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Breaks, Trends and Unit Roots in Commodity Prices: A Robust Investigation

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Breaks, Trends and Unit Roots in Commodity Prices: A Robust Investigation*

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Abstract

This paper empirically examines the time series behavior of primary commodity prices relative to manufactures with reference to the nature of their underlying trends and the persistence of shocks driving the price processes. The direction and magnitude of the trends are assessed employing a set of econometric techniques that is robust to the nature of persistence in the commodity price shocks, thereby obviating the need for unit root pretesting. Specifically, the methods allow consistent estimation of the number and location of structural breaks in the trend function as well as facilitate the distinction between trend breaks and pure level shifts. Further, a new set of powerful unit root tests is applied to determine whether the underlying commodity price series can be characterized as difference or trend stationary processes. These tests treat breaks under the unit root null and the trend stationary alternative in a symmetric fashion thereby alleviating the procedures from spurious rejection problems and low power issues that plague most existing procedures. Relative to the extant literature, we find more evidence in favor of trend stationarity suggesting that real commodity price shocks are primarily of a transitory nature. We conclude with a discussion of the policy implications of our results.

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1 Introduction

A number of developing countries, especially Sub-Saharan African countries, are dependent on the export of a few commodities for a large share of their export earnings. This level of dependency on commodities makes it important for these countries to understand the underlying trends in real commodity prices to help formulate their economic policies. For instance, for many developing countries when commodity prices experienced a declining trend over time, public expenditure and investment programs had to be abandoned or financed with foreign borrowing (Cashin et al., 2004). As a result, the nature of underlying trends in commodity prices has attracted the attention of researchers and policy makers alike. Not surprisingly, examining the historical data of primary commodity prices relative to manufactured goods has led to a large volume of empirical work investigating the time series properties of commodity prices.

When considering the possibility of the existence of a trend in real commodity prices, past studies have concentrated on the question of whether the prices are trend stationary or difference stationary by employing tests for the presence of a unit root. Perron (1988) noted that the correct specification of the trend function is important in the context of testing for a unit root in the data. If the data contain a unit root, then standard method of least squares to test for the presence of a trend will suffer from severe size distortions. On the other hand, if the data are generated by a trend stationary process but is modeled as a difference stationary process, the tests will be inefficient and will lack power relative to the trend stationary process (see Perron and Yabu, 2009a). The situation is further complicated if one entertains the possibility of structural breaks in the price series. Neglecting a break in an otherwise trend stationary process can cause the spurious appearance of unit root behavior (Perron, 1989) while a neglected trend break in a difference stationary process can lead standard unit root tests to incorrectly suggest the presence of stationarity (Leybourne et al., 1998). Accordingly, it is now common econometric practice to test for the presence of unit roots while allowing for structural changes in the trend function of the underlying time series. A problem with the application of these unit root tests is that they provide little information regarding the existence and number of trend breaks as well as whether the breaks are pure level shifts or affect both the level and slope of the trend function. Besides, the estimates of the break dates that are obtained by minimizing these unit root tests are, in general, not consistent for the true break dates (Vogelsang and Perron, 1998). On the other hand, testing whether a time series can be characterized by a broken trend is complicated by the fact that the nature of persistence in the errors is usually unknown. Indeed, inference based on a structural change test on the level of the data depends on whether a unit root is present while tests based on differenced data can have very poor properties when the series contains a stationary component (Vogelsang, 1998). This circular testing problem underscores the need to employ break testing procedures that do not require knowledge of, or are robust to, the form of serial correlation in the data.

Motivated by these considerations, this paper evaluates the time series properties of primary commodity prices by applying a range of new econometric techniques to 24 primary commodity prices using an updated version of the so-called Grilli-Yang Index over the period 1900-2008. The methods employed enable (i) robust detection of breaks in the level and/or the slope of the trend function as well as a clear demarcation between slope breaks and pure level shifts, (ii) robust estimation of the number of breaks, (iii) robust estimation of the break locations as well as the slope parameters in the regimes identified by the estimated break dates, and (iv) reliable inference regarding the presence of a unit root conditional on the presence/absence of breaks. Our procedures for evaluating the nature of trends are based on testing mechanisms recently proposed in Harvey et al. (2009, HLTb hereafter), Perron and Yabu (2009b, PYb thereafter) and Kejriwal and Perron (2010) as well as their 'no break' counterparts suggested in Harvey et al. (2007, HLTa hereafter) and Perron and Yabu (2009a, PYa hereafter). Further, a new set of powerful unit root tests allowing for structural breaks under both the null and alternative hypotheses proposed by Harris et al. (2009) and Carrion-i-Silvestre et al. (2009) is applied to determine whether the underlying commodity price series can be characterized as difference or trend stationary processes. These tests are not subject to the spurious rejection problems and low power issues associated with most existing procedures.

The rest of the paper is organized as follows. Section 2 provides some background on the issue of trends in commodity prices while Section 3 briefly reviews the relevant empirical literature on commodity price behavior. Section 4 contains a description the econometric methodology including a discussion of the various limitations of commonly employed procedures that motivate our analysis. Section 5 presents the empirical results. Section 6 discusses the policy implications of our analysis, and finally Section 7 concludes. A Supplementary Appendix (attached as a separate document) contains a description of the testing procedures as well as the notation for the various tests.

2 Nature of Commodity Price Behavior

Given the fact that many developing countries are dependent on commodity prices as their main source of income, the issue of trends in commodity prices in relation to manufactures has been of great interest in the trade and development economics literature. Classical economists such as David Ricardo and John Stuart Mill held the view that this trend should be positive as the supply of primary commodities would be constrained by the fixed amount of land while the supply of manufactures would be augmented by technical progress.

