

# Farmers' Exposure to Price Shocks in Deregulated Markets

Julie Subervie\*

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## Abstract

Post-reform markets in commodity-exporting countries in Africa and Latin America are generally characterized by the withdrawal of government organizations from commodity marketing, the emergence of new private agents operating as traders and exporters, and farmers' participation in the newly-reconfigured market regulatory authorities. However, the liberalization process varied across countries and resulted in the emergence of various marketing systems. This empirical analysis aims at providing better understanding of the situation faced by farmers likely to be exposed to both high world price volatility and possible market power of private agents. Focusing on three coffee exporting countries which have experienced market liberalization in the mid-1990s but display very different marketing systems today, I show that asymmetric adjustments of producer prices that characterized the pre-reform period and were favorable to farmers have disappeared in the post-reform period. Moreover, in spite of farmers' participation in market authorities, largest decreases in world prices may be transmitted relatively quickly to farmers when the number of exporters remains small.

Keywords: Developing countries, Market reforms, Coffee, Traders, Exporters, Price transmission, Structural break, Asymmetric adjustments

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\*INRA, UMR MOISA, 2 place Pierre Viala, 34060 Montpellier Cedex 2, France.  
julie.subervie@supagro.inra.fr

# 1 Introduction

Until the liberalization reforms, some of the stated objectives of governments of agricultural commodity exporting countries were to reduce producers' price variability and to protect farmers from possible monopolistic practices of private middlemen. Afterwards, regarding the difficulties faced by marketing boards, the objective of price stabilization has been withdrawn and the privatization of agricultural trade has been advocated as a way to ensure that farmers get the "right price". Thus, most developing countries in Sub-Saharan Africa and Latin America implemented structural adjustment reforms, which included the liberalization of export crop markets and the abolition of marketing boards, and allowed private agents to operate as traders and exporters. Earlier evidence suggests that in cases where interventions were greatest and reforms most complete, producers have indeed benefited from receiving a larger share of export prices, as producer prices started to follow more closely world prices in a quasi-mechanical way<sup>1</sup>. However, the reforms also implied the abolition of price stabilization schemes, exposing producers to the full volatility of markets. Moreover, the liberalization process has varied across countries, both in the scope of the reforms and their consequences, and has resulted in the emergence of various marketing systems (Akiyama, Baffes, Larson, and Varangis, 2003). For example, studies of the impact of liberalization on agricultural markets in Africa have shown that post-reform markets may be characterized by few large traders (Fafchamps and Hill, 2008). Such basic statistical evidence support the idea of a possible market power of new private agents, but to our knowledge, there is not any empirical evidence of such phenomenon<sup>2</sup>.

As a matter of fact, the volatility of world agricultural commodity prices has been higher over the past three decades than during the pre-1973 period (Dehn, Gilbert, and Varangis, 2005) while producers' ability to deal with the consequences of inter-annual price volatility has remained very limited (Varangis, Larson, and Anderson, 2002). The effects of commodity market reforms on producers therefore can not be sum-up to higher average prices. And yet, rather few studies tackled the issue of farmers' exposure to world price volatility after the reforms<sup>3</sup>. If the increasing variability of producers' price can be directly observed over the post-reform period, it is more difficult to put into light manifestations of market failure such as the exercise of market power by monopolistic middlemen against farmers (if there is any). This paper throws some new light on the issue of farmers' exposure to price shocks in deregulated markets, by focusing on the dynamics of price transmission, which can be seen as one of the principal external expression of market power.

The present analysis of the relationship between world coffee prices and producer prices uses monthly data series from the International Coffee Organization (ICO) database from 1975:1 to 2007:12 in three coffee exporting countries for which uninterrupted price series are available: El Salvador, Colombia and India. All of them have experienced the liberalization of coffee trade but with very different results in the new marketing chain structure. Roughly speaking, in India, liberalization has resulted in the emergence of a more competitive system, including numerous traders, middlemen and exporters with a rather limited market concentration (ITC 2008). Whereas in El Salvador and Colombia, the marketing system appear less competitive, despite the fact that farmers are well represented in the marketing regulatory authorities.

The empirical analysis includes several steps. First, the date of the reforms is determined by applying a breakpoint test to the cointegrating relationship between producer

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<sup>1</sup>As long as the benefits of increased price share were not offset by a fall in international prices.

<sup>2</sup>An exception being the analysis of Wilcox and Abbott (2004)

<sup>3</sup>Some analyses have considered the implications of structural adjustment for producers' profitability but little attention has been paid to price volatility issues (see Akiyama (2003) for a recent review).

prices and world prices. Second, I test the hypothesis of a closer cointegrating relationship after the breakpoint. Then, using a standard error correction model, I test the hypothesis of both higher short-run transmission and higher speed of transmission after the breakpoint. Fourth, I use recently developed threshold cointegration tests (usually used in analyses of spatial integration or analyses of the relationship between short-term and long-term interest rates) that allow for asymmetric adjustment towards a long-run equilibrium relationship, with a view to detecting favorable pricing policy over the pre-reform period and/or unfavorable influence of new private agents over the post-reform period.

The preliminary findings indicate that the abolition of stabilization schemes has induced a closer cointegrating relationship between producer prices and world prices in each of the countries studied. The direct impact of monthly variations in world prices on producer price variations has also increased, whereas the speed of adjustment of producer prices has not increased systematically. The central result of the paper is that the asymmetric adjustments that characterized the pre-reform period and were favorable to producers (large deviations from the long-run equilibrium resulting from increases in world prices being eliminated relatively quickly) have disappeared in the post-reform period. Moreover, in some cases the results suggest that largest decreases in world prices may be transmitted relatively quickly to farmers over the post-reform period.

The rest of the paper is organized as follows. Section 2 contains an overview of the main findings of the empirical literature on world price transmission and consequences of deregulation on producers. Models that can be used to test the hypotheses relative to the impact of deregulation on producer price adjustments are presented in Section 3. The results of the empirical analysis are shown in Section 4. Section 5 presents some concluding remarks.

## **2 Liberalization, price transmission, and market power in commodity markets**

The issue of farmers' exposure to price shocks in deregulated markets combines at least three sets of empirical literature: studies of world price transmission to farmers, studies of market power in agricultural markets, and studies of liberalization's impact on producer price variability. Some studies have analyzed the implications of structural adjustment for producers' profitability (Morales (1991), Upton (1993), Baffes and Gautam (1996), Akiyama, Baffes, Larson, and Varangis (2003)) but little attention has been paid to the price volatility issue. One exception is the analysis of Rapsomanikis and Sarris (2006). They study the impact of commodity price volatility on farmers' income in Ghana, Peru, and, Vietnam, relying on explicit formulas for income variance. They estimate income uncertainty generated by price and production volatility when farmers are exposed to the full volatility of world markets. Their results underline that market and non market uncertainties (price variations and unstable weather patterns respectively) significantly affect the variability of agricultural income of households in these countries, and especially households that are specialized in a few commodities.

*World price transmission to farmers* Evidence of the relationship between producer prices and world prices in empirical studies is scarce. Mundlak and Larson (1992) estimated a direct relationship between producer and world prices over the pre-reform period (1968-1978). They use annual averages of producer prices from the FAO database and proxied export unit values. They approximate the relationship between world prices and producer prices with a linear regression, where prices are in logarithms. Results of the estimated trans-

mission in a cross-country comparison led to surprisingly high transmission elasticities<sup>4</sup> suggesting that the commodity-pooling procedure hid some inconsistency in the data. On another hand, the analysis for individual commodities (wheat, coffee, and cocoa) yielded somewhat lower values, closer to what would be expected in countries where policy often aimed to have some smoothing effect. However, in some cases like coffee in Colombia and El Salvador, elasticity remains puzzling<sup>5</sup> (0.62 and 1.05, respectively). Other studies display more reasonable results. For example, Conforti (2004) investigate the transmission between producer prices and world prices using a cointegration framework, for a large number of country/commodity pairs over the recent decades. Results for African countries generally tend to show a lower degree of price transmission compared to that of other countries. However, information available in the data set being sometimes limited, interpretation of the results remains unclear.

