# Trends and persistence of farm-gate coffee prices around the world

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**Abstract**
This paper examines the dynamics of real coffee prices received by growers. First, we analyse the long run trends of coffee prices to determine whether producers of coffee are relatively worse off over time as has been suggested by influential reports. Given the biological nature of production of coffee we make conjectures that coffee prices can be characterised by large swings that can last several years, and accordingly, we consider whether prices can be characterised by structural breaks that cause a change in the sign and/or magnitude of the trend. Secondly, given the variability in coffee prices, an important issue for farmers' is whether any shock to the prices they receive is short-lived or not. To investigate both of these questions, we conduct robust econometric tests and exploit a unique data set for selected countries that grow coffee. We find no evidence of any structural breaks and therefore breaking trends, and little evidence of any significant secular trend. We find mixed results with reference to whether shocks to coffee prices are transitory. The results are informative as they dispel some of the beliefs about trends in farm-gate coffee prices. We conclude by outlining policy implications based on our empirical findings.

**JEL Classifications:** C22; O13; Q02; Q11
**Keywords:** trends; shocks; coffee prices; coffee farmers.

## 1. Introduction

Since 1990, which marks the period from which the coffee sector has been liberalised, global coffee production has expanded by 65% (ICO 2019a) due to an increased demand from emerging economies as well as an increased demand in speciality coffee from traditional markets. The *Coffee Development Report* published by the International Coffee Organisation (ICO 2019b) notes a continued downward trend in world prices of coffee since 2016, which could have an adverse impact on coffee growers in terms of their income, and the ability to cover the costs of their production and welfare. These findings echo a coffee market report published by *Oxfam* in 2001 which painted a bleak picture of the coffee producing countries, largely reliant on the income earned through the production and export of coffee. The report concluded that international prices of coffee have been declining especially in the recent years since the report was published, thereby spelling doom and gloom for coffee farmers. The main reasons cited for this decline was oversupply of coffee has caused stocks to rise over time leading to prices becoming depressed. The increase in supply is largely driven by new plantations of coffee, new arrivals of coffee exporters, and technological progress. In comparison, the demand for coffee shows little fluctuation due to the lack of close substitutes. Both demand and supply are known to be relatively inelastic (Mehta and Chavas 2008), which means that shifts in demand or supply, or both, can cause large swings in coffee prices. For example, coffee prices recorded significant slumps from 1990 to 1993, and from 1999 to 2003 (Cuaresma et. al. 2018) and then again in recent years since 2016 (ICO 2019b); however, there have been episodes which have recorded an upward drift in prices between 2004 and 2011 (Cuaresma et. al. 2018). The upshot is that coffee prices may be subject to structural breaks caused by sudden large shifts in supply, and therefore display periods of what might seem to be decreasing as well as and increasing trends. We address some key questions: whether this perceived notion of decline in coffee prices as documented in previous studies is a long run phenomena, or is this decline a short run temporary phase, interspersed with increasing trends or no trends? In that case is it possible that coffee prices are characterised by broken trends? This paper therefore aims to determine whether real coffee prices received by farmers has gradually declined over time, or is characterised by broken trends, which change in sign and magnitude. We focus on the period of time when most countries started to liberalise their coffee sector and in that way aim to find empirical evidence of whether such liberalisation helped to prevent a sustained decline in coffee prices that would have otherwise made coffee farmers worse off over time.

In general, measuring trends in commodity prices are difficult given the high variability that exists in these prices. As Angus Deaton notes "What commodity prices lack in trend they make up for in variance (1999; p.27)." Coffee is no exception and the large fluctuations in coffee prices are largely driven by the shifts in demand and supply curves which are relatively inelastic; and can also be attributed to the weather patterns and production lags in response to sudden increases or decreases in demand. The variability in coffee prices is believed to deter producers from making necessary investments for increasing productivity, which in turn could lead to potential shortages in coffee supply. Since liberalisation of the coffee sector, prices have become more variable and this hampers the ability of farmers to expand production, invest in inputs and service debt. A crucial question that arises here is whether any shocks to coffee prices are transitory or not. If shocks to coffee prices are not short-lived then risk management policies and government intervention may be needed to help farmers cope with smoothing the shocks to their incomes and consumption as well as providing the necessary capital for maintaining production.

This paper exploits a unique data set of coffee prices paid to growers that starts from 1990 to recent years, for selected developing countries and emerging economies. To our knowledge, this data has not been used from a time series perspective and therefore will provide useful insights on the long run dynamics. These farm-gate coffee prices are adjusted for inflation, and robust econometric methods are applied to estimate the underlying trend parameters to conclude whether they are statistically significant. As a prelude to the analysis, we use robust econometric methods to search for structural breaks in order to ascertain whether the plausible underlying trend is subject to a change in the sign and/or magnitude. The results show there are no signs of significant structural breaks in the real price of farm-gate coffee over time, and very little evidence of significant trends. Therefore, we cannot conclude that coffee farmers are worse off in the long run which is contrary to popular reports. We further investigate whether shocks to coffee prices are transitory or not. Our results are mixed with no clear pattern, which suggest that individual coffee prices need to be evaluated on a case by case basis, irrespective of coffee variety, or country of origin when designing risk management policies to protect farmers' from price shocks. The paper is organised as follows: section 2 provides an institutional background; section 3 describes the empirical framework; section 4 describes the data and the empirical analysis; and finally section 5 concludes.

## 2. Institutional Background

Coffee is the one of the most widely traded commodities (Borrella et. al. 2015) and is mainly produced in the tropics and sub-tropics which includes developing countries as well as some middle income emerging economies. For many of the developing countries, coffee accounts for a large share of export earnings making these countries highly reliant on coffee as their major source of income. Almost 70 percent of all coffee grown, is by smallholders and an estimated 25 million producers are dependent on coffee production for their livelihoods (Borrella et. al. 2015). It has been argued that rural poverty and economic vulnerability is becoming prevalent among coffee growing regions and a declining trend in coffee prices has been listed as one of the main reasons for the declining income and profitability of coffee growers, especially smallholders (Dietz et. al. 2020). This declining income hampers the achievement of the Sustainable Development Goals (SDGs) which is to raise income in rural areas, create rural employment and alleviate poverty. Clearly, a substantial drop in coffee prices to growers, threatens the livelihood of millions of smallholder producers and risks reversing any gains in living standards. Soon after the demise of the ICA, the international coffee market had undergone changes and was not immediately understood by most coffee growers, especially smallholders; however, to some larger farmers it was becoming evident that changes were to happen (Samper 2010). What was seemingly an acute but presumably short-term crisis, between 1989 and 1993, followed by subsequent recovery and then by largely speculative upturns and downturns, can be viewed as a structural change in the world coffee market.