However, this view was reversed by Prebisch (1950) and Singer (1950) claiming that commodity prices should decline in relation to manufactured goods in the long run, which was labeled as the well-known Prebisch-Singer Hypothesis (PSH hereafter). Prebisch (1950) offered a supply side theory. He argued that strong labor unions in countries that export manufacturing goods cause wages to increase during times of expansion but prevent wages from falling during times of recessions. In contrast, countries that export primary commodities have weaker labor unions that are not able to increase wages during expansions and cannot prevent wages from falling during times of recessions. Thus, primary commodity prices increase by less than manufacturing goods prices during expansions but fall more during downturns. Thus, the cost of primary commodities rises by less relative to manufactures during upswings and falls by more during downswings, creating a continuous decline in the relative cost of primary commodities. This is caused by the rightward movement in the relative supply schedule. Singer (1950) concentrated on the demand side by considering price and income elasticities. He argued that the manufacturing sector has monopoly power which would prevent technical progress from lowering prices. Moreover, the low income elasticity of demand for primary commodities would cause the decline in primary commodity prices relative to manufacturing goods.

Deaton and Laroque (2003) set out a model that shows prices of commodities in developing countries can be characterized as containing no significant trend by linking commodity price determination to the Lewis (1954) model. Lewis (1954) in his seminal paper notes that in poor countries there is an unlimited supply of labor at a fixed subsistence wage which prevents real wages from increasing. As a result, prices of commodities are unlikely to exceed the cost of production in the long run. Deaton (1999) claims that this is especially true for commodities produced in developing countries. Consequently, prices may deviate in the short run from the long run subsistence wage rate, but because there is an unlimited supply of labor, prices will eventually revert to the constant subsistence level. For agricultural commodities such as cocoa, coffee, copper, cotton and sugar, where there is an abundance of labor at a subsistence wage rate, the Lewis model would seem to be readily applicable. For non-agricultural commodities, such as metals, the owners of the reserves must receive an adequate income for holding their stock of capital. As a result, in the long run arbitrage guarantees that the discounted price of the resource stays constant, or that the commodity prices grow at the rate of interest. The formal theory in this area dates back to influential work by Hotelling (1931). He argued that the price of non-renewable resources in a competitive market would rise at the rate of interest and that the production trajectory would be monotonically declining until the resource is exhausted. However, the empirical evidence contradicts Hotelling by suggesting that the prices of minerals do not exhibit any clear trends. In fact the arbitrage argument applies to the shadow price of resources in the ground, while the observed price series are for the extracted material (Radetski, 2008). Halvorsen and Smith (1991) noted that the price in the ground may decline if extraction costs increase when the overall remaining stock is depleted. If the main component of the observed price happens to be the extraction costs, and if these costs in turn are the subsistence wages of workers employed to carry out the extraction, then Deaton and Laroque (2003) argue that the Lewis model is more relevant than Hotelling for understanding the long run time path of primary commodity prices.

Hotelling's results depend on a number of strong assumptions. For example, in his model, there is no or very little exploration or additions to reserves and no technological change. As discussed earlier, further research is called for to determine if new discoveries and technological change have caused a declining trend in real mineral commodity prices by more than offsetting the effects of mineral depletion over time. Slade (1988), Berck and Roberts (1996), Ahrens and Sharma (1997) and Lee et al. (2006) have focused on the time series properties of natural resource prices. Based on the premise that exogenous shocks can affect the time path of commodity prices, they use advances in unit root testing to examine the time paths of these non-renewable commodity prices. In general, there is not much support of commodity prices showing any positive trend. Following the view put forward by Barnett and Morse (1963), this would be an indication that natural resource scarcity was not an issue. However, as Ahrens and Sharma (1997) have argued, environmental constraints and natural resource abundance may induce a negative trend in prices, or with the introduction of backstop technology causing an inward shift to the demand for natural resources (Heal, 1976). In particular, one needs to note that Hotelling's rule states that the unit price of *unexploited* resource will rise over time at the real rate of interest. This rule was distorted into rephrasing that the price of *exploited* exhaustible resources must rise at the real rate of interest. As a result, the Hotelling rule is likely to have little relevance in practice because the volume of ultimately exploited resources is unknown.

A strand of theoretical research argues that commodity prices should be stationary, due to the biological nature of production, storage and arbitrage (Wang and Tomek, 2007). Deaton and Laroque (1992) apply a rational expectations competitive storage model to study commodity price behavior. Further models have been documented by Williams and Wright (1991) and Peterson and Tomek (2005) for agricultural commodities. Deaton and Laroque (1992) investigate how the storage model can lead to autocorrelated stationary prices. They conclude that the random walk hypothesis for commodity prices seems very implausible as it requires that all price fluctuations be permanent. However, they fail to explain the observed high degree of serial correlation in prices.

3 Literature Review

Early studies ignored the possibility that commodity prices may contain a unit root and structural breaks, simply assuming that the prices were characterized by a trend stationary process. Subsequent studies have explored the use of advances in econometric methodology for their empirical analysis. For instance, Leon and Soto (1997) applied the single break Zivot-Andrews (Zivot and Andrews, 1992) unit root test on primary commodity prices and concluded in favor of structural change while Zanias (2005) and Kellard and Wohar (2006, KW hereafter) employed the Lumsdaine-Papell test (Lumsdaine and Papell, 1997) test that allows for two trend breaks. Zanias (2005) applies the test to the extended aggregate Grilli Yang Index and concluded that the data can be adequately described by a trend stationary process with two intercept shifts. KW conduct a study using the Grilli Yang Index of disaggregated commodity prices over the period 1900-1998 and find that the deterioration of commodity prices has been discontinuous. KW argue that a single trend is a 'summary measure' of several trends which may be positive or negative and that reliance on a single trend may be misleading to policy makers. Accordingly, KW develop a measure to describe the prevalence of a trend.

Ghoshray (2011, Ghoshray hereafter) addresses a gap in the literature by employing the LM unit root test due to Lee and Strazicich (2003, 2004) which allows for structural breaks under the null hypothesis of a unit root and, unlike the Lumsdaine-Papell test, does not suffer from spurious rejection of the null. The study reveals that there are fewer cases, in relation to past studies, of commodities that display negative trends thereby weakening the case for the PSH. While the Lee and Strazicich test offers an improvement over procedures that only allow for breaks under the trend stationary alternative, it does not provide a prescription for how many breaks to include in the specification as well as whether the breaks affect only the level or both the level and slope of the trend function. Moreover, as is also confirmed by their simulation experiments, the proposed break date estimates obtained by minimizing the LM test do not provide particularly reliable approximations to the true break dates.

