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Abstract: This paper considers the effects of real exchange rate depreciation on stochastic agricultural producer prices in low-income agriculture. The conventional wisdom, that real depreciation achieved through nominal currency devaluation stimulates tradables production, does not hold in the presence of price risk. In fact, real depreciation only yields a stimulative price signal (in the Sandmo sense of higher mean, lower variance) in the case of a particular subset of nontradables: "nontraditional exports" which have indeed responded vigorously to contemporary depreciation episodes. GARCH estimation of time-series price data on several commodities from Madagascar support the hypotheses generated by the analytical model.

***JEL* codes:** O1, Q1, F31

Key words: depreciation, production under risk, GARCH estimation, Madagascar

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1. Introduction

Considerable evidence exists that overvalued currencies have a strong dampening effect on agricultural output in developing economies (Krueger et al., 1988; Jaeger and Humphreys, 1988; Elbadawi, 1992; Ghura and Grennes, 1993). Real exchange rate depreciation achieved through nominal exchange rate adjustment has therefore been a central plank of economic adjustment programs in low-income agrarian economies over the past decade. The conventional wisdom holds that this increases agricultural prices and thus stimulates agricultural production (Schuh, 1974; Krueger et al., 1988; Dornbusch, 1988). Yet the agricultural sectors of developing countries in Africa and Latin America have exhibited uneven and generally weak supply response to considerable real exchange rate depreciation (Commander, 1990; Barrett and Carter, 1997), with the notable exception of nontraditional exports, which have flourished in the wake of real depreciation (Barham et al., 1992). If exceptionally low price elasticities of supply were the explanation for weak aggregate supply response to depreciation, one would not expect to find such robust aggregate response within the class of nontraditional products. This paper offers a simple partial equilibrium model, based on price risk and nontrivial transfer costs, to explain the puzzlingly weak and cross-sectionally variable agricultural supply response to real exchange rate depreciation observed in low-income economies. Empirical evidence from Madagascar corroborates the predictions of the model.

This issue is of considerable importance given contemporary emphasis on both macroeconomic adjustment and agricultural development in low-income economies.

Especially in Sub-Saharan Africa, real exchange rate devaluation has been the most common and substantial corrective introduced in structural adjustment programs, and agriculture receives special emphasis (World Bank, 1994). Yet previous theoretical and empirical research has paid insufficient attention to two ubiquitous features of low-income agriculture that heavily condition production patterns: relatively high transfer costs that impede tradability and few opportunities to mitigate the temporal price risk introduced by biological production lags. This paper therefore considers the partial equilibrium effects of real depreciation on stochastic agricultural prices in a sector comprised of both nontradables and tradables. Section II presents a simple analytical model that demonstrates the effects of real depreciation on the mean and variance of agricultural producer prices are conditional on *ex ante* sectoral tradability. The conventional wisdom, that real depreciation stimulates tradables production, does not hold when one admits price uncertainty and producer income risk aversion. In fact, real depreciation only yields a stimulative price signal—in the Sandmo (1971) sense of higher mean, lower variance—in the case of nontraditional exports, which have indeed responded vigorously to contemporary real depreciation episodes. In section III, empirical analysis of agricultural price and tradability patterns in Madagascar supports the hypotheses generated by the analytical model. A concluding section places previous studies' findings within this more general framework.

2. The Analytical Model

There are two reasons why real exchange rate depreciation changes the mean and variance of agricultural prices faced by a small open economy, i.e., one that is a price-taker in the world market. First, exchange rates convert prevailing world price distributions into domestic currency terms. The domestic currency price distribution thus changes directly with the real exchange rate. Second, liberalization affects the partitioning of price space between tradables and nontradables by changing the domestic currency border price for a commodity, the costs of intermediation, or both, potentially shifting commodities between alternative equilibrium pricing conditions. Unless the price distribution in the world market is identical to that in the domestic market under autarky, a shift between equilibrium conditions likewise changes domestic equilibrium price distributions.

Although the agricultural sector is often modeled as fully tradable, important subsectors in virtually all economies are nontradable due to nonzero market intermediation costs. In many countries these costs often constitute an especially large part of the border parity price, rendering a large portion of agriculture internationally nontradable (Ahmed and Rustagi, 1987; Delgado, 1992; Kyle, 1992; Barrett and Carter, 1997). Moreover, market intermediation costs are themselves a function of the exchange rate since they invariably incorporate a substantial tradable component (*e.g.*, fuel).

The effects of real exchange rate depreciation on stochastic agricultural prices can be modeled using a simplified version of the classic spatial equilibrium models of Samuelson (1952) and Takayama and Judge (1971). Consider an economy in which agricultural production employs only nontradable inputs—labor, land, and livestock services—so that in partial equilibrium domestic supply is invariant to the real exchange rate. Since this is a partial equilibrium model, further assume for the sake of simplicity that domestic demand is likewise independent of the exchange rate. The autarkic conditional equilibrium price distribution for each commodity i can be characterized by the density function $\delta_i(\mathbf{y}, u)$, with conditional expectation p_{in} and conditional standard deviation s_{in}, where $\mathbf{y}$ is a vector of exogenous variables, and u is a stochastic shock.[1] Each commodity is also traded in international markets, at an exogenous world price distribution, ω_i, with expectation p_{iw} and standard deviation s_{iw}. The real exchange rate, e, represents the price of foreign currency in domestic currency units, adjusted for relative inflation. Competitive intermediaries bringing goods from port to domestic markets, or vice versa, incur nonstochastic transfer costs, m, which have both a nontradable component (*e.g.*, local warehousing) and a tradable component (*e.g.*, fuel). If the difference between the expected border price and the expected autarkic price equals or exceeds transfer costs, the possibility of arbitrage integrates the local and world markets at prevailing international prices.[2] The equilibrium price distribution, π^*, is thus (subscripts are eliminated for clarity)

$$\pi^* = \begin{cases} \iota = we + m & \text{iff} \quad p_n \geq p_w e + m \\ \delta & \text{iff} \quad p_w e - m \leq p_n \leq p_w e + m \\ \varepsilon = \omega e - m & \text{iff} \quad p_n \leq p_w e - m \end{cases} \tag{1}$$

with the nonlinear conditional expectation function p^*

$$p^* = \begin{cases} p_i = p_w e + m & \text{iff} \quad p_n \geq p_w e + m \\ p_n & \text{iff} \quad p_w e - m \leq p_n \leq p_w e + m \\ p_e = p_w e - m & \text{iff} \quad p_n \leq p_w e - m \end{cases} \tag{2}$$

and the nonlinear conditional standard deviation function s^*

$$s_i = s_w e \quad \text{iff} \quad p_n \geq p_w e + m$$

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[1]Conditioning variables are henceforth dropped to reduce clutter, and stochastic prices are described by just two parameters, an approach consistent with expected utility maximization under fairly general conditions (Meyer, 1987).

