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"The Impact of Coffe Market Reforms on Price Transmission"

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The impact of coffee market reforms on price transmission

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Abstract

This paper evaluates the impact of coffee sector reforms on shock transmission to producers using threshold cointegration tests that allow for asymmetric adjustment toward long-run equilibrium relationship. The analysis aims at showing that producers' vulnerability to shocks has been worsened by the abolition of price stabilization schemes. The findings show a closer cointegrating relationship between producer prices and world prices after the estimated reform date. The direct impact of monthly variations in world prices on producer price variations has also increased. Moreover, results show that the asymmetric price adjustment that characterized the pre-reform period and was favourable to producers, large deviations from the long-run equilibrium resulting from increases in world prices being eliminated relatively quickly, has disappeared in the post-reform period. In some cases results further show that deviations resulting from decreases in world prices are eliminated relatively quickly over the post-reform period.

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

JEL: C32, O13, O24, D40

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1 Introduction

During the 1980s and 1990s most developing countries in Sub-Saharan Africa and Latin America implemented structural adjustment reforms. These reforms 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 benefited from receiving a larger share of export prices. The degree of liberalization varied across countries. However, everywhere the reforms implied the abolition of price stabilization schemes, exposing producers to the full volatility of markets. Few papers have investigated the responsiveness of producer prices to fluctuations in the world commodity markets. Such analyses typically rely on annual data over short periods. Generally speaking, the findings indicate (surprisingly) high direct transmission between prices over the pre-reform period. However, the interpretation of the results is complicated by the fact that statistical properties of the series are usually ignored. As recent data on producer prices are scarce, the impact of the reforms on the direct relationship between world commodity prices and producer prices has not been studied much either.

The impact of the reforms on world price transmission is a crucial issue for producers in many developing countries. This is because the abolition of price stabilization schemes directly affects producers who depend on the prices of export crops. Producers' vulnerability is worsened by the volatility of agricultural commodity prices, which has been higher over the past three decades than during the pre-1973 period, and their ability to deal with the consequences of such price volatility, which is limited. Short-term effects of commodity market reforms on producers are therefore not limited to higher average prices. They make themselves felt in terms of higher volatility as well. Some analyses have considered the implications of structural adjustment for producers' profitability but little attention has been paid to price volatility issues. A way of studying the impact of commodity market reforms on producers' exposure to shocks on world prices is to estimate the relationship between world prices and producer prices. The first objective of this paper is to investigate changes in both direct relationships between prices and short-run dynamics of price systems over the 1975-2007 period in countries that implemented deep reforms in the 1990s. The second objective is to bring to light the influence on producers of government participation during the pre-reform period, and the influence of new middlemen in crop process chains during the post-reform period. Close examination of the speed of adjustment of producer prices, in a cointegrating framework where producer prices are allowed to respond asymmetrically to world price shocks, illustrates the ways in which pricing policy may be favourable (unfavourable) to producers, typically by relatively quickly absorbing deviations from the long-run equilibrium resulting from increases (decreases) in world prices.

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 price series with no missing data are available: El Salvador, Colombia and India. First, unlike previous studies, the date of the reforms is determined by applying a breakpoint test to the cointegrating relationship. Second, we test the hypothesis of a closer cointegrating relationship after the breakpoint. Then, using a standard error correction model, we test the hypothesis of both higher short-run transmission and higher speed of transmission after the breakpoint. Fourth, we use recently developed threshold cointegration tests that allow for asymmetric adjustment towards a long-run equilibrium relationship, with a view to detecting favourable pricing policy over the pre-reform period and/or unfavourable influence of new private agents over the post-reform period. The 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. Results further show that the asymmetric adjustment that characterized the pre-reform period was favourable to producers, large deviations from the long-run equilibrium resulting from increases in world prices being eliminated relatively quickly, has disappeared in the post-reform period. Moreover, in some cases the results suggest that deviations resulting from decreases in world prices are eliminated relatively quickly 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 asymmetric price transmission, and discusses the expected effects of market reforms on price transmission. Models that can be used to test the hypotheses relative to the contribution of commodity market reforms on coffee producers' exposure to world price shocks are presented in Section 3. The results of the empirical analysis are shown in Section 4. Section 5 presents some concluding remarks.

2 Main findings of empirical literature and hypotheses

The impact of commodity market reforms on producers' exposure to world price shocks combines two sets of empirical literature: studies of world price transmission, and studies of asymmetric price transmission. Although expected, the effects of market reforms on price transmission are rarely investigated in the empirical literature. Moreover, little attention has been paid to the influence of price stabilization shemes on producer price adjustment to shocks.

2.1 Findings from empirical studies of world price transmission

Some papers have examined the relationship between producer prices and world prices, using different data and methods. The estimates of the responsiveness of producer prices to world prices differ sharply from one analysis to another. Moreover, the consequences of the liberalization of export crop markets have not been investigated much in detail.