Harvey et al. (2010) apply novel time series techniques on a unique data set that comprises 25 primary commodities and spans four centuries to test the existence of trends in primary commodity prices. The procedure evaluates the significance of a linear trend based on the method developed by HLTa as well as that of a broken trend based on the procedure advocated in HLTb. Both these methods are robust to whether the commodity prices are characterized as I(1) or I(0) processes. Their results show that 11 commodity prices display a significant negative trend over the entire sample or some fraction of it thereby showing some support for the PSH. However, evidence of a negative trend in commodity prices is weakened when applying the same methods to the Grilli Yang Index where only

7 commodity prices display a significant negative trend over the entire sample or some post-break subsample of the time span.

The subject of whether significant trends exist for primary commodities has led to much debate as it has been used to explain the widening gap between developed and less developed countries. The evidence has been mixed which leads to serious policy implications as to whether developing countries should specialize in primary commodity exports. The upshot is that countries which rely on the exports of primary commodities need to understand the nature of commodity prices for their economic policies to be effective.

4 Econometric Methodology

This section borrows heavily from section 2 of Kejriwal and Lopez (2012) who employ a variant of the methodology described below to investigate the time series properties of per-capita output for a set of OECD countries. The bulk of the existing empirical literature examining the trends in commodity prices has focused on the application of unit root tests allowing for structural breaks in the trend function followed by the estimation of a level or first-differenced specification according to whether a unit root is present or not. These tests are generally obtained by minimizing the t-statistic on the unit root parameter over the set of permissible break dates or computing this t-statistic at the break date that minimizes (or maximizes) the t-statistic associated with the break parameter (or maximizes its absolute value). In order to provide the motivation for the econometric methodology advocated in this paper, it is useful to first discuss the potential drawbacks associated with the testing procedures that have typically been employed by existing studies.

First, the tests provide little information regarding the existence or number of trend breaks. At an intuitive level, it seems more natural to be first able to ascertain if breaks are at all present before proceeding to conduct unit root tests allowing for such breaks. In the absence of breaks, these tests suffer from low power due to the inclusion of extraneous break dummies thereby potentially leading the researcher to estimate a differenced specification when a level specification is in fact more appropriate. Indeed, as stressed by Campbell and Perron (1991), proper specification of the deterministic components is essential to obtaining unit root tests with reliable finite sample properties.

Second, simulation evidence presented in Vogelsang and Perron (1998) and Lee and Strazicich (2001) suggests that the estimates of the break dates obtained by minimizing/maximizing these unit root tests over all possible break dates are unlikely to provide consistent estimates of the true break dates.

Third, the unit root tests typically employed suffer from serious power and size distortions due to the asymmetric treatment of breaks under the null and alternative hypotheses. If breaks are indeed present, this information is not exploited to improve the power of the testing procedure. More importantly, these tests are subject to a spurious rejection problem when breaks are present under the unit root null hypothesis. Essentially, the problem is the presence of nuisance parameters related to the trend function under the null hypothesis. To illustrate this problem, consider the case where there is a single break ($K = 1$ in (4) below) under the unit root null ($\alpha = 1$ in (5) below). Then, under the null, differencing (4) yields

$$\Delta y_t = \beta_0 + \mu_1 \Delta DU_{1t} + \beta_1 DU_{1t} + v_t \tag{1}$$

while under the alternative (assuming an AR(1) structure), we have

$$y_t = c_0 + c_1 t + c_2 DU_{1t} + c_3 DT_{1t} + c_4 y_{t-1} + e_t \tag{2}$$

Thus, the testing regression nesting both (1) and (2) takes the form

$$y_t = d_0 + d_1 t + d_2 \Delta DU_{1t} + d_3 DU_{1t} + d_4 DT_{1t} + d_5 y_{t-1} + \varepsilon_t \tag{3}$$

Observe that omitting the impulse dummy ΔDU_{1t} in (3) will make the unit root test statistic diverge to infinity as μ_1 and/or β_1 in (1) increase(s). Then, in this context, μ_1 and/or β_1 are nuisance parameters. The resulting tests are not pivotal and spurious rejections occur under the null when the critical values derived assuming no break ($\mu_1 = \beta_1 = 0$) are employed. It is important to emphasize, however, that using the testing regression (3) will induce the same problem for currently popular endogenous break tests of Zivot and Andrews (1992), Perron (1997) and Lumsdaine and Papell (1997). The parameters μ_1 and β_1 remain nuisance parameters even when the regression model estimated is given by (3) (for more details on this problem, see Perron and Vogelsang, 1993; Perron, 2006; Lee and Strazicich, 2001).

Fourth, based on the prescription of unit root tests, the existing procedures often estimate a level specification and evaluate the joint significance of the intercept and slope dummies. However, a joint test is likely to conclude in favor of unstable growth rates even if the series has undergone a pure level shift, thereby making the interpretation of such tests quite difficult in practice. Thus, if the objective is to distinguish between changes in the level and the slope, it is essential to test for the stability of the slope parameter while allowing the intercept to vary across regimes and, conditional on the absence of slope shifts, test for level shifts.

Fifth, another common strategy is to start (before testing for a unit root) with a general level specification that incorporates both a changing slope as well as a changing intercept and then evaluate the significance of the individual t-statistics on the dummy variables. Depending on the outcome, the relevant model is estimated and used as the alternative model when testing for a unit root. There are two problems with such an approach. First, the limit distributions of the slope coefficient dummy estimates are different depending on whether a unit root is present so that prior information regarding the existence of a unit root is essential to validate significance based on t-statistics. Second, in the presence of a slope shift, the level shift parameters are not identified regardless of whether the noise component is stationary or not (see Hatanaka and Yamada, 1999; Perron and Zhu, 2005).