On the supply side, except for short-term situations, the growth of production has tended to exceed that of demand, for several reasons. An initial increase in price is seen as a signal for all large coffee producers to increase supply as well as attract new entrants who may increase production over and above the domestic demand for coffee. The area devoted to the planting of coffee trees can be expanded in response to improved market conditions such as higher prices (e.g., prices were pushed up over three years following the sudden shortfall in global coffee production such as the frosts in Brazil and the coffee-berry borer disease in Colombia in 1994). Coffee trees take time to bear fruit from the time of planting to the point of harvest which could typically be three to five years. Brazil has been a dominant producer of coffee and apart from the temporary setback the country faced when inflicted with a frost that destroyed a large number of coffee trees, it is unlikely that any such production shortfall will be faced by the country as production has gradually shifted towards the north from the frost-prone south, since the turn of the century.

The demand for coffee is inelastic, given the taste and preferences of coffee consumers, implying that a significant change in coffee prices are needed to induce a change in consumption. Shifts in the demand for coffee is a function of lower levels of income and tends to stagnate at higher income levels and is therefore known as a 'mature' market (see Ponte 2002). This has prompted roasters to cater to concentrate on the marketing of coffee by brands, and countries of origin, to increase the value added to coffee. For example, as Ponte (2002) notes, there is targeting of specific countries where there is a potential for demand as well as influencing the taste and habits of the consuming country[1]. The combination of unequal shifts in demand and supply, that are relatively inelastic, can cause the price of coffee to produce large swings.

The International Coffee Agreement (ICA) had an objective to control prices and restrict supply with an export quota system applied to coffee producing countries to keep the international price within a fixed band. The ICA collapsed in 1989 thereby leading to the removal of regulations on international coffee supply and marked a period of increased price variability. The vulnerability of farmers to this variability of the producer price of coffee, and plausible decreasing trends adds to the difficulty of farmers to plan future coffee supply, the inability to cover costs when prices slump, and the difficulty to invest in capital to modernise their farming as smallholders in particular, already face liquidity constraints. Prolonged periods of depressed prices coupled with rising costs of production, can lead to increased out-migration and this can have a profound impact on the poorer sections of a country's population given that labour accounts for more than 50% of the total cost of coffee production (ICO 2019b). If coffee prices are on a decreasing trend, and if land labour and capital costs are increasing, then farmers would have to intensify the production of coffee (Bernstein 1994) to cover their costs and maintain their income.

Since the coffee crisis from 1990 to 1993, countries have been encouraged to diversify into other crops or speciality coffees. For smallholders this is difficult as diversification can require capital investment and associated risk before potential benefits emerge. For example, the government of Brazil set up PRONAF (National Program for the Strengthening of Family Farming) to support small holders with agricultural credit at low interest rates to offset costs

[1] For example, Germany has a preference for the Mild Arabica variety whereas France and Italy prefer Robusta. UK and USA prefer a wide variety (see Ponte (2002)).

and risks; however, the process to acquire funds has turned out to be bureaucratic (David et. al. 2000). Brazil is a leader among the global exporters of coffee. However, the country specializes in low quality coffee and is driven by low costs of labour, and in the context of a deregulated market with high level of competition, smallholder end up with low prices and income (Caldarelli et. al. 2019). In Colombia, the FNC (Federación Nacional de Cafeteros) protects prices for its farmers through stabilization funds and ensuring quality control for coffee exports. The FNC provides a guarantee of purchase to coffee growers, which ensures that all producers – especially smallholders – can sell their coffee to the NFC and are protected by a price floor (Vellema et al 2015). In contrast, production conditions in Honduras are challenging for farmers whose livelihoods depend on the prices they receive from coffee. A large section of the population is engaged in coffee production which is decentralised, reliant on intermediaries and there is a low level of intervention. Therefore farmers are vulnerable to price and production shocks, and alternative employment is difficult for coffee farmers and labourers when they are faced with adverse price shocks (Dietz et al. 2019). Costa Rica has favourable natural conditions for the production of high-quality coffee and possesses a strong organizational structure in the production and marketing stages of coffee. In recent years Costa Rica has increased its production of high quality coffee. The marketing system in Costa Rica ensures that certain quality standards are met to ensure such classification of speciality coffee and farmers have been able to receive higher prices for high-quality coffee (Wollni and Brummer 2012). Ethiopia produce high-end quality Arabica coffee. In Ethiopia, coffee is considered the most important cash crop and its production is an important source of livelihood for the vast majority of smallholder farmers (Kuma et. al. 2019).

Prior to liberalisation, there was a system of regulation and intervention. During this period, governments aimed to protect coffee growers' incomes from price fluctuations, through price stabilization, which actually turned out to be costly and often inefficient, causing producer prices to fall as a share of world price levels. Since most coffee farmers produce on small holdings, the fall in price had an impact on exacerbating poverty levels. However, when coffee market sectors for various countries started to liberalise, the share of producer prices in the world market price increased, albeit with producers being exposed to the risks of volatile world coffee prices. By the mid-1990s most coffee producing countries had eliminated or reduced state-controlled marketing, such as withdrawing state trading enterprises, and giving way to market-based systems by allowing private agents to be involved in purchasing, marketing and exporting of coffee. Typically, competition was encouraged among traders; guaranteed minimum prices to farmers were withdrawn and export taxes were lowered or eliminated (Krivonos 2004). A drastic decline in international coffee prices took place in 2001/02 for both Arabica and Robusta varieties. Coffee prices fell below production costs due to oversupply, causing severe difficulties for coffee farmers (Lewin et. al. 2004). Since the liberalisation of the coffee sector, it has been argued that volatility has increased; however, McIntyre and Varangis (1999) note that farmers were receiving a higher price with volatile prices than they would have received with administered prices. Stabilisation policies were insignificant and often marketing boards were found to be corrupt (Krivonos 2004). Since liberalisation, with supposedly higher prices and increased variability, the following questions arise: (a) whether the so-called declining trend in coffee prices is significant or overshadowed by the variability? (b) whether the large swings in coffee prices has caused a break in the trend and that price trends have changed magnitude and or sign? (c) are shocks to coffee prices short-lived? These are questions that we plan to address in this paper.

