylized price risk compensation scheme to fully compensate households who incur a welfare loss from price volatility, but which neither compensates nor taxes households who gain from price volatility. Although our previous results clearly indicate that such a policy would be distributionally regressive, as would any price stabilization policy, such a scheme merits consideration as an alternative to full-blown price stabilization if political pressure (perhaps from economic élites) effectively compels the state to act in some fashion so as to reduce food price volatility.

We begin by considering the effects of full price stabilization, i.e., a policy in which households who gain from price volatility are, in effect, fully taxed for their gains while households who lose out from price volatility are, *de facto*, fully compensated for their losses. This represents the naïve benchmark of pure price stabilization, ignoring (likely important) general equilibrium effects (Acemoglu, 2010). Table 9a characterizes the (in-sample) winners and losers from such a policy. Under an assumed relative risk aversion $R = 2$, 63% of the rural Ethiopian population would lose out from price stabilization. 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 about the logic of collective action among a relatively small number of big winners, even when a majority would lose out from the policy (Olson, 1965).

Table 9b then compares the social welfare changes for two policy options, as measured against a *laissez-faire* policy under which nothing is done about commodity price volatility.[28] The first intervention option is the pure price stabilization policy discussed above (column 5). Column (6) reflects a compromise option, a price risk compensation scheme in which those households who are price risk-averse receive full compensation for their exposure to price volatility but in which those households who are price risk-neutral and price risk-loving are unaffected. As shown, 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. Moreover, only the price risk compensation is Pareto-improving, precisely because it leaves 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 more than an order of magnitude greater than the average losses to the price risk-loving subpopulation.

[28] This highly stylized analysis ignores fiscal costs and general equilibrium effects, both of which cannot be quantified with the data at hand.

While this is just a highly stylized example, it serves to underscore how the heterogeneous welfare effects of food price risk exposure may require more nuanced and creative policy responses than are commonly mooted in current popular discussions. This is an area ripe for further research using more realistic general equilibrium models that take into full consideration the distortionary effects of tax policies necessary to raise the resources for compensatory payments.

6. Conclusion

This paper tackles a highly topical policy question, viz. "Is price stabilization net welfare enhancing and, if so, for whom and at whose cost?" Our contribution is mainly empirical, establishing that price stabilization yields net welfare gains in rural Ethiopia but in a distributionally regressive fashion. These results contrast with the conventional wisdom in current food policy debates, which commonly conflates increases in mean food prices – which clearly hurt poor net food buyers – with increased fluctuations around the (perhaps higher) mean. Our approach enables isolation and direct estimation of the welfare effects of food price volatility.

This involved modestly extending the relevant microeconomic theory so as to allow studying 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, which respectively relate to the direct impacts of a price's volatility (i.e., the variance of the price of each commodity) as well as its indirect impacts through other prices (i.e., the covariance between the prices of all commodities) on household welfare. We have also shown how testing for the symmetry of the matrix of price risk aversion coefficients is equivalent to testing the symmetry of the Slutsky matrix, although the former imposes less structure on the data than the latter and is in principle less to be rejected.

In the empirical portion of the paper, we estimate the matrix of price risk aversion coefficients using panel data from rural Ethiopia. We find that these households are on average significantly price risk-averse over the prices of specific commodities as well as over covolatility in the prices of the same commodities. Although we statistically reject the hypothesis of symmetry of the matrix of price risk aversion, the estimated differences are economically insignificant, lending weak support to the underlying theory. The contrast between the statistical and economic results is likely due to the precision with which we estimate the coefficients in the matrix of price risk aversion.

More importantly, the average household's willingness to pay to fully stabilize commodity prices at their means lies between 6 and 32 percent of household income, depending on one's assumption about Arrow-Pratt relative income risk aversion. This may very well explain governments' frequent 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 a distributionally regressive benefit incidence from price stabilization policy.