[2]This does not impose an assumption of risk neutrality on intermediaries, for the risk premium in trading can be either negative or positive (Chavas, 1988).

$$s^* = \begin{cases} s_n & \text{iff} \quad p_w e - m \le p_n \le p_w e + m \\ s_e = s_w e & \text{iff} \quad p_n \le p_w e - m \end{cases} \qquad (3)$$

Assume that $s_n > s_e = s_i$ preliberalization, indicating that nontradables' prices are more volatile than tradables' prior to real depreciation.[3] Since transfer costs are a function of both distance from port and the characteristics of the commodity (*e.g.*, weight, perishability), there really exist distinct equilibrium price distributions for each local market-commodity pair. While one can proceed, without loss of generality, with the analytical model as if there were only one commodity-market pair under consideration, this will be an important consideration in the next section's empirical testing.

Mean equilibrium prices are depicted in Figure 1 for four different stylized commodities. This simple construction shows that the mean border parity price, eP_W, anchors a symmetric band with a width equal to transfer costs. Importing (exporting) occurs only if the mean autarkic market-clearing price is greater (less) than or equal to P_i (P_e), as with the commodity with demand and supply schedules D_i (D_e) and S_i (S_e), respectively. A nontradable commodity's domestic demand and supply schedules intersect at a price between the adjusted border parity prices, p_e and p_i. It will prove useful to distinguish crudely between two different classes of nontradables: near-importables (represented by D_{ni} and S_{ni}) for which autarkic equilibrium price is higher than the preliberalization border price parity ($p_i > p_{ni} > ep_w$), and near-exportables (represented by the D_{ne} and S_{ne} schedules) for which

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[3]Spatially broader markets are generally better able to self-stabilize since the positive covariance of local shocks diminishes with area. Moreover, the deeper the market in terms of numbers of participants and transaction volumes, the more substantial must a given shock be to require adjustments to equilibrium prices. For example, the coefficient of variation for annual African nontradables' prices commonly reaches 50%, far higher than for tradables' prices, due largely to the price inelasticity of demand and supply for nontradable staple foods (Sahn and Delgado, 1989).

autarkic equilibrium prices are lower than the preliberalization border price parity ($p_e < p_{ne} < ep_w$). When agricultural commodities are presumed tradable and price risk is disregarded, depreciation[4] clearly stimulates agricultural production. Relaxation of those two strong assumptions, however, yields a decidedly more complex story.

Assume an Arrow-Pratt risk-averse firm maximizes the expected utility of profits, following Sandmo (1971). Output thus increases (decreases) in the mean (variance) of the equilibrium price distribution.[5] If real depreciation increased only expected producer prices, depreciation would indeed stimulate production. But differentiation of (2) and (3) with respect to e shows that real depreciation increases both the mean and the variance of importables prices, as is apparent in equations (4) and (6) (assume for the moment that commodities are not allowed to shift among the three equilibrium conditions).

$$\partial p_i/\partial e = p_w + \partial m/\partial e \tag{4}$$

$$\partial p_e/\partial e = p_w - \partial m/\partial e \tag{5}$$

$$\partial s_i/\partial e = \partial s_e/\partial e = s_w \tag{6}$$

$$\partial p_n/\partial e = \partial s_n/\partial e = 0 \tag{7}$$

Real depreciation has ambiguous effects on mean exportables' price (5), and unambiguously increases the variability of exportables' prices (6). In this partial equilibrium framework, real depreciation has no effect on the nontradables' price distribution (7).

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[4]While real depreciation can result from either changes in relative price levels or from changes in the nominal exchange rate, I concentrate on the latter since it is most commonly used to adjust real exchange rates. Not all my results are generalizable to real exchange rate changes due to changing national price ratios. In particular, δ would no longer be independent of e.

[5]Finkelshtain and Chalfant (1991) and Barrett (1996) show Sandmo's (1971) results are not fully general. The present results apply strictly just to net seller households.

The second term in the equations (4) and (5) captures the effect of real depreciation on transfer costs. Real depreciation increases the price of tradable inputs (*e.g.*, fuel, spare parts, vehicles) required in agricultural marketing, which can reach 50% of marketing costs in Africa (Ahmed and Rustagi, 1987). Because transfer costs increase with real depreciation, $\partial p_i/\partial e > \partial p_e/\partial e$. Stated another way, depreciation induces an asymmetric shift to new boundary prices, p'_i and p'_e, because expanded marketing margins eat up a portion of the local currency gains from real depreciation, increasing the price space occupied by nontradables (Figure 2).

Now relax the earlier restriction that commodities cannot shift between categories. In Figure 2, real depreciation also renders one previously (near-exportable) nontradable commodity exportable, while the commodity that was importable becomes nontradable.[6] The endogenous contraction of importables' price space and corresponding expansion of the price space occupied by exportables and nontradables likely induces some commodities to switch equilibrium pricing conditions, with importables becoming nontradable and near-exportables becoming exportable. Indeed, the greatest potential beneficiaries from real depreciation are near-exportables producers, who can enjoy a rising expected price and falling variance if real depreciation bumps p_e above p_{ne}.

The impact of real exchange rate depreciation on the price distributions faced by agricultural producers is thus conditional on commodities' *ex ante* structural position *vis-à-vis*

[6]The asymmetric shift in tradables prices may help explain why the external adjustment of African agricultural trade balances has fallen disproportionately on the import side of the current account. The asymmetry is obviously limited since the rise in p_e is unbounded, but p_i cannot rise beyond, p_n, the commodity's autarkic equilibrium price.

the world market. For commodities importable before depreciation, both the expectation and the variance of the producer price increase, regardless of whether a shift in equilibrium conditions results. Predepreciation exportables face increasing variance and ambigous effects on mean price. For a risk-averse producer maximizing the expected utility of farm profits, the production effects of such changes to stochastic tradables prices are ambiguous. Nontradables are not affected by nominal exchange rate changes that cause real depreciation unless the shift is sufficient to render them internationally tradable. Indeed, real depreciation generates an unambiguously stimulative shift in the price distribution only for the near-exportables, "nontraditional exports" class.[7] The evolution of agricultural price distributions following real depreciation designed to stimulate agricultural production is clearly heterogeneous across subsectors because of this path-dependence. Consequently, there can be considerable intrasectoral variation in supply reponse without different price elasticities of supply, and aggregate supply response to real depreciation can vary considerably across economies of different structural characteristics.[8]

3. The Effects of Real Depreciation on Agricultural Prices in Madagascar

The simple model just developed yields several testable hypotheses. First, the endogenous repartitioning of price space in the wake of real exchange rate depreciation should increase exportables' proportion of aggregate agricultural production and decrease

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[7] This may be an important, overlooked part of the story of the stunning expansion of nontraditional cash crop exports in Latin America under neoliberal economic reforms (Barham, Clark, Katz, and Schurman, 1992).