## 3. Empirical Framework

The above discussion concludes that coffee prices may be characterised by a secular trend or by broken trends where the sign and/or magnitude may differ across regimes. To this end we employ robust procedures that allow us to determine whether there are structural breaks calling for regimes and broken trends, or whether we can simply estimate a secular trend. The trend may be secular or broken depending on the relative size of the persistent shifts in demand and supply over time.

### 3.1 *Robust tests for presence and location of multiple structural breaks*

Accordingly, we employ a robust test for structural breaks in the data to establish whether we should estimate a secular trend or whether the estimation of broken trends would be more appropriate. If we find structural breaks, that would imply the trend is not secular and that either trend estimation would include regimes where the sign and/or magnitude of the trend may be different in each regime. The tests we employ allow us to be agnostic as to whether the real coffee price series chosen in this study contain stochastic trends, that is whether they are I(1), as opposed to being I(0). To this end we adopt the test due to Sobriera and Nunes (2016) by using the following specification:

$$P_t = \alpha + \beta t + \sum_{j=1}^{n} \delta_j DU_t(\tau_j^*) + \sum_{j=1}^{n} \gamma_j DT_t(\tau_j^*) + \varepsilon_t, \qquad t = 1, 2, \dots, T \tag{1}$$

where $P_t$ denotes the logged prices and the trend estimate is given by the parameter $\beta$. The specification allows for $n$ structural breaks in the trend function. These breaks may occur at dates $T_1^*, \dots . T_n^*.$, and the level dummies $DU_t(\tau_j^*) = 1(t > T_j^*)$ detect the eventual $j$th break and the slope dummies $DT_t(\tau_j^*) = 1(t > T_j^*)(t - T_j^*)$ detect the eventual $j$th break at date $T_j^* = \lfloor \tau_j^* T \rfloor$ for $j = 1, \dots, n$, with the indicator function given by 1(.) and the integer part of the argument given by $\lfloor . \rfloor$. We can write (1) in first differenced form as:

$$\Delta P_t = \beta + \sum_{j=1}^{m} \delta_j D_t(\tau_j^*) + \sum_{j=1}^{m} \gamma_j DU_t(\tau_j^*) + v_t, \quad t = 1, 2, \dots, T \tag{2}$$

where $D_t(\tau_j^*) = 1(t = T_j^* + 1)$.

We estimate (1) and (2) by ordinary least squares (OLS) for all possible break points $\tau^n = (\tau_1, \dots, \tau_n)$ which are obtained from employing the supremum $F$ test given by:

$F_0^*(n|0) = supF_0(\tau^n)$ and $F_1^*(n|0) = supF_1(\tau^n)$

where $F_0(\tau^n)$ and $F_1(\tau^n)$ denote respectively the $F$ statistics for testing $\tau_1 = \tau_2 = \cdots . \tau_n = 0$ from (1) and (2). Sobriera and Nunes (2016) note that the asymptotic distributions of $F_0^*(n|0)$ and $F_1^*(n|0)$ are dependent on whether the underlying price series is I(0) or I(1), and so they work out a weighted statistic that yields the same asymptotic critical values in both I(0) and I(1) cases. This robust weighted statistic $F_\lambda^*(n|0)$ due to Sobriera and Nunes (2016) can be used to test the null hypothesis of no structural breaks against the alternative hypothesis of $n$ breaks, and is given by:

$$F_\lambda^*(n|0) = \lambda(\hat{\tau}^n, \tilde{\tau}^n) F_0^*(n|0) + d_\xi^n [1 - \lambda(\hat{\tau}^n, \tilde{\tau}^n)] F_1^*(n|0)$$

where $d_\xi^n$ is a constant that ensures for a given significance level $\xi$, the asymptotic critical values are the same, irrespective of the order of integration of the data.

A further sequential test due to Sobreira and Nunes (2016) is employed which is in the spirit of Bai and Perron (1998), testing the null hypothesis of $m$ against the alternative of $m + 1$

breaks constructed from the maximum value of the $supF$ type statistics. The procedure involves first starting with $m = 0$, and then using $F_\lambda^*(1|0)$ to test for the presence of one break. If the null is rejected, we then set $m = 1$ and perform the $F_\lambda^*(2|1)$ test. If the null is rejected, we continue this sequence until we cannot reject the $F_\lambda^*(m + 1|m)$ test.

Noting that this sequential procedure of detecting structural breaks may not work very well if there are two breaks in the slope of opposite sign, Sobriera and Nunes (2016) recommend that if the null hypothesis of no break is not rejected, (that is, when $F_\lambda^*(1|0)$ is not rejected), then to use the $F_\lambda^*(2|0)$ or the double maximum test $UDmaxF_\lambda^*$ or $WDmaxF_\lambda^*$. If $F_\lambda^*(2|0)$ or a double maximum test does not reject the null hypothesis, then we conclude no trend breaks, otherwise we test for 2 breaks against 3 and the sequential procedure is continued. If there is no evidence of any structural break or multiple breaks in the price series, we proceed to test for a secular trend in the price data over the full sample of observations. If we find $m$ breaks, then we demarcate $m + 1$ regimes based on the break point locations and estimate broken trends for these selected regimes.

3.2 *Robust estimation of trends*

To estimate the trends in the data, we make use of another robust test that allows us to be agnostic to the underlying order of integration in the data. To this end, we make use of the Perron and Yabu (2009a) procedure. To implement this procedure, we assume the error term in (1) to follow an autoregressive process where the lag is determined according to the modified Akaike Information Criterion (MAIC). A bias corrected version of the autoregressive parameter is created to improve the finite sample properties of the test, from which a quasi-differenced regression is estimated (see Perron and Yabu 2009a for details).