Finally, if and when the political economy of price stabilization compels a government to intervene to attenuate the impacts of commodity price volatility, we suggest a price risk compensation alternative to outright price stabilization. Holding administrative costs constant and ignoring general equilibrium effects, we demonstrate in a very simplistic illustration that a compensation scheme without market interventions might prove Pareto-superior to pure price stabilization, albeit still distributionally regressive. Given the renewed interest in this topic among policy makers at the national and international levels, the complex and heterogeneous welfare effects of multivariate commodity price volatility appears a topic that merits further exploration.

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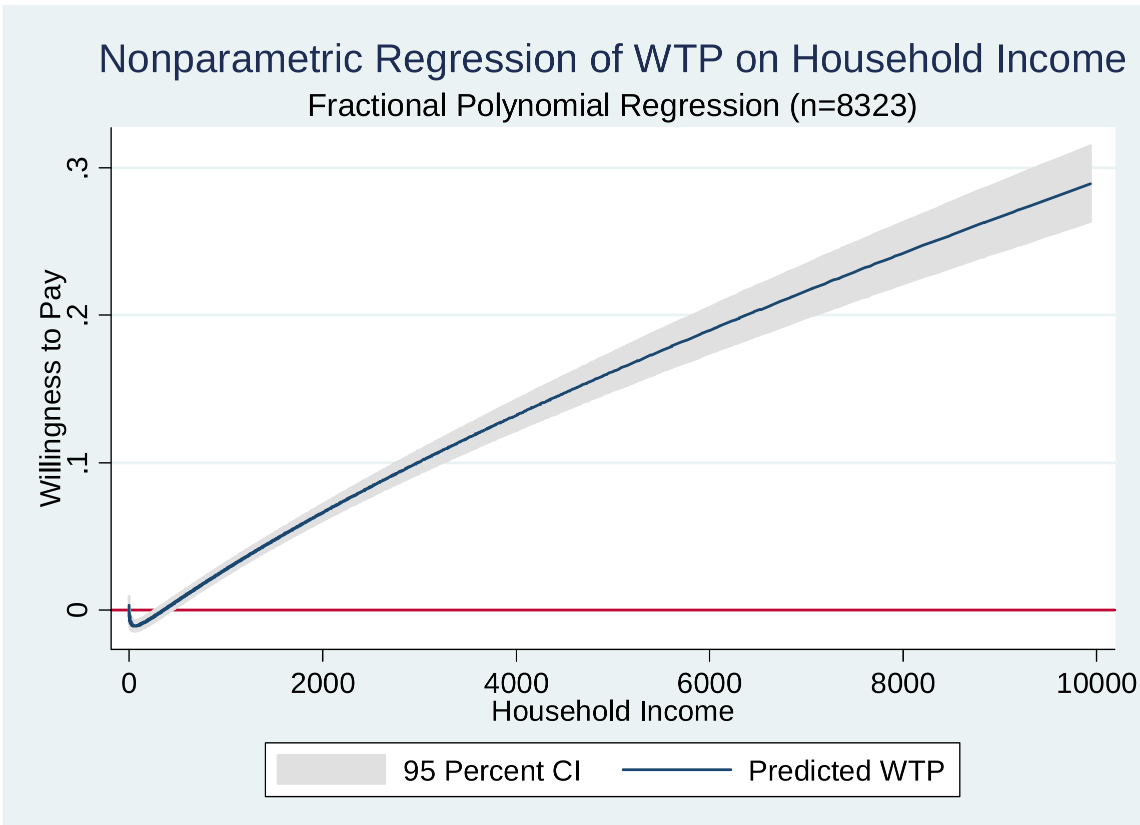

Figure 1: Fractional polynomial regression of household WTP to eliminate price volatility among seven staple commodities on household income for households whose seasonal income does not exceed 10,000 birr.