There are also few assessments of commodity market reforms showing a structural break in pricing regimes. Baffes and Gardner (2003) have examined the degree to which world price signals have been transmitted to producer prices, using a more dynamic framework which takes into account the non-stationarity of the series. Annual data from various sources, covering the period from 1970 to the mid-1990s, for eight countries and ten commodities, giving a total of 31 country/commodity pairs, were used in the study. A dynamic model<sup>6</sup> was estimated, by allowing for structural breaks in the years in which description of each country's reforms suggested they were likely to begin to have observable market effects. Evidence that policy reforms have reduced distortions in their domestic commodity price as compared to world prices is mixed. A structural break was identified in only 11 of the 31 commodity/country cases<sup>7</sup>. Moreover, only 7 of the 31 cases have a measured nominal rate of protection closer to zero after the reforms than before. Such results surprisingly suggest that the political intervention to insulate domestic markets from world commodity markets is persistent in most of the countries investigated. Nevertheless, these findings rely only on annual data covering relatively short sub-periods, thus making the interpretation of a dynamic specification quite difficult. Focusing on a group of coffee-producing countries, Krivonos (2004) shows that the reforms generally induced a closer cointegrating relationship between grower prices and world market prices, which is as expected in cases when stabilization schemes have been withdrawn. Results further show that short-run transmission of price signals from the world market to domestic producers has improved, such that domestic prices adjust faster today to world price fluctuations than they did prior to the reforms. Other studies underline that even if the short-run transmission of price signals from world to domestic markets has improved, it has remained weak in some cases (Worako, van Schalkwyk, Alemu, and Ayele, 2008).

*Market power in agricultural markets* Several studies seek evidence regarding the allega-

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<sup>4</sup>For example, the estimated transmission elasticity from pooled-commodity regressions equals 0.9 in Colombia, 0.9 in El Salvador and 0.8 in India. Within-commodity regressions yield lower though still high values (0.65 in Colombia, 0.76 in El Salvador and 0.44 in India).

<sup>5</sup>Such results may be explained by a potentially inappropriate log transformation of the variables, which often results in considerably higher coefficients than regression with original raw values. The contribution of world prices to variations in producer prices (given by the value of the  $R^2$  of the regression) may also appear higher in regressions in logarithms (0.95 and 0.93 in Colombia and El Salvador, respectively). Moreover, statistical properties of the series may give misleading results.

<sup>6</sup>Baffes and Gardner (2003) used a modified error correction model, including the difference between the world price and the producer price instead of the so-called error correction term.

<sup>7</sup>The authors tested for a structural break induced by policy reforms using a test on the parameter  $k$ , defined as the amount of adjustment which takes place in  $n$  periods:  $k = 1 - (1 - \beta)(1 - \alpha)^n$ .

tions of growing market power in the coffee roasting and retailing sector through the 1990s (e.g. Mehta and Chavas, 2008 ; Shepherd, 2004). They find that price transmission to the retail sector is asymmetric, with retail prices more responsive to increases than decreases. Such asymmetric price transmission at the retail level could thus help roasters and retailers benefit from upstream price interventions. On another hand, less attention has been paid to the influence on price transmission to producers of the emergence of new agents (traders and exporters) on post-reform markets. Two exceptions are Fafchamps and Hill (2008) and Wilcox and Abbott (2004). Focusing on the coffee value chain in Uganda over the post-reform period, Fafchamps and Hill (2008) find that a rise in the international price is readily reflected in export and wholesale prices, down to the first processing stage, but that growers receive a smaller share of the international price when it rises. In other words, when the international price rises, all domestic prices follow except for the price paid to producers, which rises much less. Their results further show that such phenomenon can be explained by the entry of traders who take advantage of farmers' ignorance about the rise in wholesale price.<sup>8</sup> Focusing on the case of West African cocoa market liberalization, Wilcox and Abbott (2004) studied the emergence of multinational processing firms which take over exporting activities and may collect rents previously captured as export taxes. They estimate the degree of market power into post-reform markets of Ivory Coast and Nigeria using a conjectural variation approach and find evidence of a market power exercised by multinational exporters and processors against farmers in the case of Ivory Coast but not in the case of Nigeria. Finally, some authors have attempted to detect the expression of market power into the dynamics of the relationship between world prices and producer prices (Krivonos, 2004 ; Worako, 2008) . Their analyses rely on modified error correction models where the short term impact of positive variations in world prices differs from the short term impact of negative ones. However, such models failed to support the hypothesis of exporters' market power. Finally, in a very original study, Coe (2006) shows that the commodity market liberalization and the privatization of market regulatory institutions have modified farmers bargaining power within the marketing regulatory authorities since the 1990s. Interestingly, he shows that the farmer share of the world price is higher in countries where farmer groups participate in the country's coffee authority, but tells nothing about shock transmission to farmers.

### 3 Modeling asymmetries in world price transmission to producers

Asymmetric price transmission has received much attention in agricultural economics (see Meyer and von Cramon-Taubadel (2004) for a recent survey). Many papers have focused on asymmetric price transmission between different stages of the marketing chain (vertical transmission) or between different locations for the same product (spatial transmission) but less attention has been paid to the possible existence of asymmetric price transmission in a framework where producer prices depend on world prices. In the vertical or spatial literature, most papers refer to non-competitive market structures and adjustment costs as an explanation for asymmetry. For example, in a commonly-used framework where retail prices are assumed to depend on farm prices, it is expected that imperfect competition in processing and retailing allows middlemen to use market power, which results in a so-called asymmetric price transmission: increases in farm prices (which squeeze middlemen's

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<sup>8</sup>The authors explain that traders' profits generated by exploiting farmers' ignorance are not eliminated by competition because excess entry of traders increases the search time of traders buying directly from producers scattered over a large area.

margins) are transmitted faster and/or more completely to consumers than are decreases (which stretch middlemen's margins). Nevertheless, some authors have suggested that market power can lead to negative asymmetric price transmission as well, if oligopolists are reluctant to risk losing their market share by increasing retail prices (Ward, 1982). On account of the statistical properties of the series, recent studies of asymmetric price transmission come within the framework of cointegration analysis. In such a framework, authors consider a possible asymmetry with respect to the speed of adjustment. For example, in a spatial framework where wholesale prices in local markets are assumed to depend on central market prices (Badiane and Shively (1998) Abdulai (2000)), the local prices are expected to adjust faster to deviations from the long-run equilibrium resulting from increases in central market prices, than to deviations resulting from decreases in those prices. Moreover, many authors aim at showing that the speed of adjustment will differ according to whether the deviation from the long-run equilibrium exceeds some specific threshold levels (Obstfeld and Taylor (1997), Balke and Fomby (1997), Goodwin and Holt (1999), Goodwin and Piggott (2001)).

In our framework, where producer prices in crop-exporting countries are driven by world prices, possible explanations for asymmetry in the speed of adjustment of producer prices strongly depend on the considered time period. Over the pre-reform period, government intervention in the form of administered producer prices may lead to positive asymmetric price transmission, in the sense that producer prices may respond faster to deviations from the long-run equilibrium resulting from world price increases. The hypothesis of a situation so favorable to producers under the pre-reform period is supported by the fact that government in developing countries was known for intervening with a view to lowering risks to producers who depend on export crop prices. More reasonably, one can consider that stabilization schemes acted towards preventing producers from high world-price volatility only in cases when deviations from the long-run equilibrium exceeded a specific threshold. In other words, producer prices may adjust faster to deviations from the long-run equilibrium resulting from largest increases in world prices, meaning that the gap between the producer price and its equilibrium value is larger than a threshold. Note that the magnitude of the estimated threshold would have an economic sense: it would correspond to the minimum gap between the producer price and its equilibrium value required to trigger government intervention towards a faster adjustment of prices. In contrast, over the post-reform period, it is expected that the main causes of negative asymmetric price transmission proposed in the vertical transmission literature also apply to the relationship between world prices and producer prices. In market structures run by private agents, where the buyers are large exporters that are able to take advantage of an unequal bargaining relationship, prices paid to producers may adjust faster to deviations from the long-run equilibrium resulting from decreases in world prices. In such situations, producer prices above their equilibrium value tend to revert quickly to the equilibrium, whereas those below their equilibrium value tend to remain there.