[8] An especially relevant contrast is between those landlocked economies (of which there are 15 in Sub-Saharan Africa) facing considerable transfer costs and those with relatively inexpensive port access.

importables' share. Second, one should find a structural shift in the correlation between the border parity price and local market prices where depreciation induces a shift between equilibrium pricing conditions. If a commodity moves from a nontradable (importable) to an exportable (nontradable) equilibrium, the correlation should jump from (to) zero to (from) significantly positive.[9] Third, the effects of reforms on the mean and variance of agricultural price series, as identified in (4)-(7), can be tested. Mean real prices of exportables, importables and near-exportables that switch equilibrium regimes should increase with depreciation, with the mean of the remaining nontradables prices unaffected by real depreciation. Price variability should increase for all tradables and decrease for near-exportables-turned-exportables.

I test these hypotheses using monthly retail price data, 1983-91, for five commodities (dried beans, maize, manioc, rice, and potatoes) from 17 different regions in Madagascar. 1983-85 is used as the prereform benchmark period as major real depreciations started in the second half of 1986. Dramatic shifts in macroeconomic policy cumulatively generated a 58% real depreciation of the Malagasy franc (FMG), 1986-88. Allowing for full adjustment, given contractual and informational lags, 1989-90 is used as the postreform comparison period.

Since the predictions of the model are conditional on a commodity's status prior to real depreciation, the first step toward empirical testing is to establish the *ex ante* structural position of different commodity price series. First we estimate transfer costs, m, for each

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[9] The shift is between zero and one if one believes literally in the law of one price.

regional commodity price series. Distance-, season-, and time-varying transfer costs per ton were modelled using equation (8).[10]

$$m = v + s\,t(s\,km_d + km_m/3) \qquad (8)$$
$$\text{with} \quad s = 1 \text{ (Apr-Sep), 1.5 (Oct, Nov, Mar), or 2 (Dec-Feb)}$$
$$\text{and} \quad t = ½\,(f+r)$$

Transfer costs per ton, m, are modelled as the sum of transport and other variable handling expenses.[11] For want of reliable data with which to make cross-regional or intertemporal adjustments, v is assumed fixed in real terms at FMG 500/ton, based on the author's unpublished data from a survey of traders. Road kilometers from the nearest international port[12] are divided between all-season macadam routes (km_m) and dirt roads (km_d), since costs on the latter average twice those on the former.[13] Seasonal cost coefficients are lowest in the dry season (April-September), higher at the start and end of the rainy season, and highest during the heaviest rains and cyclones (December-February). Because seasonal effects are more pronounced on unimproved routes than macadam, they enter quadratically for the former but only linearly for the latter. The time-varying transport cost coefficient, t, is modelled as half fuel costs, f, and half repair and maintenance costs, r. Although all fuel, spare parts, and vehicles are imported, fuel prices remain regulated in Madagascar, and have

[10]Costs are assumed constant across commodities.

[11]This specification ignores fixed costs, which may be considerable if supporting public infrastructure (*e.g.*, roads, storage depots, physical security) is lacking, or if the volumes transferred are small (Kyle, 1992).

[12]The four major international ports (Antsiranana, Mahajanga, Toamasina, and Toliary) accounted for more than 80% of gross international shipping volume to or from Madagascar, 1980-86, and were the only ports to handle at least 100,000 tons per year (BDE, 1988).

[13]Distance and road condition are from the late 1980s (UNDP/MEP, 1991). Road conditions are assumed constant over the period, which likely understates the increase in transfer costs since Madagascar's road infrastructure degraded during the latter 1980s and 1990s (World Bank, 1991).

not increased at either the same rate or time as exchange rates. The cost of maintenance and repairs, in contrast, has increased more or less directly with changes to the real exchange rate.[14] Estimated real transfer costs in Madagascar increased more than 15% in U.S. dollar terms between 1984-85 and 1989-90, to an average of $0.75 per kilometer-ton confirming that nontradables' price space expanded with real depreciation. These figures are high by international standards but not for Africa (Ahmed and Rustagi, 1987; Delgado, 1992).

The next step is to apportion each region-specific commodity price series among the *ex ante* exportable, importable, and nontradable categories for both the pre- and post-depreciation periods. If the local price series consistently fell within the nontradables band established by the border parity price plus and minus transfer costs, the commodity-region was assigned to the nontradables group. If it fell along the upper adjusted border parity price boundary it was assigned to the importables group, and price series along the lower adjusted border parity price boundary were considered exportables.[15]

As one might expect, there is considerable variation across commodities and regions in *ex ante* tradability status. The top panel of Table 1 summarizes the composition of agricultural production across exportables, importables and nontradables during the pre- and

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[14]FMG unit values for vehicle parts imports are used for r, while regulated retail diesel fuel prices, including the fuel tax surcharge, are used for f. Both data series come from unpublished BDE data and are deflated by the CPI to real FMG, base January 1983, to correspond to the real commodity price series.

[15]Unnevehr (1984) employed a similar method in a study of cassava in Indonesia. These qualitative decisions were corroborated by calculating the 1984-85 pairwise correlation coefficients between monthly local and border parity prices. Domestic and world prices were negatively correlated for more than half the nontradables, and only 7% were statistically significantly greater than zero at the 5% level (using a one-tailed test). All the tradables had nonnegative correlation coefficients and 50% of them were statistically significantly positive.

postreform periods.[16] The first column indicates what percentage of total national commodity production is covered by the data used in this estimation. It is considerable for beans, manioc, and rice, slightly better than half for potatoes, and weak for maize. The observed increase in exportables and nontradables and the corresponding decrease in importables following sharp real depreciation is consistent with the model's first hypothesis: that real exchange rate depreciation leads to an increase in the exportable proportion of agricultural production and a decrease in the importable proportion.