Evidence of the relationship between producer prices and world prices in empirical studies which focus on pre-reform periods is mixed. Hazell, Jaramillo, et Williamson (1990) examined whether instability in world market prices was transmitted to the prices paid to farmers over the 1966-1987 period. The authors used annual averages of producer prices from the FAO database and annual averages of world prices from the IFS database, for 21 developing countries exporting agricultural commodities such as coffee, cocoa, bananas, cotton or tea. The methodology consists in using a variance decomposition analysis of producer prices, which underlines the separate role of changes in the real exchange rate, in the export unit value, and in government policy and domestic factors. First, the relationship between world prices and producer prices, both in local currency, is approximated with a standard linear regression. Then, according to an approximation defined by Goodman (1960), the variance of producer prices is decomposed into several variability components. The results of the analysis for each of the selected countries and commodities indicate that government policy and the effects of the domestic market and marketing intermediaries are the primary source of variability in producer prices in 56% of the country/commodity cases considered, the variance of world prices is the most important source in 35% of the cases, and the variance of the real exchange rate is most important in the remaining 9% of the cases. In this analysis, the interpretation of the results is complicated by the fact that the contribution of world price variability can be buffered by a covariance component[1]. In the case of coffee, although variability in world prices seems to largely explain variability in producer prices (more than 100% of the variance of producer prices in some cases), the covariance component also seems to play a significant buffering role. In addition, the contribution of the residual component of the decompositiion is relatively important in some country/commodity cases, making the contribution of the other components unclear. In Colombia, where producer prices were administered over the pre-reform period, variability in world prices is a surprisingly large source of variability in producer prices (77%), whereas the covariance component is small enough to be ignored. Such results may be explained by a poor quality of the data[2] and a methodology unsuitable for the statistical properties of the series (in particular, variance-based analyses make no sense when price series are not stationary).

Mundlak et Larson (1992) estimated a direct relationship between producer and world prices over the 1968-1978 period. These authors used annual averages of producer prices from the FAO database and export unit values, calculated as the ratio of the total world value of exports for each of the commodities divided by the total world exported quantities for the corresponding commodities. The authors approximated the relationship between world prices and producer prices with a linear regression, where prices are in logarithms. Results of the estimated transmission in a cross-country comparison both led to surprisingly high transmission elasticities[3] thus suggesting that the commodity-pooling procedure hid some inconsistency in the data. Time series 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 high (0.62 and 1.05, respectively). 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). Finally, in this analysis, statistical properties of the series may give misleading results.

Some studies have analysed the implications of structural adjustment for producers' profitability (Morales (1991), Upton (1993), Baffes et Gautam (1996), Akiyama, Baffes, Larson, et Varangis (2003)) but little attention has been paid to the price volatility issue. Moreover, there are few assessments of commodity market reforms showing a structural break in pricing regimes. Baffes et 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

[1] The decomposition includes a covariance component between the real exchange rate and the export unit value. Insofar as commodities represent quite a large share in total exports, commodity price movements have the potential to explain a significant number of the terms of trade variability. Thus, the real exchange rate is likely to appreciate (depreciate) when world commodity prices increase (decrease).

[2] The FAOSTAT database has been updated since this study.

[3] 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).

model[4] 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[5]. Moreover, only 7 of the 31 cases have a measured nominal rate of protection closer to zero after the reforms than before[6]. 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.

2.2 Findings from empirical studies of asymmetric price transmission

Asymmetric price transmission has received much attention in agricultural economics (see Meyer et 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 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). It is also expected that adjustment costs that arise when firms change the quantities and/or the prices of inputs and/or outputs result in asymmetric price transmission. Peltzman (2000) explained that firms which are afraid of getting out of line with competitors by being the first to raise prices after costs increases may respond faster to cost decreases. Yet his analysis of 77 consumer and 165 producer goods showed that most of the time the price of a good will react faster to an increase of an important input than to a decrease.

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 et 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

[4]Baffes et 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.

[5]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: $k = 1 - (1-\beta)(1-\alpha)^n$.

[6]The null hypothesis on the nominal rate of protection is:

$$H_0 : \sum_{t=1}^{T-1} |\frac{p_t^d - p_t^w}{p_t^w}|/(T-1) \neq \sum_{t=T+1}^{n} |\frac{p_t^d - p_t^w}{p_t^w}|/(n-T-1)$$

where T is the reform year and $(T-1)$ and $(n-T-1)$ denote the pre and post-reform periods.

according to whether the deviation from the long-run equilibrium exceeds some specific threshold levels (Obstfeld et Taylor (1997), Balke et Fomby (1997), Goodwin et Holt (1999), Goodwin et 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 favourable 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. Following the same idea, 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 particular, producer prices may adjust faster to deviations from the long-run equilibrium resulting from large 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 has an economic sense here: it corresponds 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 can 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.