Table 1: Seasonal Descriptive Statistics for Crop Marketable Surplus (Full Sample, all in kg)

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

N=8556

Table 2: Seasonal Descriptive Statistics for Crop Marketable Surplus (Nonzero Observations)

Crop	Net Buyer Mean Marketable Surplus (kg)	(Std. Dev.)	Net Buyer Observations	Net Seller Mean Marketable Surplus (kg)	(Std. Dev.)	Net Seller Observations
Coffee	-23.44	(95.64)	6206	57.92	(95.02)	538
Maize	-397.18	(438.32)	3115	231.55	(388.10)	851
Beans	-127.14	(122.91)	2848	90.70	(95.32)	182
Barley	-459.27	(553.31)	2097	279.81	(329.47)	728
Wheat	-296.70	(337.00)	2420	434.74	(620.52)	376
Teff	-471.03	(453.10)	2136	269.06	(432.08)	530
Sorghum	-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.)
Real Commodity 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 of Marketable Surpluses		
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 zero incomes, budget shares were obtained by dividing marketable surpluses by mean nonzero income. A negative budget share indicates that the average household is a net buyer of a commodity.

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 Equation Estimates

	(1)		(2)		(3)		(4)	
Dependent Variable:	**Coffee Marketable Surplus**		**Maize Marketable Surplus**		**Beans Marketable Surplus**		**Barley Marketable Surplus**	
Coefficients	**Coeff. Est.**	**(Std. Err.)**	**Coeff. Est.**	**(Std. Err.)**	**Coeff. Est.**	**(Std. Err.)**	**Coeff. Est.**	**(Std. Err.)**
Log Coffee Price	**40.273*** **	**(0.092)**	96.297***	(0.688)	15.471***	(0.051)	-178.545***	(0.500)
Log Maize Price	-6.344	(3.703)	**389.529*** **	**(27.526)**	36.811***	(2.043)	118.277***	(20.027)
Log Beans Price	-9.567*	(5.359)	-115.378**	(39.841)	**39.952*** **	**(2.957)**	257.049***	(28.988)
Log Barley Price	-27.499***	(3.585)	-53.697*	(26.648)	7.396***	(1.978)	**305.961*** **	**(19.389)**
Log Wheat Price	30.689***	(5.884)	137.883***	(43.738)	146.613***	(3.246)	254.879***	(31.823)
Log Teff Price	104.537***	(7.905)	-66.515	(58.761)	-94.449***	(4.361)	-326.964***	(42.754)
Log Sorghum Price	-68.434***	(2.637)	-73.385***	(19.603)	-95.320***	(1.455)	-445.863***	(14.263)
Log Potatoes Price	12.659***	(1.007)	7.845	(7.488)	26.723***	(0.556)	37.081***	(5.448)
Log Onions Price	-24.624***	(3.258)	59.407**	(24.220)	-51.082***	(1.798)	-275.690***	(17.622)
Log Cabbage Price	-10.344***	(0.563)	52.844***	(4.182)	5.462***	(0.310)	73.095***	(3.043)
Log Milk Price	-13.161***	(1.182)	290.977***	(8.790)	-26.748***	(0.652)	18.605**	(6.395)
Log *Tella* Price	28.556***	(4.217)	131.295***	(31.345)	75.795***	(2.326)	307.160***	(22.806)
Log Sugar Price	11.445***	(3.310)	-151.995***	(24.602)	5.659***	(1.826)	21.166	(17.901)
Log Salt Price	6.754***	(1.944)	121.660***	(14.449)	-36.798***	(1.072)	-264.330***	(10.513)
Log Cooking Oil Price	-5.634**	(2.350)	-15.989	(17.470)	-91.614***	(1.297)	-362.422***	(12.711)
Log 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	
District-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. Because a household level of marketable surplus is regressed on the logarithm of prices and income, coefficients cannot be interpreted as elasticities.