The first step of the analysis consists in determining a break point into the Engle and Granger (1987) relationship that defines the long-run relationship between the world price and the producer price over the 1975-2007 period:

$$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t \quad (1)$$

where  $P_t^p$  and  $P_t^w$  denote the producer price and the world price respectively,  $\xi_0$  and  $\xi_1$  are parameters to be estimated, and  $\epsilon_t$  is the error term, which should be stationary if any long-run relationship exists between the two integrated price series.

Then, I examine the presence of asymmetric adjustments in producer prices over the pre- and post-reform periods following the procedure of Enders and Granger (1998) and Enders and Siklos (2001). The hypothesis of asymmetric adjustments in producer prices characterizing the pre-reform and post-reform periods is tested using a Threshold Auto Regressive (TAR) model. Unlike the standard Engle and Granger (1987) approach which assumes that  $\epsilon_t$  from Eq. 1 behave as an auto-regressive process in the form of:

$$\Delta\epsilon_t = \rho\epsilon_{t-1} + e_t \quad (2)$$

where  $\rho$  measures the speed of convergence of the system and  $e_t$  is a white-noise disturbance, Enders and Granger (1998) and Enders and Siklos (2001) introduced asymmetric adjustment by letting  $\epsilon_t$  behave as a TAR process:

$$\Delta\epsilon_t = I_t\rho_1\epsilon_{t-1} + (1 - I_t)\rho_2\epsilon_{t-1} + \sum \psi_k\Delta\epsilon_{t-k} + \mu_t \quad (3)$$

where  $I_t$  is the Heaviside indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } \epsilon_{t-d} \geq \bar{\theta} \\ 0 & \text{if } \epsilon_{t-d} < \bar{\theta} \end{cases} \quad (4)$$

and  $\bar{\theta}$  is the value of the threshold<sup>9</sup>. As in the standard model, the residuals  $\epsilon_t$  measures the deviation of  $P_t^p$  from its equilibrium defined as  $P_t^{p*} = \xi_0^* + \xi_1^*P_t^w$ . Thus, the condition  $\epsilon_{t-d} \geq \bar{\theta}$  refers to positive deviations from the threshold whereas  $\epsilon_{t-d} < \bar{\theta}$  refers to negative deviations from the threshold. In the particular case where  $\bar{\theta} = 0$ , a positive deviation implies that the producer price is higher than its equilibrium ( $P_t^p > P_t^{p*}$ ) whereas a negative deviation implies that the producer price is smaller than its equilibrium.

The consistency of Eq. 1, 3 and 4 with a wide variety of error correction models, allow an error correction representation for the system. Given the existence of a cointegrating vector in the form of Eq. 1, the error correction representation can be written as:

$$\Delta P_t^p = \eta + \lambda^+ I_t \epsilon_{t-1} + \lambda^- (1 - I_t) \epsilon_{t-1} + \sum_{k=0} \alpha_k \Delta P_{t-k}^w + \sum_{k=1} \beta_k \Delta P_{t-k}^p + \nu_t \quad (5)$$

where  $\lambda^+$  and  $\lambda^-$  are the adjustment coefficients for positive and negative deviations, respectively. As underlined by Meyer and von Cramon-Taubadel (2004), cointegration and ECM are based on the idea of a long-run equilibrium, which prevents  $P_t^p$  and  $P_t^w$  from drifting apart. Consequently, following the framework of Enders and Granger (1998) and Enders and Siklos (2001) asymmetry is considered with respect to the speed of price transmission, not the magnitude. Indeed, asymmetric price transmission implies a permanent difference between positive and negative episodes of transmission, meaning that prices may drift apart, which is incompatible with cointegration.

Enders and Granger (1998) and Enders and Siklos (2001) modified the standard cointegrating Dickey-Fuller test to allow for asymmetric adjustment. They developed a test of the null hypothesis of no-cointegration against the alternative of cointegration with TAR adjustment<sup>10</sup>: the *t-max* statistics (the largest of the individual *t* statistics<sup>11</sup>) and

<sup>9</sup>TAR models can be generalized to multiple thresholds (Balke and Fomby, 1997):

$$\Delta\epsilon_t = \rho_i\epsilon_{t-1} + e_t^{(i)} \quad \text{if} \quad \theta^{(i-1)} < \epsilon_{t-d} \leq \theta^{(i)}, \quad i = 1, \dots, K.$$

with  $-\infty = \theta^{(0)} < \theta^{(1)} < \dots < \theta^{(K)} = +\infty$  and  $e_t^{(i)}$  is a mean zero random disturbance with standard deviation  $\sigma^{(i)}$ .

<sup>10</sup>Ideally, one would like to test the no-cointegration/linearity null hypothesis against the threshold cointegration alternative. However, this cannot be done directly. Balke and Fomby (1997) suggested testing first for no-cointegration versus cointegration and then for threshold behaviour.

<sup>11</sup>Petrucelli and Woolford (1984) showed that the necessary and sufficient conditions for the stationarity of  $\epsilon_t$  in model 3 is  $\rho_1 < 0$ ,  $\rho_2 < 0$  and  $(1 + \rho_1)(1 + \rho_2) < 1$  for any value of  $\bar{\theta}$ .

the  $F$  statistic for the joint hypothesis  $\rho_1 = \rho_2 = 0$ . Critical values are tabulated in Enders and Siklos (2001).

In this analysis, there is no a priori reason to think that the thresholds equal zero. Chan (1993) showed that searching over the potential threshold values so as to minimize the sum of squared errors from the fitted model yields a super-consistent estimate of the threshold. Following the procedure of Chan (1993), the estimated residual series from the cointegrating regression are sorted in ascending order. The largest and smallest 15% of the values are discarded. For each of the remaining values, Eq. 3 is estimated. The estimated threshold yielding the lowest residual sum of squares is retained as the appropriate threshold. I then applied the test for cointegration developed by Enders and Siklos (2001) for cases when the threshold value is unknown.

## 4 Results

### 4.1 Coffee market in Salvador, India and Colombia

In the 1980s and 1990s, the degree of liberalization varied across countries but everywhere the reforms implied the abolition of price stabilization schemes. At the end of the 1980s the government of El Salvador still had a central place in the coffee sector. After 1980 the government's influence had increased, with the nationalization of marketing and exporting activities, through a public agency, Incafe. Incafe was fiercely criticized by producers, because of high export taxes. In 1989, the coffee sector switched towards a liberal form of management and Incafe was broken up (Paige, 1993). This had a direct impact on the relationship between world prices and producer prices. Today, post-reform markets in El Salvador are characterized by a small number of export markets (no more than 30) with the top three accounting for around 70% of the market (ITC 2008). On another hand, with the liberalization, the Council for Salvadoran Coffee gained private sector participation that includes board representation from farmers, processors, exporters, and government representatives.