Our econometric methodology is aimed at addressing each of the limitations discussed above. The most general model considered can be described as:

$$y_t = \mu_0 + \beta_0 t + \sum_{i=1}^{K} \mu_i DU_{it} + \sum_{i=1}^{K} \beta_i DT_{it} + u_t, \quad t = 1, ..., T \tag{4}$$

$$u_t = \alpha u_{t-1} + v_t, \quad t = 2, ..., T, \ u_1 = v_1 \tag{5}$$

where $DU_{it} = I(t > T_i),\ DT_{it} = (t - T_i)I(t > T_i), i = 1, ..., K$. A break in the trend occurs at time $T_i = [T\lambda_i]$ when $\beta_i \neq 0$. The date of the breaks, T_i, and the number of breaks, K, are treated as unknown. The error u_t is allowed to be either $I(0)$ ($|\alpha| < 1$) or $I(1)$ ($\alpha = 1$). The stochastic process $\{v_t\}$ is assumed to be stationary (but not necessarily i.i.d. thereby permitting a general error structure for u_t). We are interested in the null hypothesis H_0: $\beta_i = 0$ against the alternative hypothesis H_1: $\beta_i \neq 0$.[1]

The first step tests for one structural break (that is $K = 1$ in (4)) in the slope of the trend function using procedures that are robust to the stationarity/non-stationarity properties of the data (HLTb and PYb). The tests employed are designed to detect a break in slope while allowing the intercept to shift. A rejection by these robust tests can therefore be interpreted as a change in the growth rate regardless of whether the level has changed.[2] Given evidence in favor of a break by either of the single break tests, we then proceed to test for one versus two slope breaks (that is $K = 2$ in (4)) using the extension of PYb proposed by Kejriwal and Perron (2010). Again, this latter test allows us to distinguish between one and two breaks while being agnostic to whether a unit root is present. Given the number of sample observations available (109), we allow for a maximum of two breaks in our empirical analysis. While this may appear restrictive, allowing for a large number of breaks is not an appropriate strategy if one wants to determine if a unit root is present. The reason is that a unit root process can be viewed

[1]Strictly speaking, the null hypothesis must be re-stated as H_0: $\mu_i = \beta_i = 0$ to obtain pivotal limiting distributions for the test statistics (see section 4.2 in HLTb).

[2]A potential strategy in this case to dissociate a level from a slope shift could be to use a t-statistic to test for the significance of the level shift parameter. Such a strategy is, however, flawed since, as shown in Perron and Zhu (2005), the level shift parameter is not identified in this case.

as a limiting case of a stationary process with multiple breaks, one that has a break (permanent shock) every period. Further, as discussed in Kejriwal and Perron (2010), the maximum number of breaks should be decided with regard to the available sample size. Otherwise, sequential procedures for detecting trend breaks will be based on successively smaller data subsamples (as more breaks are allowed) thereby leading to low power and/or size distortions. It is therefore important to allow for a sufficient number of observations in each segment and choose the maximum number of permissible breaks accordingly.[3] It is useful to note that, as in Bai and Perron (1998, section 4.3), our procedure is not a purely sequential one so that at each step the break dates are estimated by minimizing the *global* sum of squared residuals.

Conditional on the presence of a stable slope at the initial step (that is $\beta_i = 0$ in (4) for $i = 1, ..., K$), the focus becomes potential changes in the level of the trend and the hypotheses tested are H_0: $\mu_i = 0$ against the alternative hypothesis H_1: $\mu_i \neq 0$. Harvey et al. (2010) propose a test for detecting multiple level breaks that is robust to the unit root/stationarity properties of the data.[4] A rejection by this robust test can therefore be interpreted as changes in the level of the series. These authors also develop a sequential procedure which allows reliable estimation of the number of level breaks.

Given the demarcation between pure level breaks and those that affect both the level and slope, we proceed to estimate the break dates. In models that involve at least one slope shift, the break date estimators are obtained by minimizing the sum of squared residuals obtained by applying ordinary least squares to (4). As shown in Perron and Zhu (2005), these estimates are consistent regardless of whether the errors are $I(1)$ or $I(0)$. In models with pure level shifts, we are not aware of a unified procedure that consistently estimates the break dates in both $I(1)$ and $I(0)$ cases. Hence, in these models, we pretest for a unit root and obtain the break date estimates using the procedure suggested by Harvey et al. (2010) in the unit root case and by minimizing the sum of squared residuals from the level specification in the stationary case.

Having obtained the break date estimates, we apply the robust procedures proposed by HLTa and PYa to test for trend significance in the subsamples determined by these estimates for models involving slope shifts. These procedures are the "no break" counterparts of the HLTb and PYb procedures respectively. With no breaks in either level or slope, these procedures are applied to the full

[3] If a unit root is indeed present, the estimates of the break dates (obtained from the first-differenced specification) from an underspecified model are consistent for those break dates inserting which allow the greatest reduction in the sum of squared residuals and therefore correspond to the most dominant breaks in this sense (see Chong, 1995, Bai and Perron, 1998).

[4] The level breaks are modeled as local to zero in the $I(0)$ case and as increasing functions of sample size in the $I(1)$ case.

sample. In models with pure level shifts, trend significance can be assessed using a first-differenced specification if the unit root pretest indicated the presence of a unit root or using a level specification otherwise.

Given evidence in favor of instability in the level and/or slope (that is $\beta_i \neq 0$ and/or $\mu_i \neq 0$ in (4) for at least one $i = 1, ..., K$), we apply a new class of unit root tests which allows for breaks in the level and the slope under both the null and alternative hypotheses (Harris et al., 2009 and Carrion-i-Silvestre et al., 2009).[5] Such a symmetric treatment of breaks alleviates these unit root tests from size and power problems that plague tests based on search procedures (for instance, Zivot and Andrews, 1992). If no evidence is found of instability either in the level or in the slope, we apply standard (no break) unit root tests developed by Elliott et al. (1996) and Ng and Perron (2001).