Table 5 (continued): Seasonal Marketable Surplus Equation Estimates

	(5)		(6)		(7)	
Dependent Variable:	**Wheat Marketable Surplus**		**Teff Marketable Surplus**		**Sorghum Marketable Surplus**	
Variable	**Coeff. Est.**	**(Std. Err.)**	**Coeff. Est.**	**(Std. Err.)**	**Coeff. Est.**	**(Std. Err.)**
Log Coffee Price	2.688***	(0.290)	113.406***	(0.632)	22.879***	(0.369)
Log Maize Price	-85.971***	(11.603)	48.316*	(25.306)	-36.271**	(14.774)
Log Beans Price	-35.692**	(16.795)	-63.054	(36.628)	-68.184***	(21.383)
Log Barley Price	-46.119***	(11.234)	-58.085**	(24.499)	50.331***	(14.303)
Log Wheat Price	**17.469**	**(18.438)**	38.188**	(40.211)	-144.397***	(23.475)
Log Teff Price	235.372***	(24.771)	**123.266****	**(54.022)**	84.194**	(31.538)
Log Sorghum Price	-45.547***	(8.263)	3.172	(18.022)	**39.693*****	**(10.521)**
Log Potatoes Price	31.551***	(3.157)	10.700	(6.884)	-34.192***	(4.019)
Log Onions Price	-64.140***	(10.210)	103.915***	(22.267)	61.417***	(12.999)
Log Cabbage Price	21.528***	(1.763)	29.668***	(3.845)	8.517***	(2.244)
Log Milk Price	-134.989***	(3.705)	111.745***	(8.081)	25.431***	(4.718)
Log *Tella* Price	106.587***	(13.213)	-0.722	(28.817)	-80.962***	(16.823)
Log Sugar Price	43.907***	(10.371)	-175.316***	(22.619)	-22.747	(13.205)
Log Salt Price	-3.396	(6.091)	140.199***	(13.284)	8.750	(7.755)
Log Cooking Oil Price	8.577	(7.364)	12.020	(16.061)	75.466***	(9.376)
Log 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	
District-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. Because a household level of marketable surplus is regressed on the logarithm of prices and income, coefficients cannot be interpreted as elasticities.

Table 6a: Estimated 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: 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: Estimated 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: Estimated 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: Estimated 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: Estimated WTP for Price Risk Stabilization, By Income Range

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 significantly price risk-averse.

Table 9a: Estimated Welfare Gains and Losses from Eliminating Price Volatility

Coefficient of Relative Risk Aversion	(1) Average Welfare Gain (Birr)	(2) Percent of Households Who Would Benefit	(3) Average Welfare Loss (Birr)	(4) Percent of Households Who Would Lose Out
R = 1	439.37	25.8	67.23	74.2
R = 2	660.41	37.0	52.96	63.0
R = 3	836.28	45.7	43.35	54.3

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 volatility, 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) Per Capita Change in Social Welfare under Price Stabilization (Birr)	(6) Per Capita Change in Social Welfare under Compensation (Birr)
R = 1	63.4	113.3
R = 2	210.8	244.4
R = 3	358.2	381.8

Note: Column 5 describes the change in social welfare under a price stabilization 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 6 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. Agricultural Household Model

The derivations in this section closely follow those in Barrett (1996), who 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 agricultural household model. In what follows, we report the basics of the model. Readers interested in more detailed explanations and derivations of these findings are encouraged to consult those prior works.

Consider a representative 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 K goods whose consumption and/or production is observed and whose associated stochastic price vector is $p_o = (p_{o1}, p_{o2}, ..., p_{oK})$; a composite c_u of all goods whose consumption and/or production is unobserved by the econometrician and whose associated stochastic composite price is p_u;[29] and leisure ℓ. The function $U(\cdot)$ is 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*.

All K goods observed and the unobserved good can, in principle, be produced and consumed by the household.[30] 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

$$F_{oi}(L_{oi}, A_{oi}),\ i \in \{1, ..., K\}, \tag{A1}$$

[29] In order simplify the exposition, we refer to the vector of commodities whose consumption and production is unobserved by the econometrician as "the unobserved good" in what follows.
[30] For example, it is quite common in developing countries for rural household to grow a staple crop (e.g., barley, wheat, maize, etc.) and many other non-staple crops (e.g., coffee, beans, etc.) 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. Finally, households may switch from one category – net buyer, autarkic, or net seller – to another from one period to the next (Bellemare and Barrett, 2006).

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

$$F_u(L_u, A_u), \tag{A2}$$

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). A general function $h(\cdot)$ maps hired labor into family labor equivalent units. 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 assumed missing.