3 Modelling regime shifts and asymmetries in world price transmission

The present analysis aims at determining a break point into the Engle et Granger (1987) relationship that defines the long-run relationship between the world price and the producer price over the 1975-2007 period:

$$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t \tag{1}$$

where P_t^p and P_t^w denote the producer price and the world price respectively, ξ_0 and ξ_1 are parameters to be estimated, and ϵ_t is the error term, which should be stationary if any long-run relationship exists between the two integrated price series. The estimated break point is then used in an error correction model (ECM) to test the hypothesis of higher speed of adjustment coefficients and higher short-run impact coefficients in the post-reform period. Changes in the speed of adjustment of producer prices are further investigated using TAR-models. The analysis examines the presence of asymmetric adjustments in producer prices over the pre- and post-reform periods following the procedure of Enders et Granger (1998) and Enders et Siklos (2001).

3.1 Cointegrating relationship and error correcting model with structural break

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 et 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[7]:

$$P_t^p = \xi_0^{'} + \xi_0^{''}\varphi + \xi_1^{'} P_t^w + \xi_1^{''}\varphi P_t^w + \epsilon_t^{'} \tag{2}$$

where $P_t^p \backsim I(1)$, $P_t^m \backsim I(1)$ and $\epsilon_t^{'} \backsim I(0)$. 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 φ is defined by:

$$\varphi = \begin{cases} 0 & \text{if } t \leq t_0 \\ 1 & \text{if } t > t_0 \end{cases} \tag{3}$$

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.

Then, on account of the non-stationarity of price series, the Engle et 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. 2 is then estimated:

$$P_t^p = \xi_0{}^{pre}(1-D) + \xi_0{}^{post}D + \xi_1{}^{pre}P_t^w(1-D) + \xi_1{}^{post}P_t^w D + \epsilon_t \tag{4}$$

where D is a dummy variable which equals 1 when $t > t_0$ or else zero, ξ_0^{pre} represents the intercept before the shift, and ξ_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 ξ_1 coefficients between sub-periods are applied using F-distribution.

The hypothesis of a higher contemporaneous response of ΔP_t^p to ΔP_t^w and a higher speed of adjustment in P_t^p after the breakpoint is then tested using an asymmetric error correction model[8]. Given the existence of a single cointegrating vector, an ECM is estimated over each sub-period in the form of:

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

[7] Gregory et Hansen (1996) developed cases where only the intercepts have a break of unknown timing but they are not relevant in this analysis.

[8] 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 λ is the speed of adjustment coefficients of ΔP_t^p, that measures the responsiveness of ΔP_t^p to the deviation of P_t^p from its equilibrium in the previous period. The coefficient α_0 measures the direct impact of ΔP_t^w on ΔP_t^p. As in the case of the long-run relationship, tests of equality of α_0 coefficients between sub-periods are applied using F-distribution. The same tests of equality are applied to λ coefficients. The dummy variable D is included as an interactive variable[9] into Eq. (5):

$$\Delta P_t^p = \lambda^{pre} Z^{pre} + \lambda^{post} Z^{post} + \alpha_0^{pre} \Delta P_t^w (1-D) + \alpha_0^{post} \Delta P_t^w D + \nu_t' \qquad (6)$$

where Z^{pre} and Z^{post} are the error correction terms from cointegration regressions run over pre-reform and post-reform periods respectively.

3.2 Asymmetric cointegration and asymmetric error correcting model

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 et Granger (1987) approach which assumes that ϵ_t from Eq. 1 behave as an auto-regressive process in the form of:

$$\Delta\epsilon_t = \rho\epsilon_{t-1} + e_t \qquad (7)$$

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

$$\Delta\epsilon_t = I_t\rho_1\epsilon_{t-1} + (1-I_t)\rho_2\epsilon_{t-1} + \sum \psi_k \Delta\epsilon_{t-k} + \mu_t \qquad (8)$$

where I_t is the Heaviside indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } \epsilon_{t-d} \geq \bar{\theta} \\ 0 & \text{if } \epsilon_{t-d} < \bar{\theta} \end{cases} \qquad (9)$$

and $\bar{\theta}$ is the value of the threshold[10]. As in the standard model, the residuals ϵ_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, 8 and 9 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:

$$\Delta P_t^p = \eta + \lambda^+ I_t\epsilon_{t-1} + \lambda^-(1-I_t)\epsilon_{t-1} + \sum_{k=0} \alpha_k \Delta P_{t-k}^w + \sum_{k=1} \beta_k \Delta P_{t-k}^p + \nu_t \qquad (10)$$

where λ^+ and λ^- are the adjustment coefficients for positive and negative deviations, respectively. As underlined by Meyer et von Cramon-Taubadel (2004), cointegration and

[9]The dummy variable also interacts with the intercept η and the lags of ΔP_t^p and ΔP_t^w but they do not appear in the equation, in the interest of readability.