There is always a potential power issue associated with unit root tests allowing for multiple breaks, given that a unit root process is observationally equivalent to a stationary process with multiple breaks in the limit. Simulation evidence presented in Carrion-i-Silvestre et al. (2009) shows that the tests allowing upto two breaks have decent finite sample power when the data generating process is driven by one or two breaks. Indeed, they have much better properties than unit root tests based on search procedures given that they exploit information regarding the presence of breaks.

5 Empirical Results

An extended data set of the original Grilli Yang Index is employed in this study. The data was updated according to the method put forward by Pfaffenzeller et al. (2007). The data set consists of 24 primary commodity prices measured annually over the period 1900 – 2008 and deflated by the Manufacturing Unit Value (MUV) index.

The initial step of the analysis tests for the presence and the number of breaks in the trend function without making an assumption regarding whether the errors are stationary or not. For the detection of slope breaks, we employ the sequential testing procedure advocated in Kejriwal and Perron (2010) while for pure level breaks, the procedure recommended by Harvey et al. (2010) is applied. The results are reported in Table 1. The test statistics $ExpW$ and t_η are the PYb and HLTb tests for the null hypothesis of no slope break respectively. The statistic $ExpW(2|1)$ is the Kejriwal and Perron (2010) sequential test of one versus two slope breaks while U is the Harvey et al. (2010) test for the null of no level breaks.

[5] Note that Perron (1989) devised unit root testing procedures that are invariant to the magnitude of the shift in level and/or slope but his analysis was restricted to the known break date case.

Table 1: Robust Tests for Breaks in Trend

	Slope Breaks				Level Breaks	
Commodity\Test	$ExpW$	$ExpW(2\|1)$	t_η	#Breaks	U	#Breaks
Coffee	-0.23	-0.20	3.16*	1	–	–
Cocoa	-0.07	-0.11	1.39	0	0.51	0
Tea	3.12*	1.80*	1.74	2	–	–
Rice	-0.25	–	2.07	0	0.32	0
Wheat	0.37	–	1.51	0	0.33	0
Maize	4.25*	3.89*	3.52*	2	–	–
Sugar	-0.28	–	1.97	0	0.41	0
Beef	-0.22	-0.22	1.47	0	0.48	0
Lamb	-0.26	-0.13	1.16	0	0.45	0
Banana	-0.14	1.32	3.05*	1	–	–
Palmoil	0.91	3.15*	3.13*	2	–	–
Cotton	13.53*	0.97	5.38*	1	–	–
Jute	2.77*	0.32	5.33*	1	–	–
Wool	2.19*	16.79*	1.72	2	–	–
Hides	0.56	–	2.16	0	0.50	0
Tobacco	3.28*	634.40*	2.73	2	–	–
Rubber	0.21	200.89*	3.35*	2	–	–
Timber	-0.19	–	2.41	0	0.45	0
Copper	0.02	–	1.76	0	0.34	0
Aluminium	0.05	-0.14	3.65*	1	–	–
Tin	-0.26	–	1.47	0	0.51	0
Silver	-0.03	0.14	1.86	0	0.38	0
Lead	-0.17	–	1.77	0	0.37	0
Zinc	0.11	–	2.47	0	0.41	0

Here '*' denotes significance at the 10% level.

Table 2: HLTa Trend Coefficient Estimates: Commodities with Slope Breaks (in %)

	One Break			Two Breaks				
	$\hat\beta_0$	$\sum_{i=0}^{1}\hat\beta_i$	D1	$\hat\beta_0$	$\sum_{i=0}^{1}\hat\beta_i$	$\sum_{i=0}^{2}\hat\beta_i$	D1	D2
Coffee	0.74 [-2.01,3.49]	-1.87 [-3.82,0.07]	1949	-	-	-	-	-
Tea	-	-	-	-1.89 [-4.25,0.46]	1.34 [0.15,2.53]	-1.83 [-3.58,-0.07]	1917	1957
Maize	-	-	-	0.45 [-1.00,1.89]	-5.29 [-5.78,-4.81]	3.60 [-2.67,9.86]	1974	1991
Banana	1.49 [-1.24,4.22]	-0.34 [-1.04,0.36]	1925	-	-	-	-	-
Palmoil	-	-	-	-0.39 [-0.88,0.11]	-6.43 [-10.34,-2.53]	4.16 [-3.64,11.96]	1974	1991
Cotton	0.51 [-1.53,2.56]	-2.71 [-3.07,-2.35]	1945	-	-	-	-	-
Jute	0.39 [-2.43,3.21]	-1.67 [-2.65,-0.68]	1946	-	-	-	-	-
Wool	-	-	-	0.71 [-1.01,2.43]	-3.93 [-4.92,-2.94]	3.46 [2.92,3.99]	1951	1991
Tobacco	-	-	-	4.37 [0.50,8.24]	0.95 [0.58,1.31]	-0.79 [-1.52,,-0.06]	1922	1969
Rubber	-	-	-	-8.15 [-17.24,0.93]	6.70 [-3.71,17.10]	-1.07 [-3.97,1.82]	1932	1951
Aluminium	-2.13 [-4.50,0.25]	-0.30 [-1.49,0.88]	1941	-	-	-	-	-

Table 3: PYa Trend Coefficient Estimates: Commodities with Slope Breaks (in %)