The Perron and Yabu (2009a) procedure to estimate the trend is carried out in the following manner. First, the following auto-regression on the error term (say $\hat{u}_t$) of a trend function is estimated:

$$\hat{u}_t = \alpha\hat{u}_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta\hat{u}_{t-i} + e_{tk} \qquad (3)$$

We estimate $\tilde{\alpha}$ from regression (3), and in order to improve the finite sample properties of the test we use a bias-corrected version denoted $\tilde{\alpha}_M$ following the recommendation by Roy and

Fuller (2001)[2]. Perron and Yabu (2009a) construct the super-efficient estimate $\tilde{\alpha}_{MS}$ as follows:

$$\tilde{\alpha}_{MS} = \begin{cases} \tilde{\alpha}_M \text{ if } |\tilde{\alpha}_M - 1| > T^{-0.5} \\ 1 \quad \text{if } |\tilde{\alpha}_M - 1| \leq T^{-0.5} \end{cases} \tag{4}$$

The super-efficient estimate allows us to implement procedures that yield nearly identical limit properties with $I(0)$ and $I(1)$ variables. The super-efficient estimate $\tilde{\alpha}_{MS}$ is then used to estimate the following quasi-differenced regression:

$$\begin{aligned} &(1 - \tilde{\alpha}_{MS}L)y_t = (1 - \tilde{\alpha}_{MS}L)x'_t\Psi^0 + (1 - \tilde{\alpha}_{MS}L)u_t; \qquad t = 2,3,\ldots,T \\ &y_1 = x'_1\Psi^0 + u_1 \end{aligned} \tag{5}$$

where $\Psi^0 = (\mu_0, \beta_0)'$. Denoting the estimate $\hat{\beta}_0$ from this regression, we construct a $100(1-\alpha)\%$ confidence interval for $\beta_0$ valid for both $I(1)$ and $I(0)$ errors, and is given as follows:

$$\hat{\beta}_0 \pm c_{\alpha/2}\sqrt{(\tilde{h}_v)\{(X^{FG\prime}X^{FG})^{-1}\}_{22}} \tag{6}$$

where $c_{\alpha/2}$ is such that $P(x > c_{\alpha/2}) = \alpha/2$ and $\tilde{h}_v$ is an estimate of $2\pi$ times the spectral density function of $v_t = (1 - \alpha L)u_t$ at frequency zero (see Perron and Yabu 2009b for details). We define $X^{FG} = (x_1^{FG}, x_2^{FG}, \ldots x_T^{FG})'$ where $x_t^{FG} = [1 - \tilde{\alpha}_{MS}, t - \tilde{\alpha}_{MS}(t-1)]$ for $t - 2,3,\ldots,T$ and $x_1^{FG} = (1,1)'$. From the estimate $\hat{\beta}_0$, we denote the corresponding t-statistic due to the Perron and Yabu (2009a) test as $t_{PY}$.

3.3 *Testing for unit roots allowing for nonstationary volatility*

The most commonly used method to analyse the persistence of agricultural commodity prices is using unit root tests (see Ghoshray 2019, Wang and Tomek 2007). These tests can determine if there is a unit root in the coffee price series, or alternatively, if the price is integrated of order one, or I(1); in which case a shock would have a permanent effect on the price. Alternatively if we reject the price has a unit root, then the price is integrated of order zero, or I(0), and shocks to prices would be transitory in nature. Given the popularity of these tests, a large

[2] See Perron and Yabu (2009b) for details.

literature has evolved, investigating the presence of a unit root in agricultural commodity prices. The results so far have been broadly inconclusive and there can be several reasons why we find mixed results. First, the unit root question in agricultural prices cannot be properly analysed until some characterisation of the underlying deterministic component is made. The uncertainty of whether or not to include a constant, or a constant and linear trend in a unit root test regression, is a problem that affects unit root tests. If a linear trend is present in the data but is not accounted for in the unit root test, then the test is likely to under-reject the unit root null. The converse is true as well; that is, the null hypothesis of a unit root will be prone to under-rejection, if a trend is not present in the data but is included in the unit root test, as it will lead to a loss of power (Marsh 2007). Secondly, the possible presence of nonstationary volatility in the underlying data can affect the size properties of the unit root tests. Conducting unit root tests on coffee prices without allowing simultaneously for nonstationary volatility, can cause the tests to suffer from poor size issues, implying false rejection of the unit root null.[3] This paper addresses these limitations through the application of a novel unit root tests proposed by Smeekes and Taylor (2012), which simultaneously deals with uncertainty regarding the deterministic trend and the possibility of nonstationary volatility in the data.

This test is useful as it avoids the erroneous conclusions that can arise from standard ADF tests. When nonstationary volatility is present in the data, the ADF test is asymptotically not correctly sized (Cavaliere and Taylor 2008). Besides, the presence or absence of a linear trend in the data series can lead to problems with unit root testing. Muller and Elliott (2003) show that the Dickey Fuller ($DF$) test with ordinary least squares detrending, denoted $DF-OLS$, suffers from low power relative to the $DF$ test with quasi-differenced ($QD$) or generalised least squares ($GLS$) detrending, denoted as $DF-QD$.[4] The upshot is that there is uncertainty with regards to which test to apply. To deal with these issues, Harvey et. al. (2009) construct a new test formed as a union of rejections of unit root tests with and without a deterministic linear trend and show that this union test can maintain high power and size irrespective of the true value of the trend. Harvey et. al. (2012) propose a four-way union of rejections of $DF-QD$

[3] A further reason, which is largely statistical and therefore not mentioned in detail, is the initial condition (defined as the deviation of the initial observation from the deterministic components) is also known to have a major impact on the power of unit root tests (see Muller and Elliot 2003, Phillips and Magdalinos 2009). A large initial condition could appear in the data if we are dealing with an unusual period such as change in agricultural policy. Ignoring the initial condition can lead to possibly erroneous results, a point that has been overlooked in the literature. The procedure by Smeekes and Taylor (2012) will address this limitation as well.

[4] This problem of low power applies if the initial condition is small; alternatively, if the initial condition is large, then the opposite happens.

and $DF-OLS$ tests, both with and without trend. Therefore, the procedure is a modified union wild bootstrap test that is robust to nonstationary volatility, which is asymptotically valid and is also shown to perform very well in finite samples. The modified union test statistic, which is a four-way union of rejections of $DF-QD^{\mu}$, $DF-QD^{\tau}$, $DF-OLS^{\mu}$ and $DF-OLS^{\tau}$ is given by:

$$UR_4 = min\left[DF-QD^{\mu}, \left(\frac{cv_{QD}^{\mu*}(\pi)}{cv_{QD}^{\tau*}(\pi)}\right)DF-QD^{\tau}, \left(\frac{cv_{QD}^{\mu*}(\pi)}{cv_{OLS}^{\mu*}(\pi)}\right)DF-OLS^{\mu}, \left(\frac{cv_{QD}^{\mu*}(\pi)}{cv_{OLS}^{\tau*}(\pi)}\right)DF-OLS^{\tau}\right] \quad (7)$$

where $cv_j^{\delta}(\pi)$ denotes the asymptotic critical value of the Dickey Fuller test which could contain an intercept, or intercept and trend, at nominal level $\pi$. This test thereby deals with the uncertainty about the trend and the initial condition.

