ELSEVIER

Agricultural Economics 20 (1999) 215–230

AGRICULTURAL ECONOMICS

# The effects of real exchange rate depreciation on stochastic producer prices in low-income agriculture

Christopher B. Barrett*

*Department of Agricultural Resource, and Managerial Economics, 351 Warren Hall, Cornell University, Ithaca NY 14853-7801, USA*

Received 24 January 1997; received in revised form 2 November 1998; accepted 4 January 1999

## Abstract

This paper considers the effects of real exchange rate depreciation on stochastic agricultural producer prices in low-income agriculture. Conventional wisdom, that real depreciation achieved through nominal currency devaluation stimulates tradables production, does not universally hold in the presence of stochastic prices. In fact, real depreciation is only stimulative in two cases – importables that remain importable and nontradables that become exportable. GARCH estimation of time-series price data on several commodities from Madagascar support the hypotheses generated by the analytical model. 

*JEL classification*: O1; Q1; F31

*Keywords:* Depreciation; Production under risk; GARCH estimation; Madagascar

## 1. Introduction

Considerable evidence exists that overvalued currencies have a strong dampening effect on agricultural output in developing economies (Jaeger and Humphreys, 1988; Krueger et al., 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. Conventional wisdom holds that this increases agricultural prices and thus stimulates agricultural production (Schuh, 1974; Dornbusch, 1988; Krueger et al., 1988). Yet the agricultural sectors of developing countries in Africa and Latin America have an exhibited uneven and generally weak supply response to considerable real exchange rate depreciation (Commander, 1989; Barrett and Carter, 1999), with the notable exception of nontraditional exports, which have flourished in the wake of real depreciation (Barham et al., 1992), and some importable foodstuffs. 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 these particular classes of products. The relationship between exchange rates and agricultural supply response thus appears more complex than

*Tel.: +1-607-255-4489; fax: +1-607-255-9984; e-mail: cbb2@cornell.edu; http://www.cals.cornell.edu/dept/arme/staff/cbb2

0169-5150/99/$ – see front matter 
PII: S0169-5150(99)00003-1

conventional wisdom admits. 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). Meanwhile, the impacts of recent, sharp exchange rate depreciation in southeast Asia on agricultural output are not yet clear. 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.

The plan of the paper is as follows. Section 2 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. Conventional wisdom, that real depreciation stimulates tradables production, does not hold universally when one admits price uncertainty and producer income risk aversion. In fact, real depreciation only yields a stimulative price signal in the cases of nontradables-turned-exportables – which I refer to as nontraditional exports[1] – and importables that do not become nontradable. In Section 3, 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, that is, 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, 1999). 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

[1]The reader should note that my definition of non-traditional exports does not include higher-value-added products (e.g. flowers, spices, fruits and vegetables) that were previously tradable but un- or under-developed. Rather, the term 'non-traditional exports' as used in this paper refers to products that were non-tradable but have become exportable following depreciation.

exchange rate. The autarkic conditional equilibrium price distribution for each commodity $i$ can be characterized by the density function $\delta_i$ ($y$, $u$), 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.[2] Each commodity is also traded in international markets, at an exogenous world price distribution, $\omega_i$, 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.[3] The equilibrium price distribution, $\pi^*$, is thus (eliminating indexing subscripts for the sake of clarity)

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

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 $s^*$

$$s^* = \begin{cases} s_i = s_w e & \text{iff} \quad p_n \geq p_w e + m \\ s_n & \text{iff} \quad p_w e - m \leq p_n \leq p_w e + m \\ s_e = s_w e & \text{iff} \quad p_n \leq p_w e - m \end{cases} \tag{3}$$

[2]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).

[3]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).

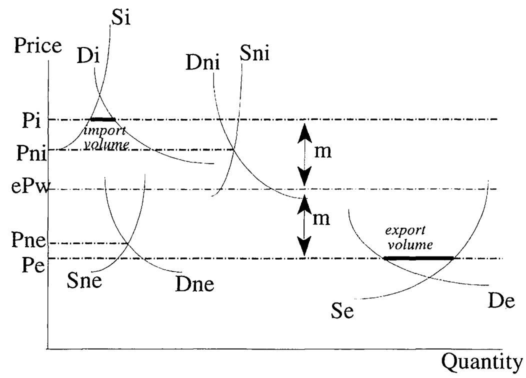


Fig. 1. Ex ante expected commodity price equilibria.