These figures may help explain the strong growth observed in nontraditional exports of beans, maize, and manioc in the wake of real depreciation, as shifting equilibrium pricing conditions from nontradability to exportability brings both higher expected producer prices and lower volatility, a combination attractive to risk-averse producers. Together these three commodities increased their share of Madagascar's merchandise export revenues from 0.3% to 2.4% between 1984-85 and 1989-90 on the strength of export volume increases of 8,530%, 575%, and 4,933% for dried beans, maize, and manioc, respectively.

The bottom panel of Table 1 presents the same information looking across regions rather than commodities. Again, one finds increasing exportables and nontradables shares and decreasing importables shares. Note the striking geographic differences in the impacts of depreciation. The regions surrounding the four major ports, where transfer costs are relatively insignificant, saw a marked shift from nontradables to tradables. But in more remote regions, macroeconomic reforms had little effect on the tradability status of

[16] Note that each crop was classified as exportable, importable, or nontradable in each region for which sufficient data were available. Thus a crop can fall in more than one category, reflecting intranational geographic diversity in marketing patterns.

agricultural commodities. Ironically, reforms, designed to integrate domestic agriculture more fully with international markets, may do the opposite in some regions, as nontradables' share of sectoral output expands due to increased transfer costs.[17]

Now consider the second and third hypotheses regarding the effects of real depreciation on the correlation between local and border parity prices and on the mean and variance of local market prices. For this I use generalized autoregressive conditional heteroskedastic (GARCH) econometric techniques, which permit simultaneous estimation of the conditional mean and variance of a dependent variable in the presence of autocorrelation in both moments, as is common in monthly price data.

The predicted responses of the mean and variance of regional price series, as well as those series' correlation with border parity prices, are conditional on both the *ex ante* and *ex post* structural status of the regional commodity price series. Consequently, several estimations were run. Available regional price series were first separated by commodity, in the belief that preferences and technology are more likely similar across regions for a given commodity than across commodities for a given region. The regional commodity price series were divided into subsamples distinguished by pre- and postdepreciation tradability conditions. Table 2 shows the distribution of region- and commodity-specific price series across the subsamples, by *ex ante* and *ex post* tradability status.[18] Since the objective of the

[17]Sectoral variation can thus be understood as the aggregate manifestation of commodities moving between different equilibrium price regimes. In this way, the present model provides microfoundations for macroeconomic findings of time- and policy-varying sectoral tradability (Mundlak, Cavallo, and Domenech, 1990).

[18] Depreciation implies a lower triangular matrix (Table 2), while appreciation implies an upper triangular matrix.

estimation is to derive "aggregate" parameters for the relation between real depreciation and stochastic agricultural prices, the regional data within each subsample were pooled and estimated as a panel. Each subsample price series, $\{P_t\}$, is represented by an autoregressive model of the general linear GARCH(p,q) form

$$\begin{aligned}
\beta(L)\, P_t &= \beta_0 + \sum_{i=1}^{m} \beta_i X_{it} + u_t \\
u_t &= h_t^{1/2} v_t \qquad (9) \\
h_t &= \alpha_0 + \sum_{j=1}^{p} \alpha_j u_{t-j}^2 + \sum_{j=1}^{q} \phi_j h_{t-j} + \sum_{k=1}^{n} \gamma_k Z_{kt}
\end{aligned}$$

where $\beta(L)$ is a polynomial lag operator, β_0 and α_0 are constants, the vector X contains the m exogenous variables included in the conditional mean equation, the vector Z contains the n exogenous variables included in the conditional variance equation, which has a GARCH(p,q) structure, and the v_t are independent, standard normal errors. This model was estimated for each of 13 commodity- and structure-specific price series panels.

Estimation proceeded first from identification of the autoregressive dimensionality of the conditional mean equations. Following classic Box-Jenkins techniques, I estimated, plotted and examined the sample autocorrelations and partial autocorrelations for each series. The smooth dampening in each series of the plotted autocorrelations and the abrupt fall of the second and successive plotted partial autocorrelations to the neighborhood of zero strongly suggested an AR(1) structure to the polynomial lag. Consequently, the conditional mean equations in (1) employ P_t as a dependent variable and P_{t-1} as an explanatory variable. None of the autocorrelations exceeded 0.92[19] and the autocorrelation series always dampened to

[19]In fact, only two exceeded 0.90 and half were 0.80 or less.

near zero within 24 months, suggesting stationarity. Llung-Box-Pierce portmanteau Q-statistics associated with the residuals of the estimated AR(1) series support this identification and augmented Dickey-Fuller tests confirm stationarity.[20]

The real exchange rate (RER), calculated as the nominal bilateral FMG/US$ rate deflated by relative CPI between the two countries, is the element of greatest interest in the X and Z vectors. The coefficient on RER tests the hypothesized relation of real depreciation to the mean and variance of prices. Since exchange rate depreciation took place in the context of considerable trade liberalization and other macroeconomic reforms, X and Z also contain a dummy variable for the reform era (REFORM), taking unit value for all observations from January 1986 on. The real FMG border parity price—the prevailing international market price multiplied by the nominal exchange rate and deflated by Madagascar's CPI[21]—was multiplied by the reform dummy variable and by its complement to create two variables: the prereform border parity price (pre-BPP), and the postreform border parity price (post-BPP). The sign and significance of the coefficients on these two variables test the hypothesis that a switch in equilibrium pricing conditions changes the correlation between domestic and international market prices.

The X and Z vectors also include regional dummy variables (except Imerina Centrale, the base region) to capture spatial fixed effects. A dummy variable was also included to capture potential disruptions attributable to a national strike, June-December 1991, and

[20] Diagnostic statistics available from the author by request.

[21] In the case of rice, which is a major component of Madagascar's consumer price index, nominal series were deflated by the CPI excluding rice.

monthly releases from the national rice buffer stock (activated in November 1986 and discontinued in 1990) were also included since the government explicitly intended buffer stock management to influence food prices. Finally, a GARCH (1,1) structure was assumed in the conditional variance equations.[22]

The results for the coefficient estimates of interest are displayed in Tables 3 through 7; asymptotic standard errors are in parentheses. Table 3 concerns commodity price series that were importable both before and after depreciation. The analytical model of section 1 predicts that both the mean and conditional variance are increasing in RER. The positive and statistically significant (at the 5% level)[23] coefficients on RER confirm those predictions. The resulting elasticity point estimates of the mean and variance of rice prices with respect to changes in the real exchange rate, evaluated at the 1983-85 period mean, are 0.46 and 0.17, respectively. Given better than 50% real depreciation, 1986-88, the estimated real effects on the mean and variance of producer prices were considerable.