	One Break			Two Breaks				
	$\hat{\beta}_0$	$\sum_{i=0}^{1}\hat{\beta}_i$	D1	$\hat{\beta}_0$	$\sum_{i=0}^{1}\hat{\beta}_i$	$\sum_{i=0}^{2}\hat{\beta}_i$	D1	D2
Coffee	0.53 [-0.60,1.66]	-1.85 [-2.79,-0.90]	1949	-	-	-	-	-
Tea	-	-	-	-1.84 [-3.10,-0.58]	1.37 [0.33,2.42]	-2.25 [-2.87,-1.62]	1917	1957
Maize	-	-	-	-0.25 [-0.64,0.15]	-4.93 [-4.93,-4.93]	3.64 [-4.44,11.71]	1974	1991
Banana	1.16 [0.59,1.73]	-0.64 [-0.87,-0.40]	1925	-	-	-	-	-
Palmoil	-	-	-	-0.33 [-0.70,0.04]	-7.24 [-7.24,-7.24]	2.36 [2.36,2.36]	1974	1991
Cotton	0.55 [-2.44,3.54]	-2.66 [-3.01,-2.31]	1945	-	-	-	-	-
Jute	0.47 [-0.58,1.53]	-2.14 [-2.75,-1.54]	1946	-	-	-	-	-
Wool	-	-	-	0.62 [-0.05,1.29]	-3.52 [-3.80,-3.24]	3.47 [3.47,3.47]	1951	1991
Tobacco	-	-	-	4.40 [0.36,8.44]	0.91 [0.61,1.20]	-0.82 [-1.21,,-0.42]	1922	1969
Rubber	-	-	-	-8.40 [-17.99,1.18]	7.65 [-3.71,19.01]	-0.92 [-5.25,3.41]	1932	1951
Aluminium	-2.02 [-2.83,-1.22]	-0.43 [-0.99,0.14]	1941	-	-	-	-	-

Table 4: Stable Trend Coefficient Estimates (in %)

Commodity\Estimate	$\hat{\beta}_0$ (HLTa)	$\hat{\beta}_0$ (PYa)
Cocoa	-0.41 [-2.96,2.14]	-0.50 [-3.59,2.59]
Rice	-0.65 [-1.98,0.69]	-0.95 [-1.27,-0.63]
Wheat	-0.38 [-1.33,0.57]	-0.69 [-0.92,-0.46]
Sugar	-1.06 [-1.79,-0.34]	-1.08 [-1.51,-0.64]
Beef	1.55 [0.23,2.88]	1.50 [0.92,2.09]
Lamb	1.77 [1.10,2.43]	1.80 [1.22,2.37]
Hides	-0.70 [-1.91,0.52]	-0.73 [-1.02,-0.43]
Timber	0.94 [0.45,1.43]	1.00 [0.74,1.26]
Copper	0.17 [-1.45,1.80]	-0.23 [-0.68,0.22]
Tin	0.57 [-1.67,2.82]	0.23 [-0.68,0.22]
Silver	0.48 [-1.88,2.84]	0.51 [-2.38,3.40]
Lead	0.16 [-1.42,1.75]	0.43 [-2.45,3.32]
Zinc	0.11 [-0.18,0.41]	0.10 [-0.10,0.30]

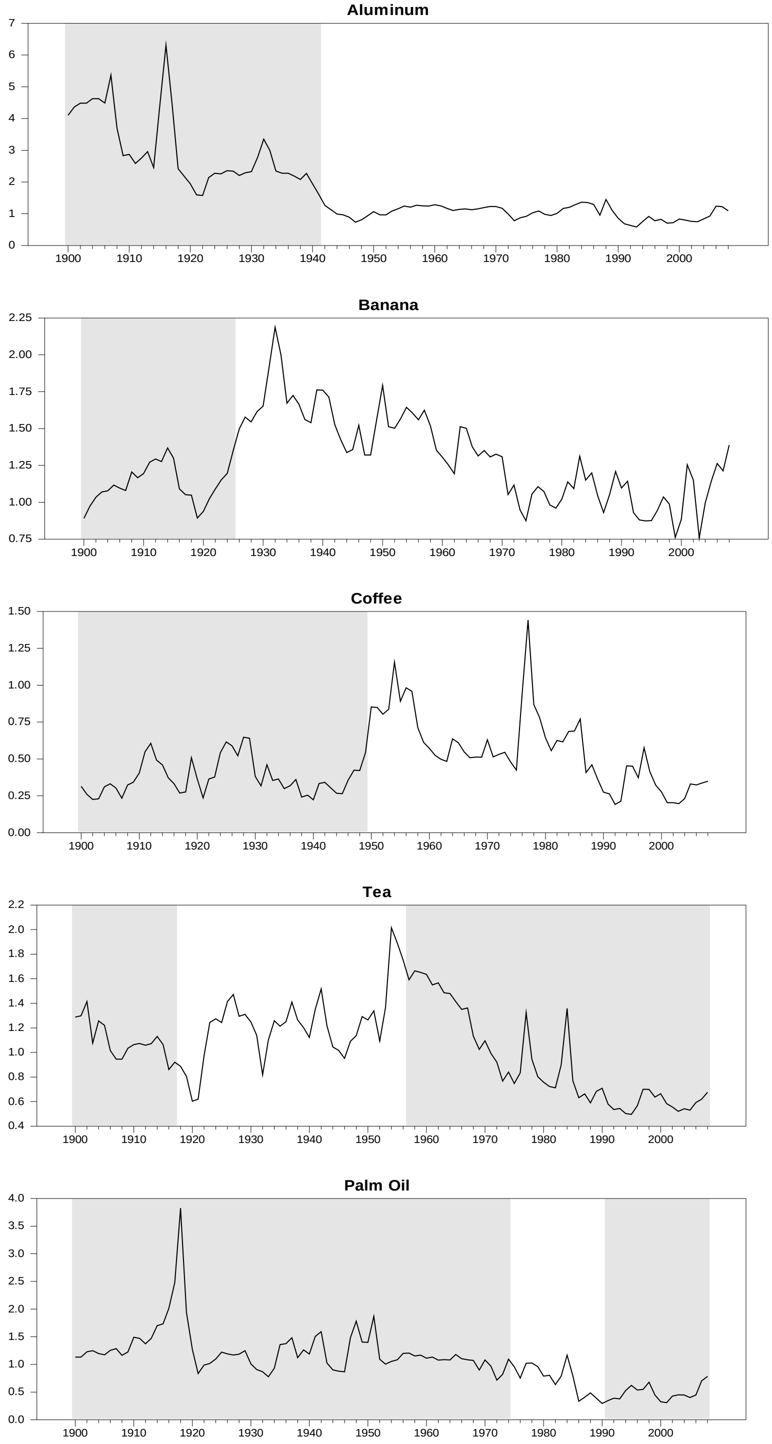

Figure 1: A Selection of Commodity Prices with Structural Breaks

The results show that 11 out of 24 commodities contained either one or two breaks in the slope. Out of the 11 commodities, 6 commodities were found to exhibit two structural breaks while the remaining 5 contained a single structural break. When applying the tests for pure level shifts we find no evidence of any breaks. Consequently no structural breaks were found for the remaining 13 commodity prices.