[10]TAR models can be generalized to multiple thresholds (Balke et Fomby, 1997):

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

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

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 et Granger (1998) and Enders et 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 et Granger (1998) and Enders et 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[11]: the *t-max* statistics (the largest of the individual t statistics[12]) and the F statistic for the joint hypothesis $\rho_1 = \rho_2 = 0$. Critical values are tabulated in Enders et Siklos (2001). Inference concerning the individual values of ρ_1 and ρ_2 and the restriction $\rho_1 = \rho_2$ can be done using classic t intervals when the true value of the threshold is known (Enders, Falk, et Siklos, 2007). However, the property of asymptotic multivariate normality has not been established when the true value of the threshold is unknown.

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. 8 is estimated. The estimated threshold yielding the lowest residual sum of squares is retained as the appropriate threshold. Enders et Siklos (2001) also developed a test for cointegration 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 (see Fig. 1).

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 (see Fig. 2).

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

[12] Petrucelli et Woolford (1984) showed that the necessary and sufficient conditions for the stationarity of ϵ_t in model 8 is $\rho_1 < 0$, $\rho_2 < 0$ and $(1 + \rho_1)(1 + \rho_2) < 1$ for any value of $\bar{\theta}$.

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 marketing 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).

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[13]. 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.10 in Appendix).

4.3 Cointegration and error correction model with regime shift

Although price series strongly suggest a regime shift in the 1975-2007 period, Eq. 1 and Eq. 5 are first estimated over the whole period. The Engle-Granger relationship is estimated and tested for standard cointegration using the ADF test. Results are displayed in Tab. 1. The estimated values of ξ_1 parameters are 0.68, 0.35 and 0.28, respectively, for El Salvador, India and Colombia. This indicates that the long-term impact of a one-unit increase in world prices generates a 0.68-unit increase in producer prices in El Salvador, but only a 0.3-unit increase in producer prices in India or Colombia. The t-statistics from the ADF test indicate that the null hypothesis of no-cointegration between prices can be rejected in all countries, in spite of stabilization schemes implemented over the first-half period. The short-run dynamics of price series is examined with an error correction model. The results are displayed in Tab. 2. In each case, the appropriate lag length is determined using an autocorrelation test[14]. The Ljung-Box $Q(4)$-statistic and the Durbin-Watson statistic indicate that the residuals are not significantly correlated. The standard errors imply that the coefficients of the error correction terms are significant at conventional levels in all countries, which means that world prices and producer prices may move apart for some months but return to a long-run equilibrium, defined by Eq. 1. In the case of El Salvador, the estimated value of the coefficient of the ΔP_t^w variable is 0.48, which means that an increase of 1 in monthly variation of world prices generates a 0.48 increase in monthly variation of producer price. The direct impact of ΔP_t^w on ΔP_t^p is smaller in the case of India (0.15) and Colombia (0.14).

[13] Arabica price series is described as Other milds, market price series, arithmetic average of El Salvador Central Standard, Guatemala prime washed, Mexica prime washed, prompt shipment, ex-dock, New-York. Average of daily quotations.

[14] Lags of ΔP^p and ΔP^w are added as long as autocorrelation tests reject the null of no autocorrelation.

Table 1: Engle-Granger cointegration results (1975-2007)

$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t$

	Salvador	India	Colombia
$\xi_1{}^a$	0.681	0.355	0.276
	(0.017)	(0.016)	(0.016)
ξ_0	-15.229	33.683	1.937
	(2.242)	(2.086)	(0.108)
N[b]	396	396	396
t-stat[c]	-4.919***	-5.079***	-3.263*

Standard errors are in parentheses.

[a] ξ_1 and ξ_1 are the parameters from the cointegrating regression.

[b] Number of usable observations.

[c] t-statistics of cointegration tests. *** (resp.**,*) : rejection of the null hypothesis at the 1% (resp. 5%, 10%) significance level.

Table 2: Results of error correction models (1975-2007)

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

	Salvador	India	Colombia
ϵ_{t-1}	-0.136(0.029)	-0.098(0.021)	-0.044(0.017)
ΔP_t^w	0.483(0.037)	0.152(0.026)	0.143(0.021)
ΔP_{t-1}^w	0.116(0.044)	0.005(0.027)	0.064(0.023)
ΔP_{t-2}^w			0.005(0.023)
ΔP_{t-3}^w			0.023(0.023)
ΔP_{t-4}^w			0.037(0.023)
ΔP_{t-5}^w			0.012(0.023)
ΔP_{t-1}^p	-0.204(0.049)	0.160(0.050)	0.065(0.052)
ΔP_{t-2}^p			-0.069(0.051)
ΔP_{t-3}^p			0.061(0.051)
ΔP_{t-4}^p			-0.200(0.051)
ΔP_{t-5}^p			-0.079(0.050)
constant	0.041(0.433)	0.110(0.305)	0.186(0.240)
N^a	394	394	390
DW^b	0.007(0.931)	0.002(0.961)	0.200(0.655)
$Q(4)^c$	5.766(0.217)	0.995(0.910)	0.118(0.998)
F-statistics[d]	58.58***	16.15***	11.20***

Standard errors are in parentheses.