Before the liberalization of the coffee market in India, a marketing board was in full control of coffee purchasing, processing and exporting. At the beginning of the 1990s the country turned to a liberalized market system, and reforms were introduced gradually. First, producers were allowed to sell a fraction of their production on the domestic market. Then, government involvement in marketing ended and coffee growers were allowed to sell their products to private agents (Krivonos, 2004). Producer prices were, in turn, aligned more closely with world prices. It is interesting to underline that liberalization of agricultural trade in India is often seen as a success. In addition to changing the marketing and pricing systems, reforms allowed the development of many new private sector organizations. Following the liberalization of domestic markets, the number of exporters increased dramatically (Akiyama, Baffes, Larson, and Varangis, 2003). Today, post-reform markets are characterized by a high number of export markets (more than one thousand) (ITC 2008). Moreover, farmers are well-represented within the new Coffee Board, which oversees market development and extension services to them. Farmers' groups also have used the Board to argue for subsidies to encourage productivity. Several farmers' associations stand as pressure groups (Coe, 2006).

In Colombia, before the reforms, the coffee sector was run by a powerful syndicate of producers, the National Coffee Fund. Cardenas (1994) analyzed the relationship between the redistribution and the stabilization functions of a marketing board using a political economy model in several developing countries where the coffee sector was run by mar-



keting boards. His analysis showed that price stabilization was successful in Colombia owing to the checks on the redistribution of coffee revenue. He underlined the fact that in Colombia producers had a direct influence on coffee policy, although government officials and producers had had equal representation since 1978. The National Coffee Fund acted as a stabilizing fund, buying coffee from producers at a guaranteed price. On the other hand, coffee was quite heavily taxed. Trade reforms began in 1990. The system was abolished in 1995, which brought producer prices closer to the world prices (see Fig. 3). In post-reform markets, only five companies account for around 70% of all the private sector exports (Shepherd, 2004). On another hand, the coffee authority still includes solely farmers as members and oversees the minimum coffee price for them.

## 4.2 Data and stationarity tests

Producer prices used in this analysis are monthly average prices paid to the grower at farm-gate level, or the minimum price guaranteed by the Government to the grower, collected by the International Organization of Coffee (ICO). World prices are monthly average prices of Arabica, compiled by the International Monetary Fund, extracted from the International Financial Statistics Database<sup>12</sup>. Both price series are in US cents per libra. The data cover the period from January 1975 to December 2007. The hypothesis that the price series are non-stationary time series over whole periods and sub-periods (determined in what follows) is tested using the Augmented Dickey Fuller (ADF) test. The results indicate that all series are  $I(1)$  at conventional significance levels (Tab.4 in Appendix).

## 4.3 Estimated breakpoints and preliminary results on price transmission

Although the timing of reforms in developing countries is approximately known, it is difficult to fix precisely a break date in the cointegrating relationship between world prices and producer prices simply by examining graphed series. Moreover, political decisions on the dissolution of marketing boards may not lead to an immediate shift of regime in the long-run relationship between prices, as the effects of the reforms on price transmission may be delayed (or even anticipated). Consequently, the residual-based test for cointegration which allows for the possibility of regime shift, developed by Gregory and Hansen (1996), is used to determine the more plausible break date in the long-run relationship defined by Eq. 1. In this alternative model, cointegration holds over some period of time and then shifts to a new long-run relationship. The case where both intercept and slope coefficient have a single break of unknown timing is considered<sup>13</sup>:

$$P_t^p = \xi_0' + \xi_0''\varphi + \xi_1'P_t^w + \xi_1''\varphi P_t^w + \epsilon_t' \quad (6)$$

where  $P_t^p \sim I(1)$ ,  $P_t^m \sim I(1)$  and  $\epsilon_t' \sim I(0)$ . The  $\xi_0'$  coefficient represents the intercept before the shift and  $\xi_0''$  represents the change in the intercept at the time of the shift. The  $\xi_1'$  denotes the cointegrating slope coefficients before the regime shift and  $\xi_1''$  denotes the change in the slope coefficients. The dummy variable  $\varphi$  is defined by:

$$\varphi = \begin{cases} 0 & \text{if } t \leq t_0 \\ 1 & \text{if } t > t_0 \end{cases} \quad (7)$$

<sup>12</sup>Arabica price series is described as Other milds, market price series, arithmetic average of El Salvador Central Standard, Guatemala prime washed, Mexico prime washed, prompt shipment, ex-dock, New-York. Average of daily quotations.

<sup>13</sup>Gregory and Hansen (1996) developed cases where only the intercepts have a break of unknown timing but they are not relevant in this analysis.

where  $t_0$  is the unknown parameter denoting the timing of the change point. Then the ADF statistic and the Phillips test statistics are calculated for all values of  $t_0 \in T$ . The smallest values of the statistics give the more plausible breakpoint  $t_0$ . Results of the residual-based tests for cointegration in models with regime shift are displayed in Tab. 1. Estimated breakpoints from the ADF test are retained because they fit better with both graphed series and timing of reforms in the countries. Estimated breakpoints are October 1994, October 1997 and October 1994, respectively, for El Salvador<sup>14</sup>, India and Colombia. These dates are shown on Figures 1, 2 and 3.

Table 1: Results of Gregory-Hansen test

	Salvador	India	Colombia
$t$ -statistics <sup>a</sup>	-3.939	-5.025**	-4.835*
$t_0$ <sup>b</sup>	238	274	238
date	October 1994	October 1997	October 1994

<sup>a</sup> Smallest  $t$ -statistics using Gregory-Hansen cointegration test among possible break points. \*\*\* (resp.\*\*,\*) : rejection of the null hypothesis at the 1% (resp. 5%, 10%) significance level.

<sup>b</sup>  $t_0$  break point corresponding to the smallest  $t$ -statistic.

Then, on account of the non-stationarity of price series, the Engle and Granger (1987) relationship (Eq. 1) is estimated over both sub-periods, defined according to the estimated breakpoint, and the null of no-cointegration is tested using the Augmented Dickey Fuller (ADF) test. A model close to Eq. 6 is then estimated:

$$P_t^p = \xi_0^{pre}(1 - D) + \xi_0^{post}D + \xi_1^{pre}P_t^w(1 - D) + \xi_1^{post}P_t^wD + \epsilon_t \quad (8)$$

where  $D$  is a dummy variable which equals 1 when  $t > t_0$  or else zero,  $\xi_0^{pre}$  represents the intercept before the shift, and  $\xi_0^{post}$  represents the intercept after the shift. The  $\xi_1^{pre}$  and  $\xi_1^{post}$  coefficients respectively denote the cointegrating slope coefficients before and after the breakpoint. Tests of equality of  $\xi_1$  coefficients between sub-periods are applied using  $F$ -distribution. Results are displayed in Tab. 5 in appendix. The  $t$ -statistics from the ADF test indicate that the null hypothesis of no-cointegration between prices can be rejected over all sub-periods in all countries. Tests of equality of  $\xi_1$  coefficients between sub-periods using  $F$ -distribution produced sample values of 30.50, 196.69 and 349.14, respectively, for El Salvador, India and Colombia, meaning significant differences in long-run transmission. As expected, the estimated coefficients indicate a much closer relationship between prices after the break. The long-run transmission reaches approximately 0.8 in El Salvador and India, and 0.6 in Colombia over the post-reform period.

Finally, the hypothesis of a higher contemporaneous response of  $\Delta P_t^p$  to  $\Delta P_t^w$  and a higher speed of adjustment in  $P_t^p$  after the breakpoint is then tested using an asymmetric error correction model<sup>15</sup>. Given the existence of a single cointegrating vector, an ECM is estimated over each sub-period in the form of:

$$\Delta P_t^p = \eta + \lambda \epsilon_{t-1} + \sum_{k=0} \alpha_k \Delta P_{t-k}^w + \sum_{k=1} \beta_k \Delta P_{t-k}^p + \nu_t \quad (9)$$

<sup>14</sup>In the case of El Salvador, the  $t$ -statistics indicate that the null of no cointegration cannot be rejected at significance levels calculated by Gregory (1996), which means that the long-run relationship between prices is not described better by a model with regime shift. In any case, as the corresponding breakpoint is the more plausible, it is retained as an arbitrary breakpoint for the remaining part of the analysis.