Assume that $s_n > s_e = s_i$ pre-liberalization, indicating that nontradables' prices are more volatile than tradables' prior to real depreciation.[4] 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.

Expected equilibrium prices are depicted in Fig. 1 for four different stylized commodities.[5] 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 mean price between the adjusted expected border parity

[4]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 price inelasticity of demand and supply for nontradable staple foods (Sahn and Delgado, 1989).

[5]Note that Figs. 1 and 2 do not depict the prevailing distributions, just the conditional means of those distributions

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 expected world market price ($p_i > p_{ni} > ep_w$), and near-exportables (represented by the $D_{ne}$ and $S_{ne}$ schedules) for which autarkic equilibrium prices are lower than the expected world market price ($p_e < p_{ne} > ep_w$). When agricultural commodities are presumed tradable and price risk is disregarded, depreciation[6] clearly stimulates agricultural production. Relaxation of those two strong assumptions, however, yields a more complex story.

Assume an Arrow–Pratt risk-averse firm maximizes the expected utility of profits, and assume for the moment only that products cannot shift between equilibrium pricing conditions (i.e. importables remain importables and nontradables remain nontradables). The effects of real depreciation can then be seen by differentiating the equilibrium price distribution reflected in Eq. (1) with respect to the exchange rate, $e$. In the case of importables, the result is strictly positive, $\partial\iota/\partial e = \omega + \partial m/\partial e > 0$, reflecting a rightward shift of the price distribution function. In other words, the post-depreciation price distribution first-order stochastically dominates the pre-depreciation distribution, so that the exchange rate shock is unambiguously stimulative. Producers face a higher probability of high prices and a lower probability of low prices. Note, however, that the assumption that the product remain importable holds only for 'small' depreciations, since $p_i$ is bounded from above by $p_n$. The stimulative effects of massive depreciation episodes – such as those that have taken place recently in southeast Asia or, over the past decade, in much of sub-Saharan Africa and Latin America – therefore remain open to question. We return to this point in a few paragraphs.

The case of nontradables is simple; exchange rate depreciation has no effect on incentives in partial equilibrium. If marketing costs are significant, as is common in poor economies, the price space occupied by nontradables may be large (Fig. 1) and this class of products unaffected by depreciation could represent a sizable share of the agricultural economy.

The effect on exportables' price distributions is complicated by the potential tradability of marketing inputs. In the highly unlikely case that there are no tradable inputs (e.g. no fuel), then the exportables' case and the importables' case are identical, $\partial\varepsilon/\partial e = \omega > 0$. In that case, a real depreciation is stimulative for all tradables because it generates a rightward shift in the price distribution. However, there is generally a substantial tradable component (e.g. fuel, spare parts, vehicles) to food marketing costs in most of the developing world (Ahmed and Rustagi, 1987), and food prices sometimes fall quite low. As a result, the post-depreciation exportables price distribution will not, in general, dominate the pre-depreciation distribution.[7] Indeed, for a sufficiently high degree of agent risk aversion, pre-depreciation distribution could be preferred.

It may be simpler to think of these results, and it will be easier to test this model empirically, in terms of the mean and variance of the equilibrium price distributions. Differentiation of Eqs. (2) and (3) with respect to $e$ captures the essence of the above results. As shown in Eq. (4)–Eq. (7), and assuming still that commodities are not allowed to shift among the three equilibrium conditions, real depreciation increases the mean and variance of importables prices, the variance of exportables prices, has no effect on either moment of the nontradables' price distribution, and has analytically ambiguous effects on the mean of the exportables' price distribution.

$$\partial p_i/\partial e = p_w + \partial m/\partial e \quad (4)$$

$$\partial p_e/\partial e = p_w - \partial m/\partial e \quad (5)$$

$$\partial s_i/\partial e = \partial s_e/\partial e = s_w \quad (6)$$

$$\partial p_n/\partial e = \partial s_n/\partial e = 0 \quad (7)$$

[6] 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, $\delta$ would no longer be independent of $e$.