[a] Number of observations.

[b] Durbin's test for serial correlation in the disturbance. χ^2-statistics and p-values in parentheses.

[c] The Q-statistics denote the Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero. p-values are in parentheses.

[d] The F-statistics measure the joint significance of the parameters.

Results of the residual-based tests for cointegration in models with regime shift are displayed in Tab. 3. 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[15], India and Colombia. These dates are shown on Figures 1, 2 and 3.

Table 3: Results of Gregory-Hansen test

	Salvador	India	Colombia
t-statistics^a	-3.939	-5.025**	-4.835*
t_0^{b}	238	274	238
date	October 1994	October 1997	October 1994

^{a} 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.

^{b} t_0 break point corresponding to the smallest t-statistic.

Eq. 1 is estimated and tested for standard cointegration over sub-periods defined according to the estimated breakpoint. Results are displayed in Tab. 4. 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 ξ_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.

Results of error correction models over sub-periods are displayed in Tab. 5. 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 α_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 ΔP_t^w on ΔP_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 λ 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 λ 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.

[15]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.

Table 4: Engle-Granger cointegration results over sub-periods

$$P_t^p = \xi_0 + \xi_1 P_t^w + \epsilon_t$$

	Salvador		India		Colombia	
	Pre-reform	Post-reform	Pre-reform	Post-reform	Pre-reform	Post-reform
${\xi_1}^a$	0.634(0.025)	0.834(0.007)	0.305(0.016)	0.855(0.026)	0.203(0.012)	0.596(0.015)
ξ_0	-11.179(3.576)	-28.838(0.862)	38.448(2.331)	-9.768(2.634)	45.579(1.740)	18.662(1.804)
N^b	238	158	274	122	238	158
t-stat^c	-3.756**	-4.542***	-5.277***	-3.511***	-3.981**	-4.886***
F-stat^d	30.50***		196.69***		349.14***	

Standard errors are in parentheses.

[a] ξ_1 and ξ_0 are the parameters from the cointegrating regression.

[b] Number of usable observations.

[c] t-statistics of the cointegration test. *** (resp.**,*) : rejection of the null hypothesis at the 1% (resp. 5%, 10%) significance level.

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

Table 5: Results of error correction models over sub-periods

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

	Salvador		India		Colombia	
	Pre-reform	Post-reform	Pre-reform	Post-reform	Pre-reform	Post-reform
ϵ_{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)
ΔP_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)
ΔP_{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)
ΔP_{t-2}^w	0.016(0.062)	0.116(0.067)	-0.024(0.031)	0.274(0.071)		
ΔP_{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)
ΔP_{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-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.

[a] Number of usable observations.

[b] Durbin's test for serial correlation in the disturbance. χ^2-statistics and p-values in parentheses.

[c] Significance level of the Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero.

[d] F-statistics measure the joint significance of the parameters.

[e] 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.

4.4 Asymmetric cointegration and asymmetric error correction model

Tab. 6 presents the test results of the TAR-models when the threshold value is set equal to zero. In each case, a TAR-model augmented by lags in $\Delta\epsilon_t$ is selected using AIC. Over the pre-reform period, the values of the *t-max* statistics are -1.96, -0.82 and -1.92, respectively, for El Salvador, India and Colombia. These values are smaller than the critical values at the 10% level, which is around -1.90 for the model with one lagged change, in the case of El Salvador and Colombia. This means that the null of no cointegration (against cointegration with threshold) can be rejected. Moreover, the sample values of the ϕ-statistics are greater than the critical values at the 5% level for both countries, which means that the null hypothesis of $\rho_1 = \rho_2 = 0$ can be rejected. But the null hypothesis of $\rho_1 = \rho_2$ tested using a standard F-distribution, cannot be rejected in both cases, which means that adjustments are not significantly asymmetric. Over the post-reform period, the values of the *t-max* statistics are -0.84, -0.09 and -1.03, respectively, for El Salvador, India and Colombia. These values are greater than the critical values at the 10% level, again suggesting that price adjustments are not significantly asymmetric over this period.

Tab. 7 presents the test results of the TAR-models when the threshold value is unknown. Here again, 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. 7, 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 ϕ-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 ρ_1 and ρ_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 ρ_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 $\bar{\theta} = -16.6$ and -13.5, respectively. The test results of the TAR-models over the post-reform period suggest an asymmetric price adjustment in the case of Colombia only. The value of the *t-max* statistic is -2.11 and the value of the ϕ-statistic is 9.97. Contrary to pre-reform results, the point estimates for ρ_1 (-0.476) and ρ_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 unfavourable to Colombian producers.