<sup>15</sup>Although parameters from the cointegrating regression are linked to the coefficients of the corresponding ECM, a shift in Eq. 1 does not imply a shift in the corresponding ECM.

where  $\lambda$  is the speed of adjustment coefficients of  $\Delta P_t^p$ , that measures the responsiveness of  $\Delta P_t^p$  to the deviation of  $P_t^p$  from its equilibrium in the previous period. The coefficient  $\alpha_0$  measures the direct impact of  $\Delta P_t^w$  on  $\Delta P_t^p$ . As in the case of the long-run relationship, tests of equality of  $\alpha_0$  coefficients between sub-periods are applied using  $F$ -distribution. The same tests of equality are applied to  $\lambda$  coefficients. The dummy variable  $D$  is included as an interactive variable<sup>16</sup> into Eq. (9):

$$\Delta P_t^p = \lambda^{pre} Z^{pre} + \lambda^{post} Z^{post} + \alpha_0^{pre} \Delta P_t^w (1 - D) + \alpha_0^{post} \Delta P_t^w D + \nu_t' \quad (10)$$

where  $Z^{pre}$  and  $Z^{post}$  are the error correction terms from cointegration regressions run over pre-reform and post-reform periods respectively. Results of error correction models over sub-periods are displayed in Tab. 6 in appendix. As suggested by the Durbin-Watson statistics and the Ljung-Box  $Q(4)$  statistics, autocorrelation in the residuals does not seem to be a problem in all the equations. Tests of equality of  $\alpha_0$  coefficients between sub-periods using  $F$ -distribution produced sample values of 28.73, 23.34 and 169.57, respectively, for El Salvador, India and Colombia, suggesting that the direct impact of  $\Delta P_t^w$  on  $\Delta P_t^p$  is far greater after the break. It ranges from 0.35 to 0.78 in El Salvador, from 0.12 to 0.55 in India, and from almost zero to 0.47 in Colombia. On contrary, tests of equality of  $\lambda$  coefficients between sub-periods give mixed results. Sample values of  $F$ -statistics imply that producer prices do not respond quicker to discrepancies in the long-run relationship between world prices and producer prices in the case of El Salvador and India. The null of equality of  $\lambda$  coefficients can be rejected at the 10% significance level in the case of Colombia. The results of asymmetric cointegration analysis give more information about the adjustment coefficients.

#### 4.4 Results of threshold cointegration analysis

Tab. 2 presents the test results of the TAR-models (when the threshold value is unknown). In each case a TAR-model augmented by lags in  $\Delta \epsilon_t$  is selected using AIC. As shown in the upper part of Tab. 2, over the pre-reform period the values of the  $t$ -max statistics are -2.47, -2.86 and -1.65, respectively, for El Salvador, India and Colombia. These values are smaller than the critical values at conventional levels, which means that the null of no cointegration (against cointegration with threshold) can be rejected in all countries. Moreover, the sample values of the  $\phi$ -statistics are greater than the critical values at conventional levels, which means that the null hypothesis of  $\rho_1 = \rho_2 = 0$  can be rejected. In each case, the point estimates for  $\rho_1$  and  $\rho_2$  suggest convergence, so that the speed of adjustment is higher for negative than for positive discrepancies from the estimated threshold. In El Salvador, the value of the threshold is  $\bar{\theta} = -17.8$ , which means that the speed of adjustment increases as the producer price is set 17.8 US cents (or more) below its equilibrium value. The point estimate of  $\rho_2$  (-0.313) indicates that approximately 31% of a negative discrepancy is eliminated within a month whereas only 11% of a positive discrepancy ( $\rho_1 = -0.107$ ) is eliminated in the same period of time. This means that discrepancies - such as the producer price far below its equilibrium value - are less persistent, which is clearly favourable to producers. Results lead to similar interpretations in India and Colombia where  $\bar{\theta}$  equals -16.6 and -13.5, respectively.

Such asymmetric price adjustments over the pre-reform period also prove to be significant in the ECM estimates (Tab. 3). In all countries, the  $t$ -statistics imply that the coefficients on the positive and negative error correction terms (respectively  $\lambda^+$  and  $\lambda^-$ )

<sup>16</sup>The dummy variable also interacts with the intercept  $\eta$  and the lags of  $\Delta P_t^p$  and  $\Delta P_t^w$  but they do not appear in the equation, in the interest of readability.

Table 2: Results of threshold cointegration analysis with  $\bar{\theta}$  unknown

Country	$\rho_1^a$	$\rho_2^b$	$\phi^c$	$\theta^d$	AIC	$Q(4)^e$
<i>Pre-reform period</i>						
Salvador	-0.107(-2.47)	-0.313(-3.67)	12.46	-17.76	4.777	4.55(0.34)
India	-0.104(-2.86)	-0.476(-6.67)	18.22	-16.64	3.768	2.58(0.63)
Colombia	-0.064(-1.65)	-0.173(-3.90)	9.21	-13.47	2.684	6.67(0.15)
<i>Post-reform period</i>						
Salvador	-0.708(-6.50)	-0.091(-1.02)	29.21	3.12	2.081	2.60(0.63)
India	-0.006(-0.09)	-0.324(-3.84)	8.02	-8.47	2.756	0.36(0.98)
Colombia	-0.476(-4.15)	-0.145(-2.11)	9.97	9.50	3.107	0.60(0.96)

<sup>a</sup> Coefficients and  $t$ -statistics for the null hypothesis  $\rho_1 = 0$ .

<sup>b</sup> Coefficients and  $t$ -statistics for the null hypothesis  $\rho_2 = 0$ .

<sup>c</sup> Sample values of  $\phi$ .  $p$ -value are in parenthesis.

<sup>d</sup> Threshold value determined along with the value of  $\rho_1$  and  $\rho_2$  such that the sum of squared errors from the fitted model is minimum.

<sup>e</sup> Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero.  $p$ -value are in parenthesis.

are significant at conventional levels, meaning that changes in producer prices respond to both negative and positive discrepancies from the estimated threshold. In the three countries, the point estimates of  $\lambda^+$  and  $\lambda^-$  suggest that producer prices adjust so as to eliminate negative deviations more quickly than positive ones. The point estimates imply that producer prices in India adjust so as to eliminate about 46% of a unit change in the deviation of the producer price from its equilibrium in the previous month, when the deviation is smaller than -16.6 (meaning  $\epsilon_{t-1} < -16.6$ ) but only 10% of a unit change in the deviation when this deviation is larger than -16.6 (meaning  $\epsilon_{t-1} \geq -16.6$ ). Results lead to similar interpretations in El Salvador and Colombia where  $\bar{\theta} = -17.8$  and  $-13.5$ , respectively (although the null hypothesis of  $\lambda^+ = \lambda^-$  cannot be rejected in the case of El Salvador). Figure 4 displays an illustration of the asymmetric price adjustment through time in the striking case of India on the pre-reform period.

Over the post-reform period, the test results of the TAR-models suggest an asymmetric price adjustment in the case of Colombia only<sup>17</sup>. The value of the  $t$ -max statistic is -2.11 and the value of the  $\phi$ -statistic is 9.97. Contrary to pre-reform results, the point estimates for  $\rho_1$  (-0.476) and  $\rho_2$  (-0.145) suggest convergence such that the speed of adjustment is higher for positive than for negative discrepancies from the estimated threshold. The value of the threshold is approximately 9.5, which means that the speed of adjustment increases as the producer price is set 9.5 US cents (or more) above its equilibrium value. This means that discrepancies - such that the producer price far above its equilibrium value - are less persistent, which is clearly unfavorable to Colombian producers. Such asymmetric adjustment also proves to be significant in the asymmetric ECM estimates. The point estimates of  $\lambda^+$  and  $\lambda^-$  indicate that producer prices adjust so as to eliminate about 48% of a unit change in the deviation when it is larger than -13.5 (meaning  $\epsilon_{t-1} \geq -13.5$ ) but only 10% of a unit change in the deviation when this deviation is smaller than -13.5 (meaning  $\epsilon_{t-1} < -13.5$ ). This result suggests again that deviations resulting from large decreases in world prices are eliminated relatively quickly. Figure 5 displays an illustration of the asymmetric price adjustment through time.