[7] Formally, the post-depreciation exportables price distribution stochastically dominates (in any degree) the pre-depreciation exportables price distribution if and only if $\inf(\omega) > \partial m/\partial e$, i.e., if the minimum price of the world distribution exceeds the induced increase in the marketing margin. Intuitively, this will not hold for large depreciation episodes, or for crops occasionally drawing very low prices or whose marketing margins are large and primarily comprised of tradables

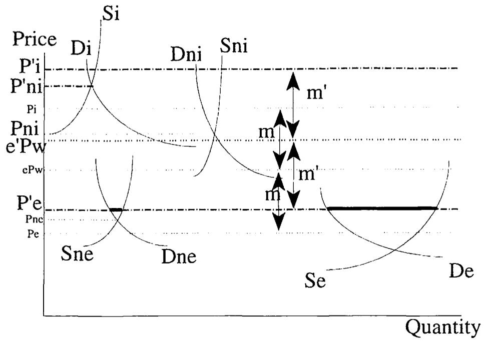


Fig. 2. Expected commodity price equilibria after real depreciation.

In practice, the circumstances under which expected exportables' prices would decrease are somewhat unusual; tradable marketing costs ex ante would have to exceed the expected world price for $\partial m/\partial e > p_w$. Nonetheless, for the reasons enumerated above, the induced change in the exportables' price distribution is not necessarily preferred by producers, so that the incentive effects of real depreciation are not unambiguously stimulative for exportable products.

The key difference between the effects of depreciation on exportables and importables arises from the effects of real depreciation on transfer costs, captured in the second term in equations Eqs. (4) and (5). Since 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 expected border parity 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 (Fig. 2). If commodities could not shift between equilibrium conditions – exportable, importable, and nontradable – then real exchange rate depreciation would be unambiguously stimulative to importables, with no effect on nontradables and ambiguous effects on exportables. These asymmetric incentive effects on tradables' price distributions may help explain why the external adjustment of African agricultural trade balances has fallen disproportionately on the import side of the current account.

Now relax the restriction that commodities cannot shift between categories. In Fig. 2, real depreciation also renders one previously (near-exportable) nontradable commodity exportable, while the commodity that was importable becomes nontradable. The endogenous contraction of importables' price space and corresponding expansion of the price space occupied by exportables and nontradables is likely to induce some commodities to switch equilibrium pricing conditions, with importables becoming nontradable and near-exportables becoming exportable. Under the above assumption that $s_n > s_e$, once depreciation induces $p_w e - m \geq p_n$, the post-depreciation (exportable) price distribution will stochastically dominate the pre-depreciation (nontradable) distribution. Real depreciation benefits near-exportables' producers by bringing rising expected prices and falling price variance if real depreciation bumps $e$ above $p_{ne}$. The price distribution of importables-turned-nontradable, by contrast, does not necessarily dominate the pre-distribution importables price because although it yields a nonnegative shift in the expected price, it also brings increased price variability, including potentially greater probability of low prices than previously.[8]

The key message of the above analysis is that the impact of real exchange rate depreciation on the price distributions faced by agricultural producers is conditional on commodities' ex ante structural position vis-à-vis the world market and on the size of the depreciation. 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. Nonetheless, 'small' depreciation episodes will be unambiguously stimulative, so long as the commodity remains importable. Predepreciation exportables experience increasing variance and analytically ambiguous effects on mean price and on the production stimulus effects of the shock. Nontradables are not affected by nominal exchange rate changes that cause real depreciation unless the shift is sufficient to render them internationally tradable, in which case the induced shift – higher mean, lower variance – is unambiguously stimulative for this 'nontraditional