However, Smeekes and Taylor (2012) note that the uncertainty about the presence of a trend and the initial condition needs to be combined with the possible presence of nonstationary volatility. To this end, Smeekes and Taylor (2012) consider union tests that are robust to nonstationary volatility, trend uncertainty, and uncertainty about the initial condition. Their test is based on the wild bootstrap approach, combined with the sieve principle to account for stationary serial correlation, designed to be robust over uncertainty about the presence of a deterministic trend and uncertainty about the initial condition. This test is an improvement over the union tests of Harvey et. al. (2012) which are found to be incorrectly sized in the presence of nonstationary volatility. The wild bootstrap variant of the union tests proposed by Smeekes and Taylor (2012) overcome these problems, showing that the tests are robust to nonstationary volatility and retain their validity. They consider two bootstrap union tests, unit root A type test ($UR_{4A}$) and unit root B type test ($UR_{4B}$); the $UR_{4A}$ test does not include a deterministic trend in the test, while the $UR_{4B}$ test does include a trend in the bootstrap data generating process. The bootstrap union tests proposed by Smeekes and Taylor (2012) would appear to constitute a valuable option if one needs to deal simultaneously with uncertainty regarding the trend and the initial condition and to provide results that are simultaneously robust to the possible presence of nonstationary volatility.

**4. Data and Empirical Results**

The data employed in this study are coffee prices paid to producers of various coffee growing countries deflated by their consumer price index to obtain the real coffee prices. The source of the data is the *International Coffee Organisation*. The countries we consider are Brazil, Colombia, Costa Rica, Ethiopia, Honduras, India, Indonesia and Uganda. For Brazil, Uganda and India we analyse both the Arabica and Robusta varieties, and for the remaining countries only the Arabica variety[5]. The start date of the sample is chosen to be in the early 1990s, a time when most countries had started to liberalise their coffee sector. Many developing countries in sub-Saharan Africa and Latin America reformed their coffee sectors as the existing system of marketing was costly and inefficient. The speed of reforms differed across countries, but by the mid-1990s most coffee growing counties had replaced state controlled marketing with private agents. The data is monthly and has varying start and end dates depending on the availability of the data. The details of the sample range and size of each country's coffee prices and the currency unit are given in Table 1 along with the descriptive statistics of each of the prices. The results are contained in Table 1 below.

[Table 1 about here]

The coefficient of variation is a relative measure of dispersion that expresses the sample standard deviation in terms of its mean and provides a unit-less measure of dispersion. The coefficient is the highest for India (Robusta), recording 67% variation. Honduras and Ethiopia have a significantly high level of dispersion recording figures over 40%. Relatively the dispersion of real coffee prices for Brazil (both Robusta and Arabica) along with Colombia and Costa Rica are relatively low with the coefficient of variation being less than 30%. Besides, we find significant positive skewness for all coffee prices, except Coast Rica and Uganda (Robusta). This would imply that there are few or no downward spikes to match the pronounced upward spikes. Except for Uganda (Robusta) none of the coffee prices show significant negative skewness. All coffee prices also display substantial kurtosis, with tails much thicker than those of the normal distribution, a feature that, is not uncommon in agricultural commodities. The skewness and kurtosis measures are designed to be zero and 3 respectively for a normal distribution, and since the coffee prices are found to exhibit positive skewness and excess kurtosis, it is not surprising that the test for normality of coffee prices is rejected. A plot of the different coffee prices considered in this study are shown in Figure 3 below.

[5] This is based on the availability of data.

[Figure 3 about here]

From the figure it is difficult to discern whether there is an underlying trend in the data. The data appears to be highly variable with several large positive spikes – which are expected given the prevalence of positive skewness. A common feature for all Arabica coffee prices is the large positive spike around 1997/98 and then a gradual decline albeit with a considerable and varying range of volatility, to reach a low at around 2001/02. Thereafter Arabica prices tend to increase until 2011 before the volatility increases. The graphs for Robusta prices seem to follow a different dynamic path to Arabica coffees. One may conclude that the data is characterised by large upswings and downswings which either could be thought of as breaking trends in coffee prices or high volatility. We treat these features of the data with high importance when conducting robust tests for structural breaks and trend estimates.

Accordingly, before estimating a secular trend, we test for structural breaks to ascertain the need for estimating broken trends. We employ the robust procedure due to Sobreira and Nunes (2016) to detect structural breaks in the trend of the data series where the number and dates of the breaks are unknown and are robust to the order of integration of the data. The results are shown in Table 2 below.

[Table 2 about here]

The application of the test statistics $F_{\lambda}(m|0)$ for $m = 1,2,3$ and the double maximum statistics $UDmaxF_{\lambda}$, along with the $WDmaxF_{\lambda}$ to the various coffee prices are compared against the 10% critical values. Following the procedure by Sobriera and Nunes (2016) if the null hypothesis of no break is not rejected, that is, when $F_{\lambda}^{*}(1|0)$ is not rejected, then we proceed to test $F_{\lambda}^{*}(2|0)$ or the double maximum test $UDmaxF_{\lambda}^{*}$ or $WDmaxF_{\lambda}^{*}$. We allow up to 3 breaks so we also test for $F_{\lambda}^{*}(3|0)$. If the $F_{\lambda}^{*}(3|0)$ or double maximum tests do not reject the null hypothesis, then we conclude no trend breaks. For all coffee prices, the $F_{\lambda}(m|0)$ statistic for $m = 1,2,3$ and the double maximum statistics $UDmaxF_{\lambda}$, along with the $WDmaxF_{\lambda}$ lead us to conclude no rejection of the null hypothesis of no trend break. For example, in the case of Brazil (Arabica) the null of no break against the alternative of a single break returns a test statistic of 2.26 which is less than the critical value at conventional levels (at least 10%) and therefore we cannot reject the no break null. This implies that the null hypothesis of no trend breaks cannot be rejected for all coffee prices. Given that there is not enough evidence to conclude that coffee prices have any structural break in the trend, we can assume that the evidence favours an estimation of an unbroken trend for these coffee prices. At this juncture, we can infer that the large upswings and downswings in the data as shown in Figure 1 are not linked to broken trends, but are likely due to the large volatility.