The household's time constraint is such that $L^m + \ell + \sum_i L_{oi}^f + L_u^f \leq E^L$, 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 $A^m + A^f \leq E^A$, where A^m is the amount of household land leased out on the tenancy market at parametric rental rate r; and $A^f \equiv \sum_i A_{oi}^f + A_u^f$ is the amount of household land devoted to the production of the observable and unobservable commodities, respectively. Likewise, A_{oi}^h and A_u^h are the amounts of leased in land devoted to the production of the observable and unobservable commodities, respectively, so that $A_{oi} \equiv A_{oi}^f + A_{oi}^h$ and

$A_u \equiv A_u^f + A_u^h$ 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. That is, all (stochastic) product prices are unknown when labor allocation decisions are made, but post-harvest prices are revealed before consumption decisions are made. The household's problem is thus to

$$\max_{\{A_{oi}^h, A_u^h, A_u^f, L^m, L_{oi}^h, A_{oi}^f, L_{oi}^f, L_u^h, L_u^f, A^m, \ell\}} E \max_{\{c_o, c_u\}} U(c_o, c_u, \ell) \tag{A3}$$

subject to

$$p_o c_o + p_u c_u \leq Y^*, \tag{A4}$$

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

$$+ \sum_i p_{oi} F_{oi}(L_{oi}, A_{oi}) + p_u F_u(L_u, A_u) + I, \tag{A5}$$

$$L_{oi} \equiv h(L_{oi}^h) + L_{oi}^f \quad \forall i, \tag{A6}$$

$$L_u \equiv h(L_u^h) + L_u^f, \tag{A7}$$

$$L^m + \ell + \sum_i L_{oi}^f + L_u^f \leq E^L, \tag{A8}$$

$$A^f \equiv \sum_i A_{oi}^f + A_u^f \tag{A9}$$

$$A^h \equiv \sum_i A_{oi}^h + A_u^h \tag{A10}$$

$$A^m + A^f \leq E^A \tag{A11}$$

$$h(L_{oi}^h) \in [0, L_{oi}^h], \text{ and} \tag{A12}$$

$$h(L_u^h) \in [0, L_u^h]. \tag{A13}$$

Given that the household's utility function is strictly increasing, preferences are locally non-satiated and so the constraints in equations (A4), (A8) and (A11) bind. 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 $y = Y^*/p_u$. Finally, assume that the household is income risk-averse, in the sense that $\frac{\partial^2 V}{\partial y^2} = V_{yy} < 0$.[31]

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

$$\max_{\{A_{oi}^h, L_{oi}^h, A_{oi}^f, L_{oi}^f, L_u^h, L_u^f, A^m, \ell\}} EV(\ell, p_i, y) \tag{A14}$$

subject to

$$\begin{aligned} Y = w[E^L - \ell - \textstyle\sum_i L_{oi}^f - \sum_i L_{oi}^h - L_u^f - L_u^h] + r[E^A - \sum_i A_{oi}^f - \sum_i A_{oi}^h - A_u^f - A_u^h] \\ + \textstyle\sum_i p_i F_{oi}(L_{oi}, A_{oi}) + F_u(L_u, A_u) + I\,. \end{aligned} \tag{A15}$$

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

with respect to L_{oi}^h: $E\left\{V_y\left(p_i \frac{\partial F_{oi}}{\partial L_{oi}^h} - w\right)\right\} \le 0$ ($=0$ if $L_{oi}^h > 0$), (A16)

with respect to A_{oi}^h: $E\left\{V_y\left(p_i \frac{\partial F_{oi}}{\partial A_{oi}^h} - r\right)\right\} \le 0$ ($=0$ if $A_{oi}^h > 0$), (A17)

[31] 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.

with respect to L_{oi}^{f}: $E\left\{V_y\left(p_i\frac{\partial F_{oi}}{\partial L_{oi}^{f}}-w\right)\right\}\leq 0$ ($=0$ if $L_{oi}^{f}>0$), (A18)

with respect to A_{oi}^{f}: $E\left\{V_y\left(p_i\frac{\partial F_{oi}}{\partial A_{oi}^{f}}-r\right)\right\}\leq 0$ ($=0$ if $A_{oi}^{f}>0$), and (A19)

with respect to ℓ: $E\{V_\ell - V_y w\}\leq 0$ ($=0$ if $\ell>0$). (A20)