[8] Formally, given the assumptions that $p_i < p_n$ and $s_i < s_n$ for importables, the post-depreciation nontradables' price distribution stochastically dominates (in any degree) the pre-depreciation importables price distribution if and only if $\inf(\iota) < \inf(\delta)$, i.e., if nontradables prices never fall lower than importables prices, a condition unlikely to hold.

exports' class.[9] The evolution of agricultural price distributions following real depreciation designed to stimulate agricultural production is clearly heterogeneous across subsectors. Consequently, there can be considerable intra-sectoral variation in supply response without different price elasticities of supply, and aggregate supply response to real depreciation can vary considerably across economies of different structural characteristics.[10]

## 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 (decrease) the proportion of crops that are exportable (importable). Second, one should find a structural shift in the correlation between the border parity price and local market prices where depreciation induces a shift among 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, to (from) one if the law of one price holds strictly. Third, the effects of reforms on the mean and variance of agricultural price series, as identified in Eq. (4)–Eq. (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, 1983–1991, for five commodities (dried beans, maize, manioc, rice, and potatoes) from 17 different regions in Madagascar. The period 1983–1985 is used as the pre-reform 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), 1986–1988. Allowing for full adjustment, given contractual and informational lags, 1989–1990 is used as the post-reform 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 regional commodity price series. Distance-, season-, and time-varying transfer costs per ton were modelled using equation (Eq. (8)).[11]

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

Transfer costs per ton, $m$, are modelled as the sum of transport and other variable handling expenses.[12] For want of reliable data with which to make cross-regional or inter-temporal 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[13] are divided between all-season macadam routes ($\mathrm{km_m}$) and dirt roads ($\mathrm{km_d}$), since costs on the latter average thrice those on the former.[14] 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). As seasonal effects are more pronounced

[9] 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 et al., 1992).

[10] 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.

[11] Costs are assumed constant across commodities.

[12] 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).

[13] The four major international ports (Antsiranana, Mahajanga, Toamasina, and Toliary) accounted for more than 80 % of gross international shipping volume to or from Madagascar, 1980–1986, and were the only ports to handle at least 100 000 tons per year (BDE, 1988).

[14] 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).

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 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.[15] Estimated real transfer costs in Madagascar increased more than 15% in U.S. dollar terms between 1984–1985 and 1989–1990, 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 world 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.[16]

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 post-reform periods.[17] 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 1984–1985 and 1989–1990 on the strength of export volume increases of 8530, 575, and 4933% 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 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.[18]

[15] 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.

[16] Unnevehr (1984) employed a similar method in a study of cassava in Indonesia. These qualitative decisions were corroborated by calculating the 1984–1985 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.

[17] 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.

[18] 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 et al., 1990).

Table 1
Pre- and post-reform commodity subsector structure

| Commodity (%) | 1984–1985 | | | | 1989–1990 | | | |
|---|---|---|---|---|---|---|---|---|
| | 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* |
| *Region (%)* | 1984–1985 | | | | 1989–1990 | | | |
| 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*[a] | *65* | *0* | *53* | *47* | *64* | *32* | *54* | *14* |
| *Remote regions*[b] | *56* | *0* | *1* | *99* | *57* | *0* | *0* | *100* |
| *Total* | *62* | *8* | *32* | *60* | *65* | *20* | *11* | *69* |

[a] Port regions are Antsiranana, Mahajanga, Toamasina and Toliary.
[b] 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).

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 to be 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 post-depreciation tradability conditions. Table 2 shows the distribution of region- and commodity-specific price series across the sub-

Table 2
Distribution of regional series by a commodity-specific subsample

| EX ANTE | | Importable | Nontradable | Exportable |
|---|---|---|---|---|
| | Importable | Rice (4) | | |
| EX POST | Nontradable | Dried beans (5)<br>Rice (7) | Maize (2)<br>Manioc (7)<br>Potatoes (2)<br>Rice (6) | |
| | Exportable | | Dried beans (2)<br>Maize (4)<br>Manioc (6)<br>Potatoes (3) | Dried beans (2)<br>Potatoes (5) |

Number of regions in parentheses.

samples, by ex ante and ex post tradability status.[19] Since the objective of the 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

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

where $\beta(L)$ is a polynomial lag operator, $\beta_0$ and $\alpha_0$ 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[20] and the autocorrelation series always dampened to 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.[21]

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[22]

[19]Depreciation implies a lower triangular matrix (Table 2), while appreciation implies an upper triangular matrix. Recall from the previous section that of the six cells resulting from a depreciation epsiode, only the upper left and the lower center ones yield unambiguously stimulative price distribution shocks.