We test for the presence of significant trends in coffee prices for the entire sample considered, since there is no evidence of trend breaks and therefore no need to estimate broken trends at points where structural breaks could have been identified. Accordingly, we first apply the Perron and Yabu (2009a) procedure for robust trend estimation. The results of the estimation are shown in Table 3, along with the 95% and 90% confidence intervals. The robust t-statistics are also reported for reference.

[Table 3 about here]

For example, in the case of Brazil (Arabica) the trend estimate is –0.35; however, the associated 95% confidence interval has a lower bound of –1.10 and an upper bound of 0.41. Since the confidence interval contains zero the trend estimate is insignificant. The same can be said when using the 90% confidence interval, where the lower bound is –1.25 and the upper bound is 0.55. The associated insignificant t-statistic confirms the result that the trend estimate is insignificant. Using this robust procedure, we find no evidence of a significant trend in any of the real coffee prices, irrespective of the country of origin, or the variety of coffee, except for Honduras and India (Robusta) coffee prices. The parameter estimates for Honduras and India (Robusta) are negative and the t-statistics indicate significance at both the 95% and 90% confidence levels, implying that the real prices for these two coffee prices have been declining over time. The trend estimates of India (Robusta) show a fair amount of variability with a lower bound decline of 1.32% and an upper bound decline of 0.28% at the 5% significance level; or a lower bound decline of 1.41% and an upper bound decline of 0.19% if one were to consider the 10% significance level. The range in both cases exceed 1 percentage point. In contrast the variability in the trend estimates for Honduras are lower, with a lower bound decline of 0.25% and an upper bound decline of 0.07% at the 5% significance level. The estimates hardly change when choosing a higher level of significance, but the trend estimate is smaller in magnitude and the range is smaller at approximately 0.18 percentage points relative to India (Robusta) at a little over 1 percentage point.

Some of our results depart from the conclusion made by other studies. While Minten et. al. (2019) document an increasing trend in coffee prices for Ethiopia, we find an insignificant trend. Further, Gong and Sullivan (2017) note that coffee prices have been increasing in Uganda, whereas we find no trend for both Arabica and Robusta varieties. The lack of trends could result from producer prices displaying large upswings and downswings which may be in response to government policy that could indirectly affect production and marketing in the coffee sector. In Honduras and India (Robusta) we show coffee prices have a declining trend, implying producers have systematically lost purchasing power when exchanging a unit of coffee for a bundle of consumer goods over time. If increases in productivity and efficiency are not sufficient to offset this effect, farmers could be economically worse off today. Over the past two decades productivity has increased in Honduras, while it has broadly stagnated in India for the Robusta variety.

Table 4 shows the results of the bootstrapped union tests for unit roots allowing for nonstationary volatility on all the coffee prices. Using the robust procedure due to Smeekes and Taylor (2012) we find that only for 4 cases (that is, Costa Rica, Honduras India (Arabica) and Uganda (Arabica)) we can reject the null hypothesis at least at the 10% significance level. This is true for both the $UR_{4A}$ and $UR_{4B}$ tests where we include and exclude a trend in the data.

[Table 4 about here]

This implies that for these coffee prices any shocks will be short-lived and priced will revert to their long run equilibrium price; in the case of Honduras the reversion will be to the long run equilibrium trend, whereas for the remaining three prices the reversion will be to the long run mean. For the remaining 6 prices, where we cannot reject the null hypothesis of a unit root. We conclude that shocks to these prices are not going to be transitory. Summarising the results of the unit root tests, we can conclude that there is no clear pattern to shocks being transitory for the range of coffee prices. For example, we cannot conclude any distinction in terms of variety such as Arabica/Robusta; or geographical region such as Sub-Saharan Africa and Latin America; or economic growth, such as emerging economies Brazil/India compared to poor economies such as Uganda and Honduras. What is important, is that each coffee price, irrespective of country of origin or variety, should be analysed individually, without making sweeping assumptions about the underlying trend of the data; for example, whether the trend contains stochastic trends and whether any underlying deterministic trend is present and if so whether it is positive or negative. Incorrect assumptions can lead to wrong policy recommendations with respect to farmers' welfare and risk management policies, besides the possible wrong modelling techniques with regards to the relationship between coffee prices and other variables such as farmers' household income, consumption and other related variables.

**5. Concluding Remarks**

In this paper we analyse the trend in real coffee prices paid to farmers.
Therefore, on one hand, we are inclined to make a conjecture that supply could be outstripping demand thereby leading to a declining trend; and on the other hand, we could make a conjecture that the variability in prices leads farmers to be risk averse and cut back on investments that could lead to decreased production, causing an upward pressure on prices. The structure of the coffee sector is complex and which of these two opposing arguments are persistent that can cause the long run trend in coffee prices to be significant (or not) is an empirical question, which we aim to address in this paper.

We determine whether the prices are characterised by secular or broken trends and estimate the growth or decline of prices if a trend does exist. We find no evidence of any structural breaks in the price series for the individual coffee prices received by farmers in different coffee producing countries. Apart from Honduras and India (*Robusta*) coffee which show a secular declining trend, none of the prices have a significant trend. The methods we apply are robust and lead us to conclude that at least in the long term, for most countries, real coffee prices do not show any signs of a declining trend, contrary to the grim reports of ICO and Oxfam and other popular media outlets.

In fact almost all of the coffee prices show no significant trends at all (positive or negative), rather, they exhibit a fair amount of variability as shown by the wide confidence intervals. The variability of coffee prices received by producers makes it difficult for them to manage risks. In the face of such variability, an important issue for farmers' is whether any shock to the coffee prices they receive is transitory or not. Our results show that a little less than half of the prices considered in this study conclude that shocks are transitory. No clear pattern emerges with respect to coffee prices classified by variety or country of origin, implying that conclusions about persistence of shocks to coffee prices should be considered on a case by case basis.