<sup>17</sup> Asymmetric adjustment is only suggested in the case of El Salvador.

Table 3: Results of asymmetric error correction models with  $\bar{\theta}$  unknown

$$\Delta P_t^p = \eta + \lambda^+ I_t \epsilon_{t-1} + \lambda^- (1 - I_t) \epsilon_{t-1} + \sum_{k=0} \alpha_k \Delta P_{t-k}^m + \sum_{k=1} \beta_k \Delta P_{t-k}^p + \nu_t$$

	$\lambda^{+a}$	$\lambda^{-b}$	$\lambda^- = \lambda^{+c}$	$Q(4)^d$	$DW^e$
<i>Pre-reform period</i>					
Salvador	-0.109(-2.64)	-0.211(-2.51)	1.21(0.27)	5.21(0.27)	0.51(0.47)
India	-0.097(-2.69)	-0.462(-6.32)	21.08(0.00)	0.25(0.99)	0.52(0.47)
Colombia	-0.06(-2.12)	-0.148(-4.21)	3.37(0.07)	1.67(0.79)	0.35(0.55)
<i>Post-reform period</i>					
Salvador	-0.644(-5.29)	-0.033(-0.35)	17.03(0.00)	0.45(0.98)	1.81(0.18)
Colombia	-0.478(-4.76)	-0.104(-1.75)	10.44(0.00)	3.65(0.45)	0.94(0.33)

*t*-statistics are in parentheses.

<sup>a</sup> Error correction terms showing adjustments to positive deviations from the long-run.

<sup>b</sup> Error correction terms showing adjustments to negative deviations from the long-run.

<sup>c</sup> Sample *F*-statistics for the null hypothesis that the speed of adjustment coefficients are equal. *p*-value are in parenthesis.

<sup>d</sup> Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero. *p*-value are in parenthesis

<sup>e</sup> Durbin's test for serial correlation in the disturbance.  $\chi^2$ -statistics and *p*-values in parentheses.

## 5 Conclusion

Some studies have analyzed the implications of structural adjustment for producers' profitability in crop exporting countries, showing that producers may benefit from the reforms under some conditions. Focusing on producer prices, earlier evidence suggests that the reforms increased the share of producer prices in world prices. However, little attention has been paid to producers' exposure to the full volatility of markets after the reforms. This paper aims to show that the reforms led not only to a closer cointegrating relationship and a higher short-run transmission between prices, but also to the emergence of asymmetric adjustments of producer prices unfavorable to farmers. Indeed, a close examination of speed of adjustment in producer prices indicates that pre- and post-reform periods are characterized by asymmetric adjustments, reflecting the influence of public and private agents on price transmission. Results indicate that, in the pre-reform period, largest increases in world prices were transmitted relatively quickly to growers. Whereas in the post-reform period, negative asymmetric transmission has replaced the positive one, as largest decreases in world prices appear to be transmitted faster to farmers. Such results suggests that, on the one hand, government intervention was favorable to producers in terms of price adjustment over the pre-reform period, and on the other hand, over the post-reform period, new private agents are more likely to transmit world price decreases since early response in this case saves them from diminishing their margin.

This paper contributes to the literature on the impact of commodity market reforms on producers, by addressing the topical issue of price shock transmission to producers, using recent developments in cointegrating analysis. As a matter of fact, at least four results lead to the conclusion that the reforms may have worsened producers' vulnerability to world price volatility: higher transmission in the long run, higher transmission in the short run, disappearance of favorable asymmetries in producer price adjustment, and appearance of unfavorable asymmetries in producer price adjustments.

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Table 4: Results of ADF unit root tests

	World market			Salvador			India			Colombia		
<i>1975-2007</i>	level	-1.023	[1]	-1.194	[1]	-0.421	[1]	0.045	[1]	0.045	[1]	
	first diff.	-14.95***	[1]	-21.284***	[1]	-17.178***	[1]	-11.222***	[1]	-11.222***	[1]	
<i>Pre-reform period</i>	level	-0.58	[1]	-0.531	[1]	-0.029	[1]	0.593	[1]	0.593	[1]	
	first diff.	-10.953***	[1]	-17.482***	[1]	-15.068***	[1]	-10.934***	[1]	-10.934***	[1]	
<i>Post-reform period</i>	level	-1.041	[1]	-1.264	[1]	-0.463	[1]	-0.511	[1]	-0.511	[1]	
	first diff.	-10.87***	[1]	-11.564***	[1]	-7.787***	[1]	-11.683***	[1]	-11.683***	[1]	

[1]: Model without constant nor deterministic trend, [2]: Model with constant without deterministic trend, [3]: Model with constant and deterministic trend. \*\* (resp.\*\*\*): Rejection of the null hypothesis at the 5% (resp. 1%) significance level. In the case of world prices, the pre-reform period goes from 1975:1 to 1994:10, as in El Salvador and Colombia.



Table 5: Engle-Granger cointegration results over sub-periods

	Salvador			India			Colombia		
	Pre-reform	Post-reform	$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t$	Pre-reform	Post-reform	$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t$	Pre-reform	Post-reform	$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t$
$\xi_1^a$	0.634(0.025)	0.834(0.007)	0.305(0.016)	0.305(0.016)	0.855(0.026)	0.203(0.012)	0.203(0.012)	0.596(0.015)	0.596(0.015)
$\xi_0$	-11.179(3.576)	-28.838(0.862)	38.448(2.331)	38.448(2.331)	-9.768(2.634)	45.579(1.740)	45.579(1.740)	18.662(1.804)	18.662(1.804)
$N^b$	238	158	274	274	122	238	238	158	158
$t$ -stat <sup>c</sup>	-3.756**	-4.542***	-5.277***	-5.277***	-3.511***	-3.981**	-3.981**	-4.886***	-4.886***
$F$ -stat <sup>d</sup>	30.50***		196.69***	196.69***		349.14***	349.14***		

Standard errors are in parentheses.

<sup>a</sup>  $\xi_1$  and  $\xi_0$  are the parameters from the cointegrating regression.

<sup>b</sup> Number of usable observations.

<sup>c</sup>  $t$ -statistics of the cointegration test. \*\*\* (resp. \*\*, \*) : rejection of the null hypothesis at the 1% (resp. 5%, 10%) significance level.

<sup>d</sup> Sample  $F$ -statistic for the null hypothesis that the coefficients  $\xi_1^{pre}$  and  $\xi_1^{post}$  are equal in the following model:  
 $P_t^p = \xi_0^{pre}(1 - D) + \xi_0^{post}D + \xi_1^{pre}P_t^w(1 - D) + \xi_1^{post}P_t^wD + \epsilon_t$ .