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

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

[22]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.

Table 3
Results for commodities importable both before and after reforms

Rice

| Conditional mean equation[a] | Estimate | SE | Conditional variance equation[a] | Estimate | SE |
|---|---|---|---|---|---|
| $P_{t-1}$ | 0.80 | (0.07) | Constant | 15.75 | (2.21) |
| REFORM | 4.06 | (2.02) | $\alpha$ | 0.05 | (0.10) |
| RER | 0.23 | (0.11) | $\varphi$ | 0.88 | (0.15) |
| | | | REFORM | −40.77 | (0.04) |
| | | | RER | 0.72 | (0.05) |
| $n$ | 432 | | | | |
| Log-likelihood | −2128 | | | | |
| $R^2$ [b] | 0.72 | | | | |

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

– was multiplied by the reform dummy variable and by its complement to create two variables: the pre-reform border parity price (pre-BPP), and the post-reform 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 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.[23]

The results for the coefficient estimates of interest are displayed in Tables 3–7; 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 2 predicts that both the mean and conditional variance increase in RER. The positive and statistically significant[24] 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 1983–1985 period mean, are 0.46 and 0.17, respectively. Given better than 50% real depreciation, 1986–1988, the estimated real effects on the mean and variance of producer prices were considerable. As shown in Table 2, rice was the only one of these commodities to remain importable in the wake of real exchange rate depreciation. Rice output increase considerably faster than that of the other crops. The effects hypothesized in this paper may be one of several explanations for that outcome.[25]

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 variance of prices' increase with RER and that domestic prices were positively correlated with border parity prices before devaluation, but not

[23]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, 1975). As 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.

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

[25]Barrett (1998) posits a different reason related to this some observed change in price distributions.

Table 4
Results for commodities importable before and nontradable after reforms

| | Dried beans | | Rice | |
|---|---|---|---|---|
| | Estimate | SE | Estimate | SE |
| *Conditional mean equation*[a] | | | | |
| $P_{t-1}$ | 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*[a] | | | | |
| Constant | 383.07 | (4.71) | 450.35 | (2.62) |
| $\alpha$ | 0.14 | (0.19) | 0.08 | (0.12) |
| $\varphi$ | 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 | |
| $R^2$ [b] | 0.72 | | 0.69 | |

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

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 pre-reform 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 1983–1991 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

Table 5
Results for commodities nontradable both before and after reforms

| | Maize | | Manioc | |
|---|---|---|---|---|
| | Estimate | SE | Estimate | SE |
| *Conditional mean equation*[a] | | | | |
| $P_{t-1}$ | 0.75 | (0.08) | 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*[a] | | | | |
| Constant | 1576.20 | (10.44) | 384.61 | (41.91) |
| $\alpha$ | 0.36 | (0.09) | 0.62 | (0.13) |
| $\varphi$ | 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 | |
| $R^2$ [b] | 0.69 | | 0.66 | |
| | Potatoes | | Rice | |
| *Conditional mean equation*[a] | | | | |
| $P_{t-1}$ | 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*[a] | | | | |
| Constant | 3397.10 | (5.06) | 1187.10 | (88.39) |
| $\alpha$ | 0.12 | (0.16) | 0.47 | (0.29) |
| $\varphi$ | 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 | |
| $R^2$ [b] | 0.77 | | 0.75 | |

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

between RER and mean price, and a negative relation between RER and 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.[26] Most of the

[26] 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 world potato price was 0.31.