Table 4 presents the estimation results for the panels of commodity price series that were importable before reforms but became nontradable as the band of nontradability widened and shifted upward in price space following depreciation. The model predicts mean and

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[22]Each panel was estimated by maximum-likelihood, employing a quasi-Newton method, numerical derivatives, and starting values derived from OLS in the maximization of the log-likelihood function. Specifically, the Broyden-Fletcher-Goldfarb-Shanno (BFGS) algorithm was chosen because, like the more commonly employed Davidon-Fletcher-Powell (DFP) algorithm, it has superlinear convergence but, unlike DFP, BFGS updates the Hessian approximation directly. This avoids inversion of the Hessian approximation, which often tends DFP toward singularity (Walsh, 1985). Because direct updating algorithms sometimes take many iterations to build up the covariance matrix, a low convergence criterion (0.00001) was chosen to ensure sufficient iterations.

[23]The term "statistical significance" is henceforth used with respect to hypothesis testing at the 5% level of significance, unless indicated otherwise.

variance of prices increase with RER and that domestic prices were positively correlated with border parity prices before devaluation, but not after. Again, the coefficients on RER in the conditional mean and variance equations are positive and statistically significant, with the exception of the conditional variance equation for rice, where the estimate is negative and insignificantly different from zero. The estimated elasticities of mean domestic price with respect to the real exchange rate, again evaluated at the prereform mean, are 0.90 and 1.57 for beans and rice, respectively. The elasticity of the variance of bean prices, with respect to RER, is estimated at 0.49. The estimates from Tables 3 and 4 are consistent with the analytical finding that real exchange rate devaluation positively affects the mean and variance of prices for *ex ante* importables, whether or not the commodity subsequently becomes nontradable.

By contrast, the estimated coefficients of pre-BPP and post-BPP do not support the hypothesis that the shift in equilibrium pricing conditions changed the correlation between domestic and world market prices. The estimated correlations were quite weak and not statistically significant in either commodity series, reflecting adversely on the relevance of the law of one price in this setting, perhaps due to the persistence of government intervention in commodity pricing even after substantial market-oriented reforms.

One would expect depreciation to have little direct effect on commodities that were nontradable both before and after reforms. That is precisely what one finds in Table 5, which reports estimation results derived from the regional price series for maize, manioc, potatoes, and rice that remained nontradable over the 1983-91 period. The coefficient on RER is of low magnitude in all four of the conditional mean equations, and statistically insignificantly different from zero in each of the conditional mean and variance equations. Real devaluation had no consistent impact, either positive or negative, on commodity prices in regions insulated from international trade by high intermediation costs.

Promotion of “nontraditional” exports occupies an important place in the objectives and rhetoric of economic liberalization, no less in Madagascar than elsewhere. Of the five commodities considered here, only potatoes were traditionally exported from Madagascar prior to the massive real depreciation of the 1980s. As mentioned earlier, export volumes of beans, maize, and manioc exploded in the latter half of the 1980s. The analytical model suggests that commodities shifting from a nontradable equilibrium to an exportable equilibrium exhibit a positive relation between RER and mean price, and a negative relation between RER and the variance. The correlation between domestic and world market prices should also become significantly positive once the commodity is tradable.

Table 6 presents the estimation results derived from the regional price series for dried beans, maize, manioc, and potatoes that were nontradable prior to major depreciation but then became exportable in the late 1980s and early 1990s. As predicted, the coefficients on RER are generally positive in the conditional mean equations and negative in the conditional variance equations. The exception is the coefficient estimate in the conditional variance equation for potatoes. In investigating this anomaly, I discovered that while the other four commodities met the analytical model's assumption that $s_n > s_e$, the assumption was violated for potatoes. The seemingly contrarian sign of the RER coefficient estimate in the conditional variance equation for this potatoes subsample is thus entirely consistent with the analytical model.[24] Most of the coefficient estimates are statistically significantly different from zero. For those variables, the estimated elasticities of mean price with respect to the real exchange rate, evaluated at 1984-85 means, are 0.63, 1.33, and 0.96 for beans, maize, and manioc, respectively. Contrast these values with the near-zero elasticities estimated for the same commodities in regions where the commodities remained nontradable in the wake of reforms. Nontraditional exports promotion is plainly region-specific. The same results apply to the elasticity of the variance of local market prices with respect to the real exchange rate. Those estimated values are -1.02 and -0.20 for maize and manioc, respectively. Real depreciation generally dampens price risk for nontradables becoming exportables.

Unlike in the case of tradables becoming nontradables (Table 4), the hypothesized structural switch in the correlation between domestic and international market prices finds support in the estimates reported in Table 6. All four of the pre-BPP coefficient estimates were near zero and statistically insignificant, while all four of the post-BPP coefficients were positive, larger than the corresponding pre-BPP coefficients and, with the exception of manioc, statistically significant. All of the coefficient estimates were well below unity, again suggesting that the law of one price does not hold strictly in the current setting. The evidence from Tables 4 and 6 is thus mixed regarding the hypothesis that structural change in equilibrium pricing conditions, from tradability to nontradability or vice versa, engenders discontinuity in the correlation between domestic and world market prices, with the pairwise correlation coefficient equalling zero under nontradability and being positive under tradability.

[24]The coefficient of variation of the domestic nontradables price series for potatoes varied from 0.07 to 0.17, while the coefficient of variation on the border parity potato price was 0.31.

The present data and estimation methods are indeterminate as to whether tradability really brings with it closer correspondence to international market price signals.

The final structural permutation estimated was for regional price series exportable both prior to and following depreciation. The model predicts a positive relation between real depreciation and the variance of prices for these subsamples. As reported in Table 7, the estimates weakly support this hypothesis. The coefficient estimates for the RER variable are positive in each of the conditional mean and variance equations, and statistically significant in all except the conditional variance equation for beans. The estimated elasticities with respect to the real exchange rate are 0.60, 0.33, and 0.40 for mean bean prices, mean potato prices, and the variance of potato prices, respectively. Note that the elasticity estimate of mean bean prices in this subsample, 0.60, is essentially the same as it was in the subsample moving from a nontradable equilibrium to an exportable one: 0.63. The empirical results presented in Tables 3 through 7 generally support the second and third hypotheses derived from the analytical model. The effect of real exchange rate depreciation on the moments of domestic commodity price distributions depends fundamentally on *ex ante* structural situation of particular commodities. Statistically, this conclusion was validated by Chow tests of structural shift in the parameter estimates. The GARCH models were rerun for each commodity, stacking the data from a pair of distinct structural subsamples (*e.g.*, the rice data underlying the estimates in Tables 3 and 4). Dried beans, potatoes, and rice each had three different subsamples and thus three different possible pairings for which Chow tests statistics were computed for these commodities, one each for maize and manioc. Each of the eleven test statistics far exceeded the relevant χ^2 critical value, yielding rejection at the 1%

significance level of the hypothesis that the true parameters are equal across the structural distinctions.