What do these results imply for coffee producing countries? Farm size and terrain has favoured mechanization of cultivating and harvesting of coffee in Brazil, allowing capital to substitute the relatively costly hired labour. If costs have been relatively constant over time, then the insignificant trend estimates for coffee prices received by growers in Brazil would suggest that the farmers are no better off, but also not worse off in the long run. Countries such as Colombia and Ethiopia are at a disadvantage competing on production costs, as these origins are associated with superior quality that generally command a premium in the coffee market. Continuously improving quality (rather than yields) and tapping into the high-value market segment can provide a way out for farmers. Governments in producing countries can support their farmers through provision of targeted extension services as well as the establishment of a strong brand related to the origin (as in the case of Colombia).

With very little evidence of any significant long run negative trend, farmers can consider different approaches, such as gradual elimination of coffee trees and replacing them with environmentally friendly shade trees that can enhance the demand for speciality coffee (for which a premium can be charged) and at the same time provide an alternative source of income through the production of timber. This approach could be extended to farm level diversification to allow for other cash crops, or food crops that could be used for domestic consumption. Investment in coffee tree nutrition and disease or pest control can prevent short term positive price shocks. Farmers need to be provided with the liquidity to make such investments and to cover unforeseen costs when coffee has proven unprofitable or when farmers are faced with successive crises, pests, and other difficulties. Indeed, it may well be possible that out-migration occurs and farming other agricultural commodities may not be financially viable, leading farmers to sell their land to cover debts; and permanently emigrating.

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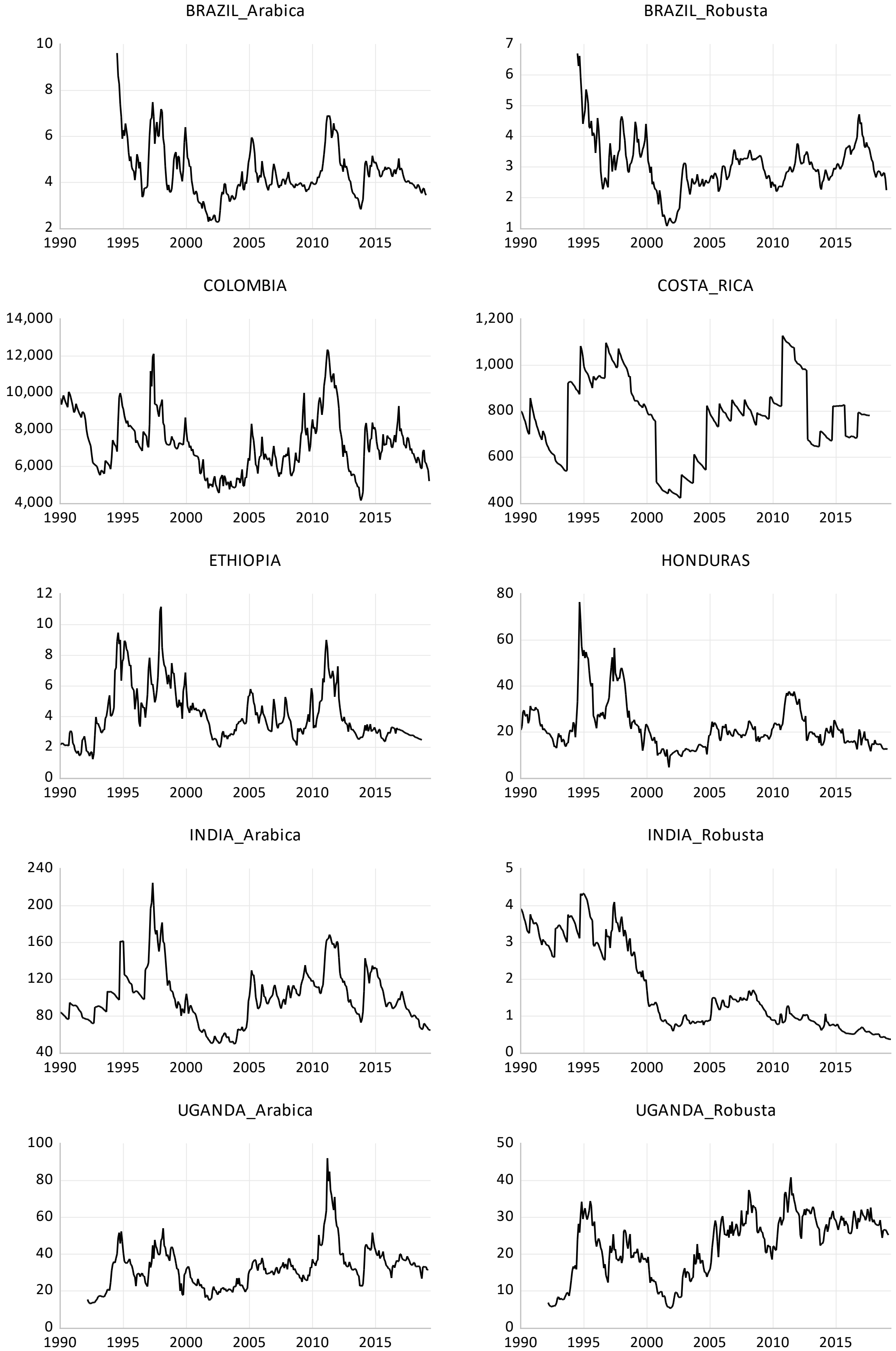
# FIGURES

**Figure 1.** Real coffee prices paid to producers.



## TABLES

**Table 1.** Data and Descriptive Statistics

| | C.V. | Skewness | Kurtosis | Normality |
|---|---|---|---|---|
| Brazil (*Ara*) | 0.257 | 1.098*** | 5.036*** | 110.18*** |
| Brazil (*Rob*) | 0.280 | 0.816*** | 5.717*** | 123.46*** |
| Colombia | 0.222 | 0.674*** | 3.256 | 27.61*** |
| Costa Rica | 0.220 | -0.143 | 2.470* | 5.03* |
| Ethiopia | 0.434 | 1.194*** | 4.246*** | 104.29*** |
| Honduras | 0.470 | 1.775*** | 6.937*** | 409.79*** |
| India (*Ara*) | 0.303 | 0.880*** | 4.149*** | 64.95*** |
| India (*Rob*) | 0.675 | 0.727*** | 2.003*** | 45.69*** |
| Uganda (*Ara*) | 0.355 | 1.415*** | 7.285*** | 357.17*** |
| Uganda (*Rob*) | 0.370 | -0.491*** | 2.241*** | 20.85*** |

**Notes:** Coefficient of variation (C.V.) is given by the ratio of the standard error to the mean of the data series**:** The notation, ***and *denote rejection of the null hypothesis at the 1%, and 10% significance levels. The significance of skewness is measured against the null hypothesis of zero skewness and the significance of kurtosis is measured against the null of no excess kurtosis. The Jarque-Bera test is used to test for normality with the null being a normal distribution.