Table 6: Results of error correction models over sub-periods

	Salvador			India			Colombia		
	Pre-reform	Post-reform	$\Delta P_t^p = \eta + \lambda\epsilon_{t-1} + \sum_{k=0} \alpha_k \Delta P_{t-k}^w + \sum_{k=1} \beta_k \Delta P_{t-k}^p + \nu_t$	Pre-reform	Post-reform	Pre-reform	Post-reform	Pre-reform	Post-reform
$\epsilon_{t-1}$	-0.128(0.039)	-0.266(0.078)	-0.162(0.034)	-0.078(0.044)	-0.098(0.022)	-0.200(0.053)			
$\Delta P_t^w$	0.355(0.056)	0.785(0.023)	0.120(0.030)	0.550(0.053)	0.007(0.015)	0.468(0.037)			
$\Delta P_{t-1}^w$	0.123(0.062)	0.285(0.076)	-0.009(0.031)	0.184(0.070)	0.041(0.016)	0.262(0.054)			
$\Delta P_{t-2}^w$	0.016(0.062)	0.116(0.067)	-0.024(0.031)	0.274(0.071)					
$\Delta P_{t-1}^p$	-0.212(0.069)	-0.389(0.092)	0.156(0.062)	-0.021(0.086)	0.297(0.058)	-0.124(0.075)			
$\Delta P_{t-2}^p$	0.015(0.067)	-0.155(0.083)	0.076(0.062)	-0.222(0.080)					
$N^a$	235	155	271	119	236	156			
$DW^b$	0.064(0.801)	0.581(0.446)	0.560(0.454)	0.021(0.886)	1.330(0.249)	0.023(0.880)			
$Q(4)^c$	4.190(0.381)	1.231(0.873)	0.504(0.973)	0.253(0.993)	1.250(0.870)	3.865(0.425)			
$F\text{-stat}^d$	13.52***	195.33***	6.49***	28.70***	17.75***	64.12***			
$\alpha_0^{pre} = \alpha_0^{post}^e$	28.73***	23.34***			169.57***				
$\lambda^{pre} = \lambda^{post}$	0.37	1.19			4.01*				

Standard errors are in parentheses.

<sup>a</sup> Number of usable observations.

<sup>b</sup> Durbin's test for serial correlation in the disturbance.  $\chi^2$ -statistics and  $p$ -values in parentheses.

<sup>c</sup> Significance level of the Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero.

<sup>d</sup>  $F$ -statistics measure the joint significance of the parameters.

<sup>e</sup>  $F$ -statistics for the null hypothesis that  $\alpha_0^{pre} = \alpha_0^{post}$  and  $\lambda^{pre} = \lambda^{post}$  in an ECM including a dummy variable for the break date.

Figure 1: World price and producer price in Salvador (1975-2007)

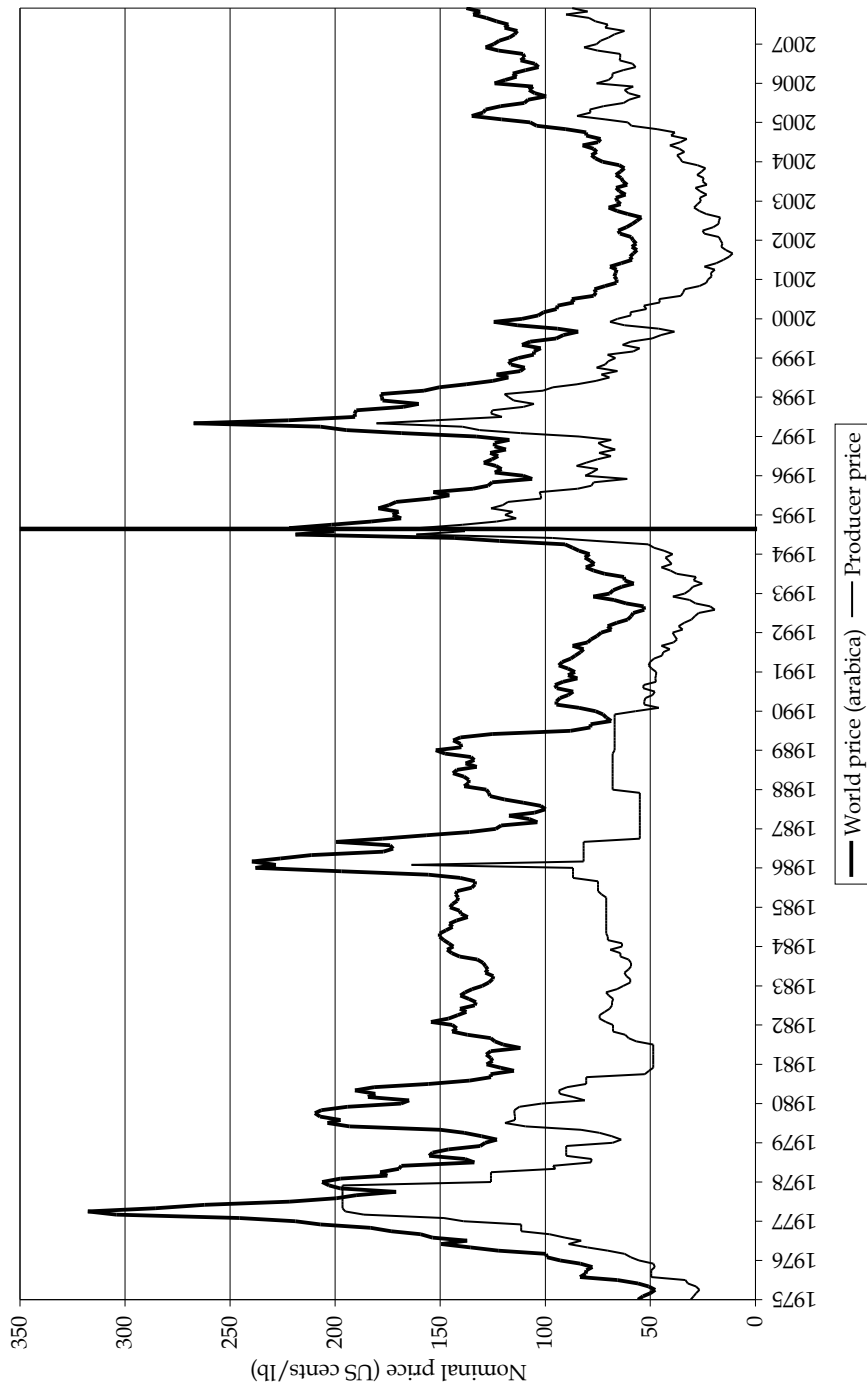


Figure 2: World price and producer price in India (1975-2007)