4. Conclusions

Although real exchange rate depreciation is commonly thought to be stimulative to tradables, in an environment of stochastic prices it actually has ambiguous net effects on the incentives faced by risk-averse tradables producers. Indeed, real currency depreciation favors nontraditional exports over all other commodity classes, in that only commodities moving from nontradability to exportability might enjoy rising mean prices coupled with falling price variability, an attractive combination to risk-averse producers maximizing the expected utility of profits. This may help account for the robust export volume response of nontraditional exports to real exchange rate depreciation around the world, as exemplified by dried beans, maize, and manioc in Madagascar. Just as the gains from depreciation are localized in commodity space, so do the benefits tend to concentrate spatially, in regions with sufficiently good access to ports that increasing intermediation costs do not preclude international commerce.

The analytical model developed here provides a very simple partial equilibrium foundation for several, previously unintegrated empirical observations surrounding the relation between exchange rate depreciation and the moments of stochastic agricultural price distributions. For instance, it qualifies the econometric results of Balassa (1990) and Diakosavvas and Kirkpatrick (1990), and the (CGE) simulation results of Dorosh (1994), who find a positive correlation between real depreciation and agricultural exports as a percentage of national income. These results simply capture the one-sided expansion of exportables' price-space depicted in Figure 2 and shown empirically in Table 1. Many previous studies have too quickly generalized from an analysis of export crops to the agricultural sector more broadly in asserting the stimulative effects of real exchange rate depreciation on prices and production. Given that nonexportables may dominate in low-income agriculture (Kyle, 1992; Barrett and Carter, 1997), this is a serious error of interpretation that may have contributed to excessively optimistic expectations of what exchange rate reforms might accomplish for low-income agriculture.

The present results are also consistent with the existing literature linking macroeconomic performance and sector-specific risk phenomena. Both empirical and theoretical studies have found that "macroeconomic conditions are an important source of agricultural price variability" (Lapp and Smith, 1992, p. 8). In particular, inflation, which is fueled by depreciation, tends to be positively related to relative price variability in agriculture (Lapp and Smith, 1992) and economywide (Fischer, 1981).[25] Guillaumont (1992) found a direct empirical relation between exchange rate policy and producer price variability in lower income countries. He observed that "the instability of real producer prices (both for food and export crops) was highest in the countries where monetary depreciation was strongest. It was lowest in countries adjusting without depreciation or with moderate depreciation" (p. 217).

[25]Although Lapp and Smith (1992) found the estimated direct relationship between foreign exchange rate movements and agricultural price variability was not robust to changes in specification; this may result from their use of an aggregate price variability index. Since the sign of the relation between real exchange rate movements and agricultural price variability differs across classes of commodities defined by *ex ante* equilibrium conditions, real depreciation is ambiguously related to the aggregate sectoral price index.

Overvalued currencies appear to have had a strong dampening effect on agricultural output in many low-income economies prior to the economic reforms of the 1980s (Krueger, Schiff, and Valdés, 1988; Jaeger and Humphreys, 1988; Elbadawi, 1992; Ghura and Grennes, 1993). As a consequence, many analysts and policymakers expected real depreciation to have expansionary effects for agriculture. But those expectations have been based on assumptions of full tradability and risk neutrality that are likely untenable in low-income agriculture. This paper makes the case that where a nontrivial portion of the agricultural sector is nontradable, considerable cross-sectional variation in the response of stochastic prices to real exchange rate depreciation is to be expected. Thus, the aggregate sectoral effects on the incentives faced by risk-averse producers are inherently ambiguous, depending fundamentally on the *ex ante* structure of the sector.

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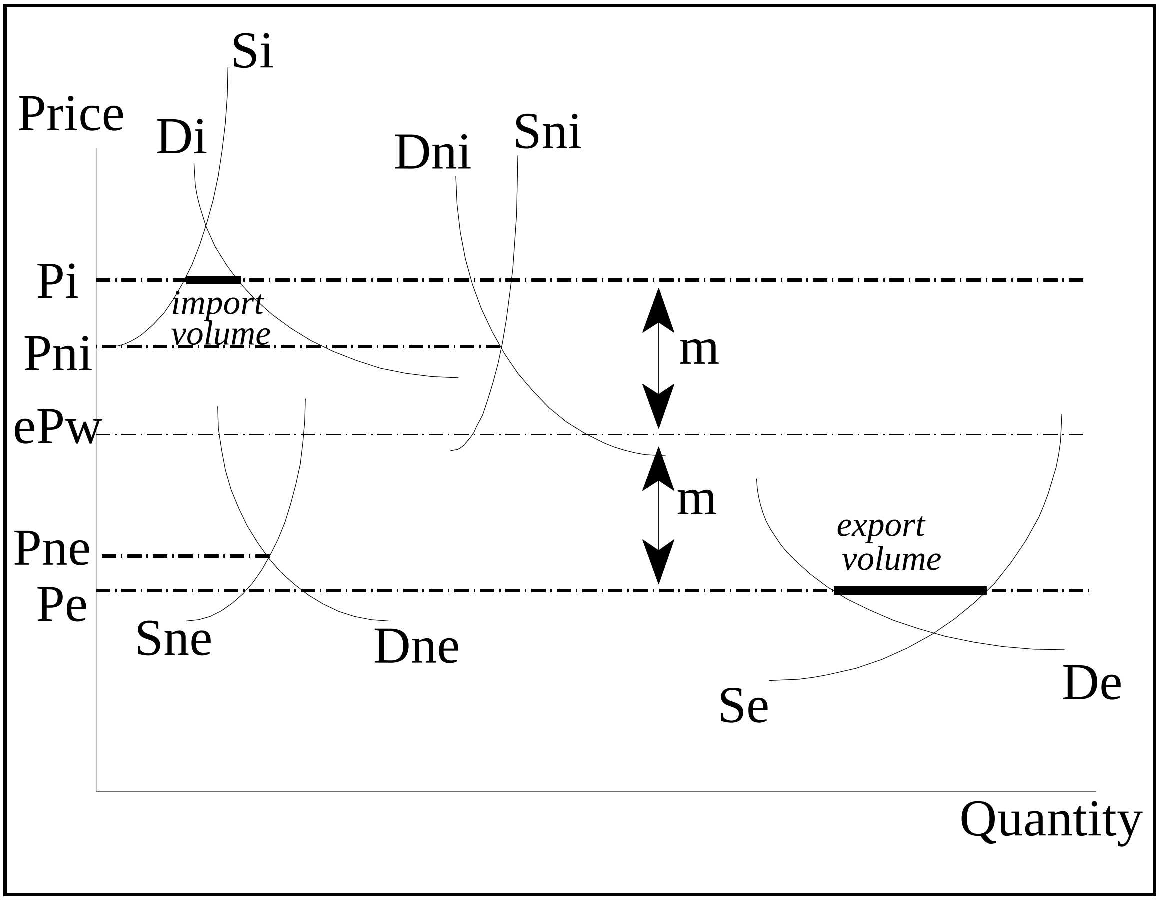