**Table 2.** Robust sequential tests for structural breaks

| | $SupF^*(1\|0)$ | $SupF^*(2\|0)$ | $SupF^*(3\|0)$ | $UDmaxF_\lambda^*$ | $WDmaxF_\lambda^*$ |
|---|---|---|---|---|---|
| Brazil (*Ara*) | 2.26 | 2.51 | 1.99 | 2.41 | 2.54 |
| Brazil (*Rob*) | 3.85 | 3.30 | 2.39 | 4.01 | 3.76 |
| Colombia | 1.21 | 1.95 | 2.20 | 2.03 | 2.57 |
| Costa Rica | 3.20 | 3.53 | 3.02 | 3.40 | 3.58 |
| Ethiopia | 2.44 | 2.42 | 2.49 | 2.53 | 2.92 |
| Honduras | 1.72 | 4.27 | 3.71 | 4.11 | 4.34 |
| India (*Ara*) | 2.76 | 3.96 | 4.06 | 3.81 | 4.75 |
| India (*Rob*) | 2.64 | 3.03 | 2.91 | 2.92 | 3.41 |
| Uganda (*Ara*) | 2.39 | 3.91 | 4.68 | 4.32 | 5.48* |
| Uganda (*Rob*) | 2.46 | 2.89 | 3.55 | 3.27 | 4.15 |

**Notes:** none of the estimated statistics can reject the null hypothesis of no break (all the estimated test statistics are less than the critical values at the 10% significance level). Whether it be the sequential trend break statistics such as $F_\lambda^*(m|0)$ or the break tests statistics such as the *Dmax* tests, or the modified sequential test statistics – all due to the procedures by Sobriera and Nunes (2016). The only exception is Uganda, where one of the *Dmax* tests is rejected; (the notation, * denotes rejection of the null hypothesis at the 10% significance level) but this is only a borderline case, and is not supported by the sequential tests.

**Table 3.** Robust tests for trend estimation

| | $\hat{\beta}$ (%) | 95% C.I. $\hat{\beta}$ | 90% C.I. $\hat{\beta}$ | Lag | t-stat |
|---|---|---|---|---|---|
| Brazil (*Ara*) | –0.35 | (–1.10, 0.41) | (–1.25, 0.55) | 6 | –0.764 |
| Brazil (*Rob*) | –0.37 | (–1.24, 0.49) | (–1.40, 0.65) | 3 | –0.711 |
| Colombia | –0.11 | (–0.50, 0.27) | (–0.57, 0.34) | 6 | –0.487 |
| Costa Rica | –0.08 | (–0.65, 0.49) | (–0.76, 0.60) | 1 | –0.230 |
| Ethiopia | –0.14 | (–0.71, 0.42) | (–0.81, 0.52) | 17 | –0.423 |

| | | | | | |
|---|---|---|---|---|---|
| Honduras | –0.16*** | (–0.25, –0.07) | (–0.26, –0.06) | 3 | –3.08*** |
| India (*Ara*) | –0.08 | (–0.65, 0.49) | (–0.76, 0.60) | 2 | –0.230 |
| India (*Rob*) | –0.80** | (–1.32, –0.28) | (–1.41, –0.19) | 5 | –2.55** |
| Uganda (*Ara*) | 0.22 | (–0.69, 1.14) | (–0.86, 1.31) | 1 | 0.403 |
| Uganda (*Rob*) | 0.40 | (–0.54, 1.35) | (–0.72, 1.52) | 1 | 0.707 |

**Notes:** The notation, ***and **denote rejection of the null hypothesis at the 1%, and 5% significance levels respectively. The confidence intervals are denoted by C.I. The lag length is chosen according to the modified Akaike Information Criterion (MAIC).

**Table 4.** Results of the bootstrap union tests for unit roots in the presence of non-stationary volatility

| | Test Statistic | $UR_{4A}$ Bootstrap crit. val. [p–val] | $UR_{4B}$ Bootstrap crit. val. [p–val] |
|---|---|---|---|
| Brazil (*Arabica*) | –1.931 | –2.007 [0.132] | –2.006 [0.131] |
| Brazil (*Robusta*) | –1.960 | –2.025 [0.126] | –2.028 [0125] |
| Colombia | –2.035 | –2.044 [0.103] | –2.052 [0.105] |
| Costa Rica | –2.397*** | –1.807 [0.009] | –1.806 [0.009] |
| Ethiopia | –1.789 | –1.920 [0.161] | –1.919 [0.161] |
| Honduras | –2.566** | –1.964 [0.012] | –1.973 [0.013] |
| India (*Arabica*) | –2.275** | –1.977 [0.039] | –1.977 [0.038] |
| India (*Robusta*) | –1.259 | –2.001 [0.711] | –1.894 [0.640] |
| Uganda (*Arabica*) | –2.067* | –1.944 [0.066] | –1.947 [0.065] |
| Uganda (*Robusta*) | –1.665 | –1.991 [0.283] | –1.963 [0.265] |

## Appendix

**Table A.** Description of Data

| Country | Time period | No. of obs. | Currency Unit |
|---|---|---|---|
| Brazil (*Arabica*) | July1994 – January 2019 | 295 | BRL/60KG |
| Brazil (*Robusta*) | July1994 – January 2019 | 295 | BRL/60KG |
| Colombia | January 1990 – April 2019 | 352 | COP/125KG |
| Costa Rica | January 1990 – September 2017 | 333 | CRC/SQ |
| Ethiopia | January 1990 – September 2018 | 345 | ETB/17KG |
| Honduras | January 1990 – June 2019 | 350 | HNL/SQ |
| India (*Arabica*) | January 1990 – May 2019 | 353 | INR/50KG |
| India (*Robusta*) | October 1990 – May 2019 | 353 | INR/50KG |
| Uganda (*Arabica*) | March 1992 – March 2019 | 325 | UGX/KG |
| Uganda (*Robusta*) | March 1992 – March 2019 | 325 | UGX/KG |

Source: International Coffee Organisation