Having determined that structural breaks are present in 11 of the 24 commodity prices, we estimate the trend coefficients over the regimes that are delineated by the estimated break dates. The regime-specific trend estimates as well as the associated 90% confidence intervals obtained using the approaches suggested by HLTa and PYa are presented in Tables 2 and 3 respectively.

A number of interesting features with respect to the characterization and estimation of trends appear from these results. In the case of coffee where we find a single structural break in 1949, the sign of the estimates of the trend coefficient using the two methods differ. While the regime between 1900 to 1949 for both methods results in an insignificant trend estimate, the PYa method finds the trend coefficient to be significantly negative in the regime spanning 1950 to 2008. In contrast, the HLTa method finds the trend in this latter regime to be insignificant. Similarly, in the case of tea and aluminum, we find that for the first regime the trend estimate is insignificant using the HLTa method, whereas it is negative using the PYa method. For palm oil, the difference in trend estimates is found in the third regime when comparing the two approaches. The overall conclusion of the trend function in the case of banana is quite different using the two approaches. Under the HLTa method, we conclude that allowing for a single structural break the trend is insignificant over the entire sample. This result is in stark contrast to the PYa approach that finds significant trends in both regimes. In the first regime, the trend is found to be positive, while the trend is found to be negative in the second regime. The remaining 6 commodities show that the sign estimates of the trend for each regime are the same. One potential explanation for the observed difference in results from employing the two methods is that the PYa procedure generally has higher power than the HLTa procedure, as has been demonstrated through simulation experiments in PYa. Consequently, the confidence intervals based on PYa are usually shorter than those based on HLTa, as is evident from a comparison of Tables 2 and 3.

Figure 1 plots a selection of commodity prices that experience structural breaks and where the regimes provide different confidence intervals (computed by the HLTa and PYa methods) for the trend estimates across sub-samples. The successive regimes that are obtained from the estimated break dates are highlighted by the shaded and unshaded regions of the graph. One can observe by eyeballing the data over the different regimes, that where a difference in the sign of the estimated trend coefficient is found for the two methods, the PYa estimates of the trend appears more plausible.

For the 13 commodities that do not experience any structural breaks, we proceed to estimate the trend function employing the HLTa and PYa methods. The results of the estimates of the trend function are given in Table 4. Except for rice, wheat and hides, both the HLTa and PYa methods produce estimates of the trend function of the same sign. Out of the 13 commodities, 6 commodities do not show any evidence of a significant positive or negative trend. We can conclude that a negative trend exists for rice, wheat, sugar and hides whereas for beef, lamb and timber, the sign of the trend function is positive. Note that using the HLTa method, there is no evidence of a significant negative trend in rice, wheat and hides, but using the PYa method we find a significant negative trend.

Table 5: Prevalence of Trends

	HLTa			PYa		
Commodity	$\Psi(-)$	$\Psi(+)$	$\Psi(.)$	$\Psi(-)$	$\Psi(+)$	$\Psi(.)$
Coffee	0.000	0.000	1.000	0.542	0.000	0.458
Cocoa	0.000	0.000	1.000	0.000	0.000	1.000
Tea	0.477	0.358	0.165	0.642	0.358	0.000
Rice	0.000	0.000	1.000	1.000	0.000	0.000
Wheat	0.000	0.000	1.000	1.000	0.000	0.000
Maize	0.156	0.000	0.844	0.156	0.000	0.844
Sugar	1.000	0.000	0.000	1.000	0.000	0.000
Beef	0.000	1.000	0.000	0.000	1.000	0.000
Lamb	0.000	1.000	0.000	0.000	1.000	0.000
Banana	0.000	0.000	1.000	0.762	0.238	0.000
Palmoil	0.156	0.000	0.844	0.156	0.156	0.688
Cotton	0.578	0.000	0.422	0.578	0.000	0.422
Jute	0.569	0.000	0.431	0.569	0.000	0.431
Wool	0.477	0.367	0.156	0.477	0.367	0.156
Hides	0.000	0.000	1.000	1.000	0.000	0.000
Tobacco	0.358	0.642	0.000	0.358	0.642	0.000
Rubber	0.000	0.000	1.000	0.000	0.000	1.000
Timber	0.000	1.000	0.000	0.000	1.000	0.000
Copper	0.000	0.000	1.000	0.000	0.000	1.000
Aluminium	0.000	0.000	1.000	0.385	0.000	0.615
Tin	0.000	0.000	1.000	0.000	0.000	1.000
Silver	0.000	0.000	1.000	0.000	0.000	1.000
Lead	0.000	0.000	1.000	0.000	0.000	1.000
Zinc	0.000	0.000	1.000	0.000	0.000	1.000

Following KW, we synthesize the results from the analysis of the above tables by constructing a measure of the prevalence of trends. For each commodity we calculate $\Psi(-) = \lambda(-)/T$, where $\lambda(-)$ equals the number of years that a statistically significant negative trend exists. In the same way, we calculate the measure of the prevalence of a positive trend [defined as $\Psi(+) = \lambda(+)/T$] and trendless behavior [defined as $\Psi(.) = 1 - \Psi(-) - \Psi(+)$]. Table 5 displays the relative measure results for all 24 commodities.