Figure 1: *Ex ante* commodity price equilibria

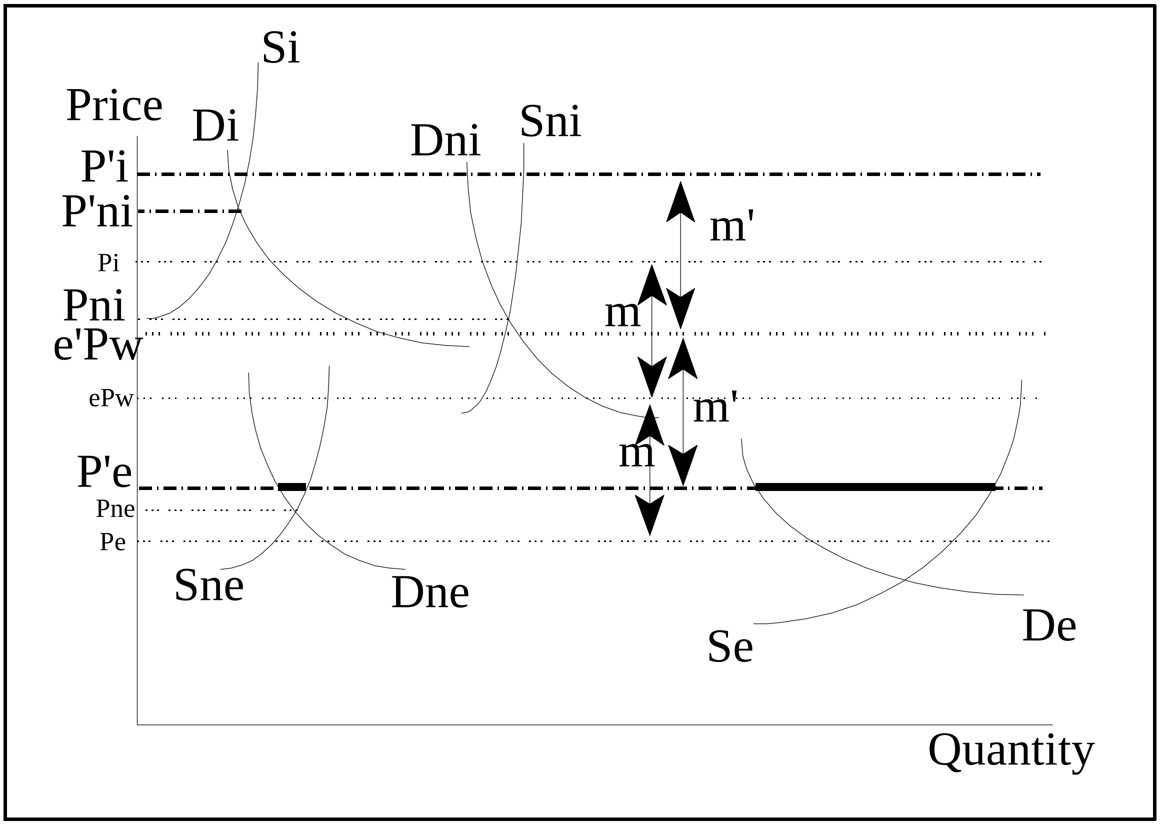

Figure 2: Commodity price equilibria after real depreciation

Table 1: Pre- and Post-Reform Commodity Subsector Structure

(Percent)	1984-1985				1989-1990			
Commodity	Coverage	EXP	IMP	NT	Coverage	EXP	IMP	NT
Dried beans	85	6	7	88	86	93	0	7
Maize	27	0	0	100	27	58	0	43
Manioc	72	0	0	100	72	31	0	69
Potatoes	51	100	0	0	51	100	0	0
Rice	100	0	74	27	100	0	19	81
Total	*62*	*8*	*32*	*60*	*65*	*20*	*11*	*69*

(Percent)	1984-1985				1989-1990			
Region	Cover	EXP	IMP	NT	Cover	EXP	IMP	NT
Imerina Centrale	83	19	36	44	85	48	0	52
Vakinankaratra	21	0	69	31	33	16	0	84
Itasy	81	39	27	34	81	33	0	68
Fianarantsoa	85	8	26	66	85	13	0	87
Mananjary	21	0	100	0	20	0	0	100
Farafangana	83	0	1	99	75	0	0	100
Toamasina	71	0	47	53	64	42	58	0
Ambatondrazaka	94	4	65	32	93	6	0	94
Fenerive Est	65	0	48	52	60	53	48	0
Mahajanga	74	1	70	29	78	18	82	0
Antsohihy	84	0	2	99	88	15	0	85
Maintirano	90	0	0	100	92	0	0	100
Toliary	67	0	28	72	67	59	0	42
Taolagnaro	9	0	0	100	16	0	0	100
Morondava	39	0	0	100	48	0	0	100
Antsiranana	45	0	74	26	48	0	85	16
Antalaha	63	0	1	99	72	0	0	100
Port regions [*]	*65*	*0*	*53*	*47*	*64*	*32*	*54*	*14*
Remote regions [**]	*56*	*0*	*1*	*99*	*57*	*0*	*0*	*100*
Total	*62*	*8*	*32*	*60*	*65*	*20*	*11*	*69*

[*] Port regions are Antsiranana, Mahajanga, Toamasina and Toliary.
[**] Remote regions are Antalaha, Farafangana, Maintirano, Morondava and Taolagnaro.

Commodity figures are output share weighted averages of regional series.
Regional and sectoral figures are revenue share weighted averages of commodity series.
Production and regional price data as reported by SMTIS/MPARA. International price data from IMF (maize, rice) and FAO (beans, manioc and potatoes).