The results of asymmetric error correction models when the threshold value is set equal to zero are presented in Tab. 8. In accordance with the test results of the TAR-models, producer price adjustments do not prove to be asymmetric in the ECM, as both the coefficients of speed of adjustment (λ^+ and λ^-) are not significantly different from zero over either the pre-reform period or the post-reform period. On the other hand, asymmetric price adjustments suggested by test results of TAR-models over the pre-reform period prove

Table 6: Results of threshold cointegration analysis with $\bar{\theta} = 0$

Country	${\rho_1}^a$	${\rho_2}^b$	ϕ^c	${\rho_1 = \rho_2}^d$	AIC	$Q(4)^e$
Pre-reform period						
Salvador	-0.101(-1.96)	-0.251(-3.04)	11.42	2.03(0.15)	4.789	4.04(0.40)
India	-0.043(-0.82)	-0.382(-5.55)	14.29	11.38(0.00)	3.805	3.33(0.50)
Colombia	-0.140(-1.92)	-0.095(-1.94)	8.11	0.18(0.67)	2.697	4.62(0.33)
Post-reform period						
Salvador	-0.611(-5.15)	-0.093(-0.84)	23.93	8.48(0.00)	2.149	3.43(0.49)
India	-0.025(-0.28)	-0.226(-0.09)	5.99	1.78(0.18)	2.814	0.21(0.99)
Colombia	-0.373(-2.91)	-0.109(-1.03)	8.84	1.79(0.18)	3.141	0.44(0.98)

[a] Coefficients and t-statistics for the null hypothesis $\rho_1 = 0$.
[b] Coefficients and t-statistics for the null hypothesis $\rho_2 = 0$.
[c] Sample values of ϕ. p-value are in parenthesis.
[d] Sample F-statistic for the null hypothesis that $\rho_1 = \rho_2$. p-value are in parenthesis.
[e] Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero. p-value are in parenthesis.

Table 7: Results of threshold cointegration analysis with $\bar{\theta}$ unknown

Country	${\rho_1}^a$	${\rho_2}^b$	ϕ^c	$\bar{\theta}^d$	AIC	$Q(4)^e$
Pre-reform period						
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)
Post-reform period						
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)

[a] Coefficients and t-statistics for the null hypothesis $\rho_1 = 0$.
[b] Coefficients and t-statistics for the null hypothesis $\rho_2 = 0$.
[c] Sample values of ϕ. p-value are in parenthesis.
[d] Threshold value determined along with the value of ρ_1 and ρ_2 such that the sum of squared errors from the fitted model is minimum.
[e] Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero. p-value are in parenthesis.

to be significant in the asymmetric ECM estimates for an unknown threshold. The results are displayed in Tab. 9. In all countries, the t-statistics imply that the coefficients on the positive and negative error correction terms (respectively λ^+ and λ^-) 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 λ^+ and λ^- 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).

Table 8: Results of asymmetric error correction models with $\bar{\theta} = 0$

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

	λ^{+a}	λ^{-b}	$\lambda^- = \lambda^{+c}$	$Q(4)^d$	DW^e
Pre-reform period					
Salvador	-0.120(-2.45)	-0.147(-1.81)	0.07(0.80)	4.89(0.30)	0.06(0.80)
India	-0.027(-0.54)	-0.331(-4.98)	9.82(0.00)	1.68(0.79)	0.83(0.36)
Colombia	-0.089(-1.60)	-0.103(-2.70)	0.03(0.87)	1.31(0.86)	1.38(0.24)
Post-reform period					
Salvador	-0.570(-4.35)	-0.104(-0.88)	6.18(0.01)	0.31(0.99)	1.80(0.18)
India	-0.07(-0.94)	-0.08(-0.95)	0.00(0.96)	0.25(0.99)	0.02(0.87)
Colombia	-0.467(-4.27)	0.018(0.19)	7.72(0.01)	3.33(0.50)	0.15(0.70)

t-statistics are in parentheses.

[a] Error correction terms showing adjustments to positive deviations from the long-run.

[b] Error correction terms showing adjustments to negative deviations from the long-run.

[c] Sample F-statistics for the null hypothesis that the speed of adjustment coefficients are equal. p-value are in parenthesis.

[d] Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero. p-value are in parenthesis

[e] Durbin's test for serial correlation in the disturbance. χ^2-statistics and p-values in parentheses.

Over the post-reform period, the price asymmetric adjustment that characterizes the case of Colombia proves to be significant in the asymmetric ECM estimates. The point estimates of λ^+ and λ^- 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.

5 Conclusion

Some studies have analysed 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.