Intuitively, the terms in parentheses in equations (A16) to (A19) mean that the household is a profit maximizer (i.e., it sets the value of its marginal product of labor equal to the wage rate, and the value of its marginal product of land equal to the rental rate), and equation (A20) means that the household will set its (expected) marginal utility of leisure equal to the marginal cost of leisure. This set of FONCs is similar to what is usually derived from the basic agricultural household model (Singh et al., 1986; Bardhan and Udry, 1999).

B. Additional Estimation Results and Robustness Checks

Table B1a: Estimated Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 1

	Coffee	Maize	Beans	Barley	Wheat	Teff	Sorghum
Coffee	**4.657*****	2.596***	0.860***	4.484***	1.732***	2.830***	0.650***
	(1.342)	(0.510)	(0.166)	(0.715)	(0.250)	(0.481)	(0.203)
Maize	2.582***	**159.592*****	3.907***	15.115***	11.381***	34.324***	5.871***
	(0.507)	**(18.598)**	(0.511)	(3.302)	(3.005)	(6.379)	(1.615)
Beans	0.900***	4.108***	**12.964*****	25.043***	10.759***	14.847***	2.802***
	(0.174)	(0.537)	**(1.101)**	(2.665)	(1.603)	(1.650)	(0.505)
Barley	4.387***	14.865***	23.418***	**231.808*****	31.455***	28.840***	7.267***
	(0700)	(3.247)	(2.492)	**(26.193)**	(4.279)	(4.880)	(2.263)
Wheat	1.808***	11.945***	10.737***	33.564***	**69.239*****	34.858***	4.340***
	(0.261)	(3.153)	(1.600)	(4.565)	**(15.110)**	(6.703)	(1.173)
Teff	2.843***	34.667***	14.259***	29.616***	33.549***	**131.799*****	7.362***
	(0.483)	(6.443)	(1.584)	(5.012)	(6.452)	**(15.074)**	(1.253)
Sorghum	0.638***	5.794***	2.629***	7.288***	4.082***	7.193***	**24.339*****
	(0.199)	(1.593)	(0.474)	(2.270)	(1.103)	(1.224)	**(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 B1b: 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 B2a: Estimated Matrix of Price Risk Aversion for Relative Risk Aversion *R* = 3

	Coffee	Maize	Beans	Barley	Wheat	Teff	Sorghum
Coffee	**13.493*** **	7.496***	2.567***	12.809***	5.162***	8.226***	1.860***
	(3.887)	(1.473)	(0.497)	(2.043)	(0.746)	(1.398)	(0.581)
Maize	7.482***	**460.839*** **	11.661***	43.174***	33.926***	99.761***	16.806***
	(1.470)	**(53.702)**	(1.524)	(9.431)	(8.956)	(18.539)	(4.622)
Beans	2.607***	11.862***	**38.699*** **	71.534***	32.072***	43.151***	8.020***
	(0.505)	(1.551)	**(3.286)**	(7.613)	(4.779)	(4.795)	(1.446)
Barley	12.712***	42.924***	69.909***	**662.159*** **	93.766***	83.822***	20.798***
	(2.027)	(9.376)	(7.440)	**(74.820)**	(12.754)	(14.184)	(6.477)
Wheat	5.239***	34.490***	32.050***	95.875***	**206.392*** **	101.315***	12.424***
	(0.757)	(9.105)	(4.776)	(13.041)	**(45.042)**	(19.484)	(3.357)
Teff	8.240***	100.104***	42.563***	84.598***	100.007***	**383.074*** **	21.077***
	(1.400)	(18.603)	(4.729)	(14.316)	(19.232)	**(43.813)**	(3.586)
Sorghum	1.848***	16.729***	7.847***	20.819***	12.166***	20.907***	**69.675*** **
	(0.577)	(4.601)	(1.415)	(6.484)	(3.287)	(3.557)	**(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 B2b: 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.