Table 2: Distribution of Regional Series by Commodity-Specific Subsample

		EX ANTE		
		Importable	**Nontradable**	**Exportable**
	Importable	Rice (4)		
EX POST	**Nontradable**	Dried beans (5) Rice (7)	Maize (2) Manioc (7) Potatoes (2) Rice (6)	
	Exportable		Dried beans (2) Maize (4) Manioc (6) Potatoes (3)	Dried beans (2) Potatoes (5)

Number of regions in parentheses

Table 3: Results for Commodities Importable Both Before and After Reforms

Rice

Conditional Mean Equation*	Estimate	Std. Error	Conditional Variance Equation*	Estimate	Std. Error
P_{t-1}	0.80	(0.07)	Constant	15.75	(2.21)
REFORM	4.06	(2.02)	α	0.05	(0.10)
RER	0.23	(0.11)	ϕ	0.88	(0.15)
			REFORM	-40.77	(0.04)
			RER	0.72	(0.05)
n		432			
Log Likelihood		-2128			
R^2 **		0.72			

* Estimates for regional and strike dummies, buffer stock releases, and constant of conditional mean equation omitted.
** Between observed and predicted values.

Table 4: Results for Commodities
Importable Before and Nontradable After Reforms

	Dried Beans		Rice	
	Estimate	Std. Error	Estimate	Std. Error
Conditional Mean Equation*				
P_{t-1}	0.86	(0.06)	0.82	(0.07)
REFORM	53.13	(1.81)	-5.88	(1.59)
RER	0.88	(0.12)	0.78	(0.09)
pre-BPP	-0.02	(0.13)	-0.02	(0.21)
post-BPP	0.25	(0.26)	-0.07	(0.32)
Conditional Variance Equation*				
Constant	383.07	(4.71)	450.35	(2.62)
α	0.14	(0.19)	0.08	(0.12)
ϕ	0.38	(0.23)	0.81	(0.27)
RER	0.98	(0.31)	-0.43	(99.14)
REFORM	388.41	(0.02)	156.05	(0.15)
n	480		756	
Log Likelihood	-2384		-3647	
R^2 **	0.72		0.69	

* Estimates for regional and strike dummies, buffer stock releases, and constant of conditional mean equation omitted.
** Between observed and predicted values.

Table 5: Results for Commodities Nontradable Both Before and After Reforms

	Maize		Manioc	
	Estimate	Std. Error	Estimate	Std. Error
Conditional Mean Equation*				
P_{t-1}	0.75	(0.00)	0.84	(0.03)
REFORM	-103.09	(16.81)	37.06	(24.50)
RER	-0.01	(0.05)	-0.03	(0.02)
Conditional Variance Equation*				
Constant	1576.20	(10.44)	384.61	(41.91)
α	0.36	(0.09)	0.62	(0.13)
ϕ	0.14	(0.17)	0.41	(0.07)
RER	-1.99	(3.49)	-0.49	(0.72)
REFORM	61.36	(45.25)	98.22	(34.89)
n	192		672	
Log Likelihood	-877		-3167	
R^2 **	0.69		0.66	

	Potatoes		Rice	
	Estimate	Std. Error	Estimate	Std. Error
Conditional Mean Equation*				
P_{t-1}	0.89	(0.09)	0.66	(0.07)
REFORM	3.15	(2.23)	110.74	(331.64)
RER	0.01	(0.03)	0.03	(5.47)
Conditional Variance Equation*				
Constant	3397.10	(5.06)	1187.10	(88.39)
α	0.12	(0.16)	0.47	(0.29)
ϕ	0.81	(0.20)	0.56	(0.13)
RER	-6.16	(26.77)	-1.96	(7.93)
REFORM	74.86	(0.57)	200.82	(828.41)
n	192		648	
Log Likelihood	-1024		-3230	
R^2 **	0.77		0.75	

* Estimates for regional and strike dummies, buffer stock releases, and constant of conditional mean equation omitted.

** Between observed and predicted values.

Table 6: Results for Commodities
Nontradable Before and Exportable After Reforms

	Dried Beans		Maize	
	Estimate	Std. Error	Estimate	Std. Error
Conditional Mean Equation				
P_{t-1}	0.84	(0.06)	0.61	(0.09)
REFORM	73.51	(2.02)	16.31	(1.76)
RER	0.52	(0.11)	0.49	(0.09)
Pre-BPP	-0.01	(0.11)	-0.14	(0.36)
Post-BPP	0.34	(0.16)	0.25	(0.05)
Conditional Variance Equation				
Constant	841.99	(0.06)	2043.10	(6.28)
α	0.17	(31.62)	0.57	(0.36)
ϕ	0.61	(42.99)	0.10	(0.28)
RER	-0.27	(0.39)	-7.87	(1.90)
REFORM	-590.34	(0.06)	-408.86	(0.01)
n	384		384	
Log Likelihood	-1851		-1861	
R^2	0.68		0.57	

	Manioc		Potatoes	
	Estimate	Std. Error	Estimate	Std. Error
Conditional Mean Equation				
P_{t-1}	0.92	(0.02)	0.92	(0.08)
REFORM	2.54	(21.12)	-0.26	(1.49)
RER	0.21	(0.01)	0.90	(0.88)
Pre-BPP	0.10	(0.07)	-0.04	(0.13)
Post-BPP	0.27	(0.37)	0.66	(0.19)
Conditional Variance Equation				
Constant	206.04	(56.99)	43.78	(4.11)
α	0.80	(0.15)	0.39	(0.29)
ϕ	0.61	(42.99)	0.10	(0.28)
RER	-0.25	(0.10)	0.29	(37.48)
REFORM	29.37	(50.77)	-75.59	(0.02)
n	576		288	
Log Likelihood	-2505		-1327	
R^2	0.67		0.83	

Table 7: Results for Commodities Exportable Both Before and After Reforms

	Dried Beans Estimate	Dried Beans Std. Error	Potatoes Estimate	Potatoes Std. Error
Conditional Mean Equation*				
P_{t-1}	0.93	(0.00)	0.96	(0.02)
REFORM	85.26	(122.51)	2.45	(9.03)
RER	0.68	(0.12)	0.13	(0.01)
Conditional Variance Equation*				
Constant	580.36	(10.03)	91.65	(45.91)
α	0.42	(21.57)	0.32	(0.08)
ϕ	0.22	(0.39)	0.67	(0.05)
RER	0.40	(0.78)	1.49	(0.68)
REFORM	218.73	(10.03)	-162.53	(46.63)
n	192		480	
Log Likelihood	-926		-2119	
R^2 **	0.89		0.84	

* Estimates for regional and strike dummies, buffer stock releases, and constant of conditional mean equation omitted.
** Between observed and predicted values.