Table 6
Results for commodities nontradable before and exportable after reforms

| | Dried beans | | Maize | |
|---|---|---|---|---|
| | Estimate | SE | Estimate | SE |
| *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) |
| $\alpha$ | 0.17 | (31.62) | 0.57 | (0.36) |
| $\varphi$ | 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 | |
| | Potatoes | | Rice | |
| *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) |
| $\alpha$ | 0.80 | (0.15) | 0.39 | (0.29) |
| $\varphi$ | 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 | |

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 1984–1985 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 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

Table 7
Results for commodities exportable both before and after reforms

| | Dried beans | | Potatoes | |
|---|---|---|---|---|
| | Estimate | SE | Estimate | SE |
| *Conditional mean equation*[a] | | | | |
| $P_{t-1}$ | 0.93 | (0.08) | 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*[a] | | | | |
| Constant | 580.36 | (10.03) | 91.65 | (45.91) |
| $\alpha$ | 0.42 | (21.57) | 0.32 | (0.08) |
| $\varphi$ | 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 | |
| $R^2$ [b] | 0.89 | | 0.84 | |

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

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. 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 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 3–7 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, and one each for maize and manioc. Each of the 11 test statistics far exceeded the relevant $\chi^2$ 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 many risk-averse tradables producers. For small depreciation episodes, in which importables do not become nontradable, exchange rate shocks will stimulate importables production. However, massive real currency depreciation, of the sort common in structural adjustment programs and during emerging markets' financial crises in the past 15 years, favors nontraditional exports over all other commodity classes, in that only commodities moving from nontradability to exportability experience a stochastically dominant shift in price distributions – evidenced by rising mean prices and falling price variability – regardless of the size of the depreciation shock. These observations may help account for two empirical regularities of the past 15 years' adjustment experience. First, import compression has accounted for a far larger share of agricultural current account adjustments than has export expansion. Second, within the exportables sector, nontraditional exports have a had disproportionately robust response to real exchange rate depreciation, as exemplified by the cases of 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), 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 Fig. 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, 1999), 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).[27] 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).

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 (Jaeger and Humphreys, 1988; Krueger et al., 1988; Elbadawi, 1992; Ghura and Grennes, 1993). As a consequence, many analysts and policy makers expected real depreciation to have expansionary effects for agriculture. But those expectations have been based on assumptions of full tradability and certain prices that are untenable in low-income agriculture. This paper makes the case that where a nontrivial portion of the agricultural sector is nontradable and marketing costs are partly tradable, one should expect considerable cross-sectional variation in the response of stochastic prices to real exchange rate depreciation. Thus, the aggregate sectoral effects on the incentives faced by risk-averse producers are

[27] Although Lapp and Smith (1992) found that 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.

inherently ambiguous, depending fundamentally on the ex ante structure of the sector.

## Acknowledgements

I thank Michael Carter, Jean-Paul Chavas, Paul Dorosh, Stan Johnson, John Rust, an anonymous referee, and seminar participants at UC-Davis, Wisconsin, and Utah State for helpful comments. This research was supported by the Institute for the Study of World Politics, and the Social Science Research Council. The usual disclaimer applies.

## References

Ahmed, R., Rustagi, N., 1987. Marketing and price incentives in African and Asian countries: A comparison, In: Elz, D. (Ed.), Agricultural Marketing Strategy and Pricing Policy, World Bank, Washington, DC.

Balassa, B., 1990. Incentive policies and export performance in sub-Saharan Africa. World Development 18, 383–391.

Banque des Donnees de l'Etat, 1988. Inventaire Socio-Economique, 2 vols. Antananarivo.

Barham, B.L., Clark, M., Katz, E., Schurman, R., 1992. Nontraditional agricultural exports in Latin America: Towards an appraisal. Latin American Research Review 27, 43–82.

Barrett, C.B., 1998. Immiserized growth in liberalized agriculture. World Development 26, 743–753.

Barrett, C.B., Carter, M.R., 1999. Economies. Microeconomically coherent agricultural policy reform in Africa, In: Paulson, J. (Ed.), African Economies in Transition, Macmillan, London.

Chavas, J.-P., 1988. On competitive speculation under uncertainty: An alternative view of the inverse-carrying charge. Journal of Economics and Business 40, 117–128.