Table 9: Results of asymmetric error correction models with $\bar{\theta}$ unknown

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

	λ^{+a}	λ^{-b}	$\lambda^- = \lambda^{+c}$	$Q(4)^d$	DW^e
Pre-reform period					
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)
Post-reform period					
Salvador	-0.644(-5.29)	-0.033(-0.35)	17.03(0.00)	0.45(0.98)	1.81(0.18)
India	-0.014(-0.24)	-0.312(-4.45)	9.65(0.00)	7.20(0.12)	0.27(0.60)
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.

[a] Error correction terms showing adjustments to positive deviations from the long-run.

[b] Error correction terms showing adjustments to negative deviations from the long-run.

[c] Sample F-statistics for the null hypothesis that the speed of adjustment coefficients are equal. p-value are in parenthesis.

[d] Ljung-Box statistic that the first four of the residual autocorrelations are jointly equal to zero. p-value are in parenthesis

[e] Durbin's test for serial correlation in the disturbance. χ^2-statistics and p-values in parentheses.

This paper aims to show that the reforms led not only to a closer cointegrating relationship but also to a higher short-run transmission between prices. Moreover, 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. In particular, empirical results indicate that producer prices relatively quickly corrected large deviations resulting, in the pre-reform period, from increases in world prices and, in the post-reform period, from decreases in world prices. This suggests that, on the one hand, government intervention was favourable to producers in terms of price adjustment over the pre-refrom period, and on the other hand, private agents are more likely to transmit world price decreases since early response in this case saves them from diminishing their margin over the post-reform period.

This paper contributes to the literature on the impact of commodity market reforms on producers, by addressing the topical issue of world price transmission to producers, using recent developments in cointegrating analysis. 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 favourable asymmetries in producer price adjustment, and appearance of unfavourable asymmetries in producer price adjustments. For all that, the question of did the reforms benefit producers remains difficult to answer, as liberalization of the crop sector may affect producers in many ways. However, in the short term it seems to result in a trade-off between higher price volatility and higher price levels.

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Table 10: Results of ADF unit root tests