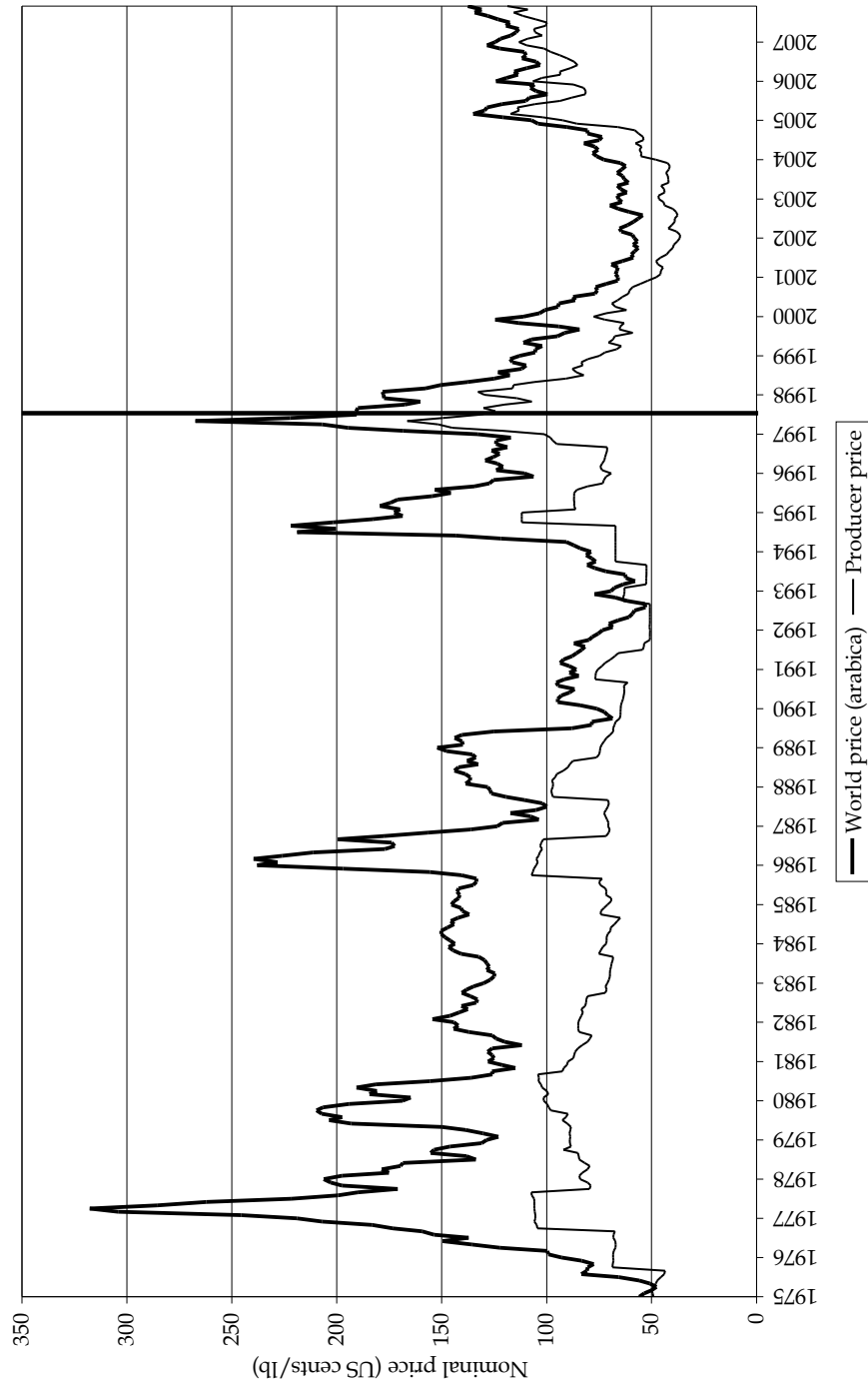


Figure 3: World price and producer price in Colombia (1975-2007)

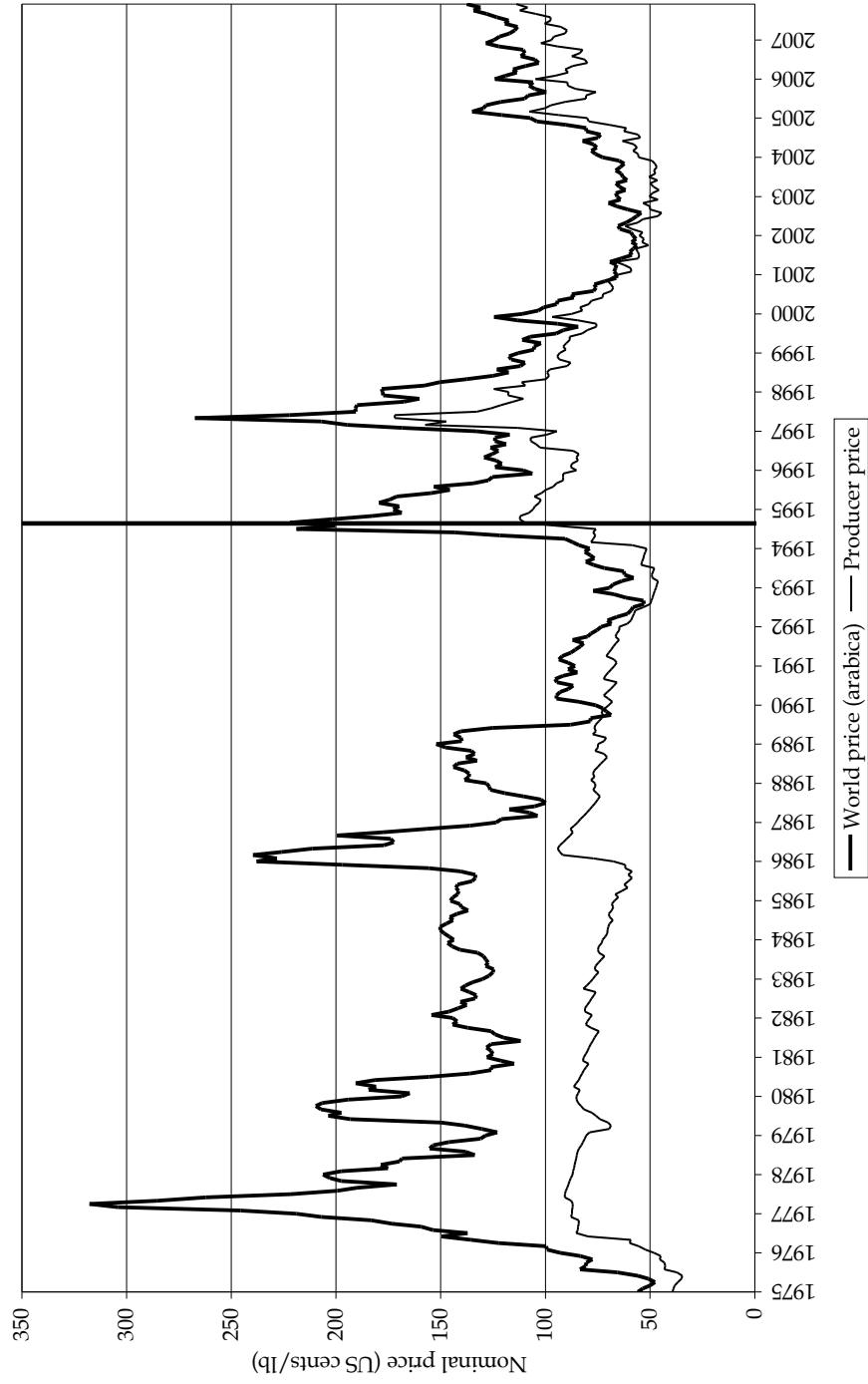


Figure 4: Impulse function of India over the pre-reform period

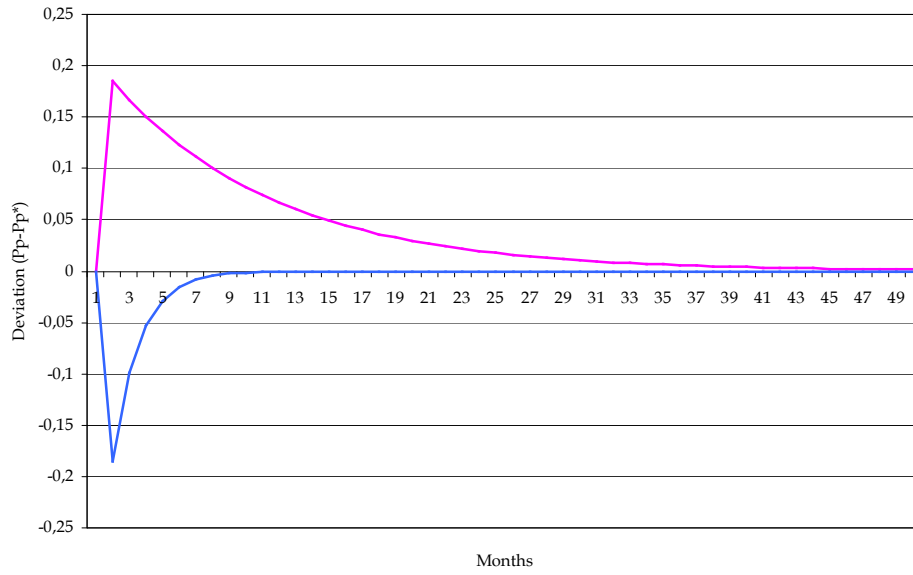


Figure 5: Impulse function of Colombia over the post-reform period