Commander, S. (Ed.), 1989. Structural Adjustment and Agriculture: Theory and Practice in Africa and Latin America, Heinemann, Portsmouth, NH.

Delgado, C.L., 1992. Why domestic food prices matter to growth strategy in semi-open West African agriculture. Journal of African Economies 1, 446–471.

Diakosavvas, D., Kirkpatrick, C., 1990. Exchange-rate policy and agricultural exports performance in sub-Saharan Africa. Development Policy Review 8, 29–42.

Dornbusch, R., 1988. Overvaluation and trade balance, In: Dornbusch, R., Helmers, F.L.C.H. (Eds.), The Open Economy, Oxford University Press, Oxford.

Dorosh, P.A., 1994. Structural Adjustment, Growth, and Poverty in Madagascar: A CGE Analysis, Cornell University Food and Nutrition Policy Program monograph 17.

Elbadawi, I.A., 1992. Real overvaluation, terms of trade shocks, and the cost to agriculture in sub-Saharan Africa: The case of the Sudan. Journal of African Economies 1, 59–85.

Fischer, S., 1981. Relative Shocks, Relative Price Variability, and Inflation, Brookings Papers on Economic Activity, pp. 381–431.

Food and Agriculture Organization of the United Nations, Quarterly Bulletin of Statistics, various issues.

Ghura, D., Grennes, T.J., 1993. The real exchange rate and macroeconomic performance in sub-Saharan Africa. Journal of Development Economics 42, 155–174.

Guillaumont, P., 1992. Adjustment policy and agricultural development, In: Peters, G.H., Stanton, B.F. (Eds.), Sustainable Agricultural Development: The Role of International Cooperation, Proceedings of the 21st International Conference of Agricultural Economists, Dartmouth Publishing, Brookfield, VT.

International Monetary Fund, International Financial Statistics, Various issues.

Jaeger, W., Humphreys, C., 1988. The effect of policy reforms on agricultural incentives in sub-Saharan Africa. American Journal of Agricultural Economics 70, 1036–1043.

Krueger, A.O., Schiff, M., Valdés, A., 1988. Agricultural incentives in developing countries: Measuring the effects of sectoral and economywide policies. World Bank Economic Review 2, 255–271.

Kyle, S., 1992. Pitfalls in the measurement of real exchange rate effects on agriculture. World Development 20, 1009–1019.

Lapp, J.S., Smith, V.H., 1992. Aggregate sources of relative price variability among agricultural commodities. American Journal of Agricultural Economics 74, 1–9.

Meyer, J., 1987. Two-moment decision models and expected utility maximization. American Economic Review 77, 421–430.

Mundlak, Y., Cavallo, D., Domenech, R., 1990. Effects of macroeconomic policies on sectoral prices. World Bank Economic Review 4, 55–79.

Sahn, D.E., Delgado, C.L., 1989. The nature and implications for market interventions of seasonal food price variability, In: Sahn, D.E. (Ed.), Seasonal Variability in Third World Agriculture, Johns Hopkins University Press, Baltimore, MD.

Samuelson, P.A., 1952. Spatial price equilibrium and linear programming. American Economic Review 42, 283–303.

Schuh, G.E., 1974. The exchange rate and U.S. agriculture. American Journal of Agricultural Economics 56, 1–13.

Takayama, T., Judge, G.G., 1971. Spatial and Temporal Price and Allocation Models, North-Holland, Amsterdam.

United Nations Development Programme and Ministère De L'Economie Et Du Plan, 1991. Régions Et Développement: Etudes Régionales, 6 vols. Antananarivo.

Unnevehr, L.J., 1984. Transport costs, tariffs and the influence of world markets on Indonesian domestic cassava prices. Bulletin of Indonesian Economic Studies 20, 30–43.

Walsh, G.R., 1975. Methods of Optimization, Wiley, New York.

World Bank, 1991. Republic of Madagascar: Rural Road Sub-Sector Strategy, World Bank, Washington, DC.

World Bank. 1994. Adjustment in Africa: Reforms, Results, and the Road Ahead, Oxford University Press, Oxford.