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

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

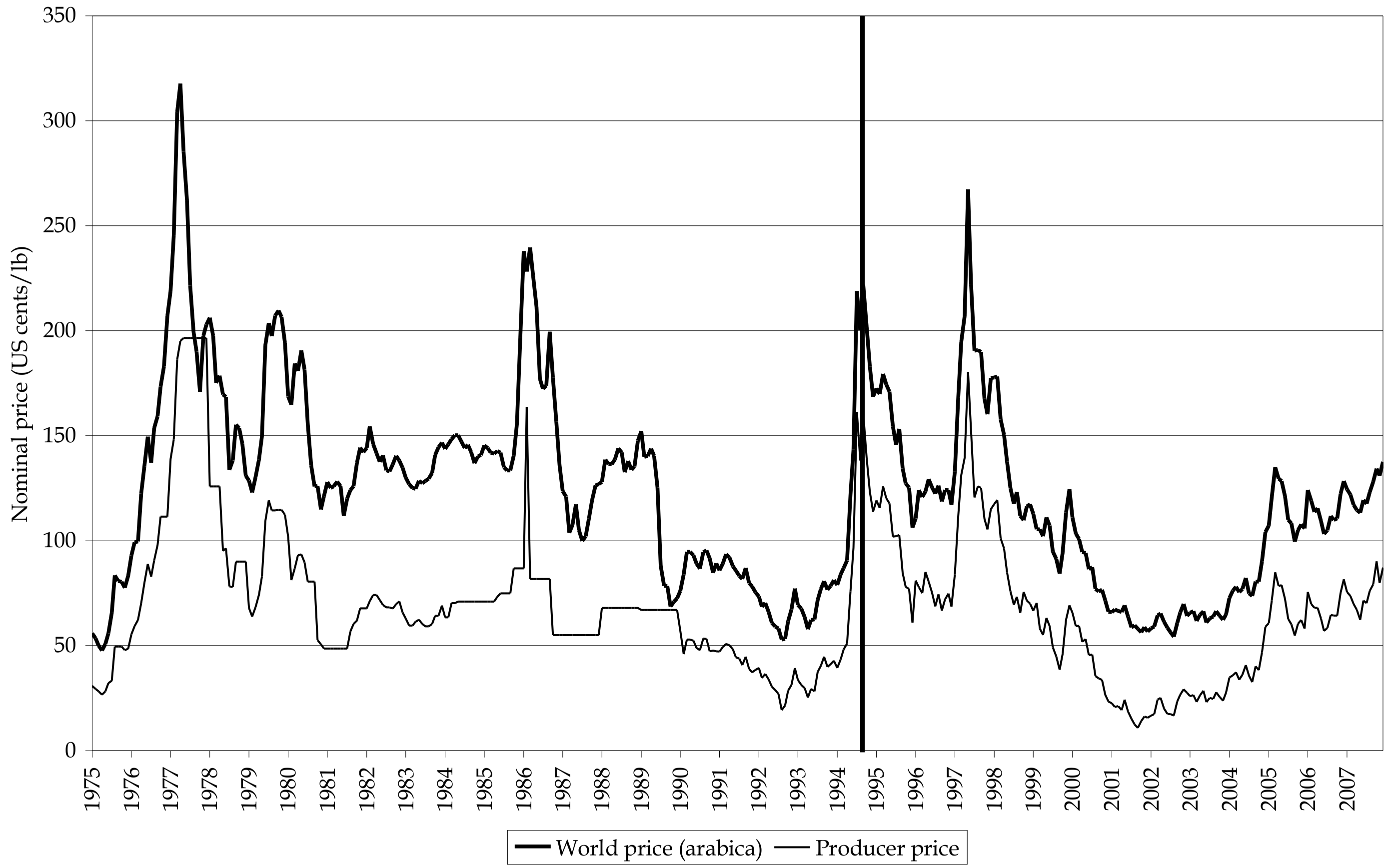

Figure 2: World price and producer price in India (1975-2007)

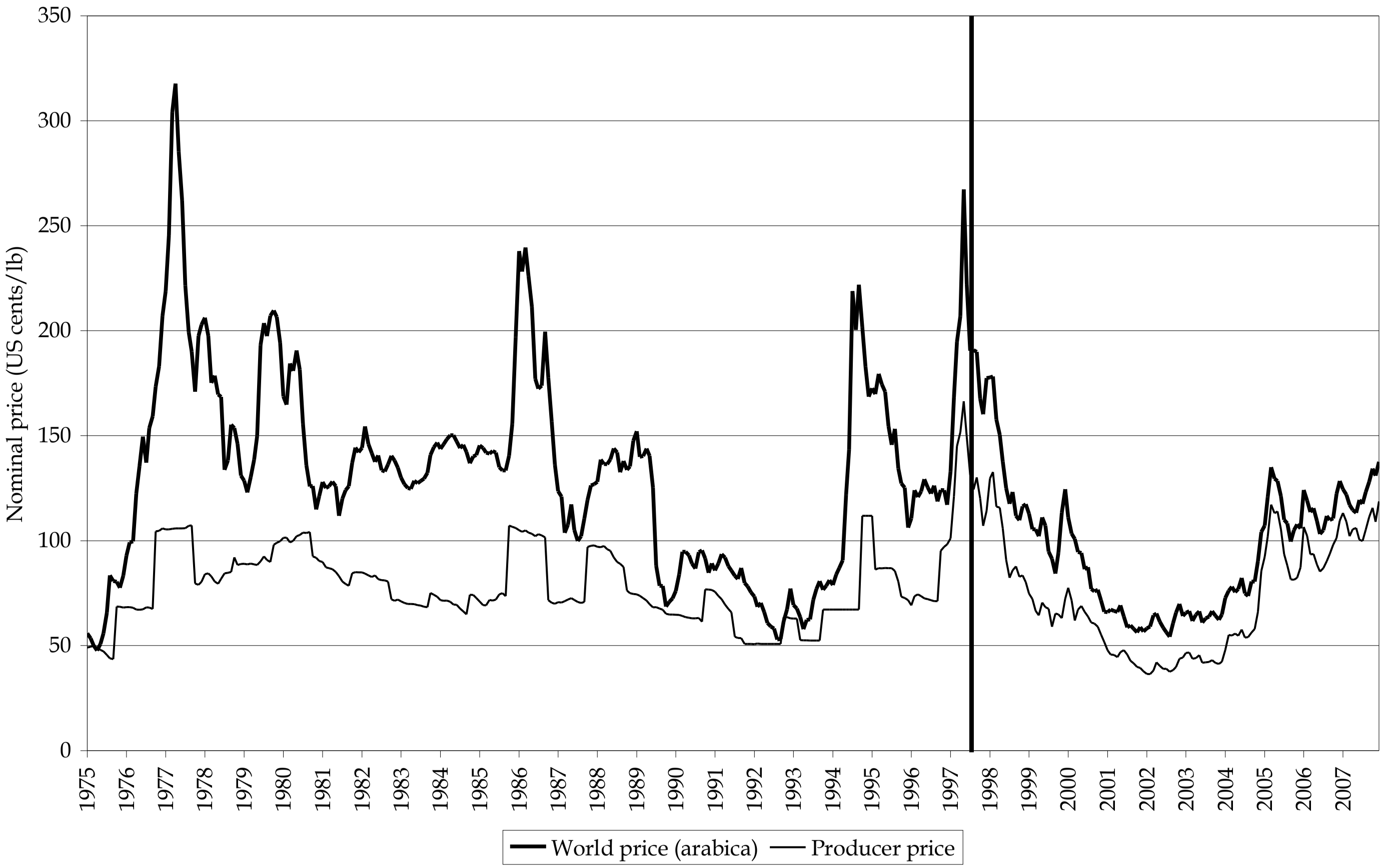

Figure 3: World price and producer price in Colombia (1975-2007)