Table 6a: Unit Root Tests with Slope Breaks

Commodity\Test	MZ_{α}^{gls}	MSB^{gls}	MZ_t^{gls}	P_T^{gls}	MP_T^{gls}	H
Coffee	-20.32*	0.16*	-3.15*	6.80	6.63	-3.44*
Tea	-23.72*	0.14*	-3.43*	8.41*	8.17*	-
Maize	-25.15	0.14*	-3.46	9.21	8.84	-
Banana	-5.85	0.29	-1.71	19.30	16.75	-1.67
Palmoil	-34.12*	0.12*	-4.10*	6.14*	5.70*	-
Cotton	-25.69*	0.14*	-3.57*	6.31*	6.06*	-1.83
Jute	-19.19*	0.16*	-3.09*	6.39*	6.24*	-2.87
Wool	-7.01	0.25	-7.01	23.53	18.98	-
Tobacco	-23.66	0.15	-3.43*	8.77	8.48	-
Rubber	-18.44*	0.16	-2.97	10.07	9.41	-
Aluminium	-15.39	0.18	-2.71	9.50	9.62	-2.86

Here '*' denotes significance at the 10% level.

The prevalence of trends according to the HLTa method shows that 8 out of the 24 commodities display at least one significant negative trend segment in contrast to the 13 commodities indicated by the PYa method. For the HLTa method, only 1 commodity (sugar) shows a significant negative trend for the entire sample, whereas for the PYa method we find so for three further commodities: rice, wheat and hides. Using the HLTa approach, no other commodity shows a negative trend for at least 70% of the sample period. If we were to consider at least 50% of the sample period, then there are two further commodities (cotton, jute). Contrasting this result with the PYa method, for 70% of the sample period a negative trend is prevalent in one further commodity (banana), and the number rises by four (coffee, tea, cotton, jute) if we were to consider at least 50% of the sample. Comparing our results with KW and Ghoshray, we find that there is less evidence of a prevalent negative trend. While KW find 8 out of 24 commodities contain a prevalent (that is, at least 70% of the sample period) negative trend, and Ghoshray finds only 6, our results show one commodity (sugar) using the HLTa and a further four (tea, wheat, banana, hides) commodities using the PYa method. Harvey et al. (2010), in their study of the Grilli Yang Index of commodity prices, find 3 commodities (aluminum, rice, sugar) contain a stable negative trend and a further 4 commodities (banana, coffee, jute, lead) contain a post break negative trend. Comparing our results with Harvey et al. (2010), we find a match for 2 commodities (rice, sugar) that are found to contain a stable negative trend, and a match for 3 commodities (coffee, banana, jute) that contain a broken negative trend. It needs to be noted, however, that the results may vary due to the different sample sizes chosen in these studies.

Table 6b: Unit Root Tests without Breaks

Commodity\Test	MZ_{α}^{gls}	MSB^{gls}	MZ_{t}^{gls}	P_{T}^{gls}	MP_{T}^{gls}
Cocoa	-8.82	0.24	-2.09	10.90	10.37
Rice	-19.61*	0.16*	-3.07*	5.17*	5.03*
Wheat	-17.05*	0.16*	-2.82*	6.38*	5.94*
Sugar	-20.41*	0.16*	-3.19*	4.49*	4.48*
Beef	-15.31*	0.18*	-2.75*	6.09*	6.07*
Lamb	-16.35*	0.17*	-2.86*	5.58*	5.58*
Hides	-4.12	0.34	-1.40	22.76	21.74
Timber	-20.70*	0.15*	-3.20*	4.62*	4.49*
Copper	-11.53	0.19	-2.18	9.52	9.05
Tin	-11.78	0.20	-2.40	7.92	7.88
Silver	-7.58	0.24	-1.83	12.97	12.31
Lead	-14.64*	0.18*	-2.58	7.02	6.98
Zinc	-32.05*	0.12*	-3.99*	2.92*	2.90*

Here '*' denotes significance at the 10% level.

Finally, we conclude our empirical analysis by examining whether the commodity prices are characterized by difference or trend stationary processes. Following the results in Table 1 where we determine whether or not the prices contain structural breaks, we employ a new class of unit root tests proposed by Harris et al. (2009) [denoted by H] and Carrion-i-Silvestre et al. (2009) [the M-tests] which allow for breaks in the slope under both the null and alternative hypotheses. For commodities with no breaks in either level or slope, the standard (no break) unit root tests proposed by Elliott et al. (1996) and Ng and Perron (2001) [the no break M-tests]. The results of the tests are reported in Tables 6a and 6b.

The results show that 16 of the 24 commodity prices can be classified as a trend stationary process. For the remaining 8 prices (cocoa, banana, wool, hides, copper, aluminum, tin and silver) the null hypothesis of a unit root was not rejected. Our results show that for the 16 commodities characterized as a trend stationary process, exogenous shocks to these commodity prices are likely to be transitory in nature. On the other hand, for the 8 commodity prices found to be difference stationary one may conclude that any exogenous shocks to these commodities are likely to be relatively more persistent than commodities that are found to display trend stationary behavior.

Overall, our results show some clear differences when compared to recent studies by Ghoshray and KW. When considering the issue of a prevalent trend, we find that there is less evidence of a prevalent negative trend (that is, at least 70% of the sample period); our results show one commodity (sugar) using the HLTa and a further four (tea, wheat, banana, hides) commodities using the PYa method display a prevalent negative trend. While KW find 8 commodities out of 24 to be characterized by a prevalent negative trend, Ghoshray finds such a feature to be relevant for only 6 commodities. Further, our results using the HLTa and PYa methods indicate evidence favoring a positive trend for 3 commodities (beef, lamb and timber). In contrast, a prevalent positive trend is obtained by Ghoshray for only one commodity (timber) and KW for only two commodities (tin and zinc). Regarding the presence of unit roots, Ghoshray and KW find 11 and 10 commodities to be difference stationary, respectively, while our analysis indicates that there are fewer commodities (8 in total) that can be classified as difference stationary. Comparing with Ghoshray, a match is found for only 4 commodities (being cocoa, aluminum, hides, silver) whereas with KW a match is found for 3 commodities (being cocoa, banana, copper). Two commodities (wool, tin) do not match either of the two studies. However, one must note that the sample size chosen in this study is slightly longer than Ghoshray and more so compared to KW.

6 Policy Implications

Tables 1-4 describe whether the primary commodity prices chosen in this study experience any structural breaks and if so, the date/timing of such breaks. A key contribution of the paper is that the number of breaks, whether in level only or in both level and slope, is consistently estimated without requiring any a priori knowledge regarding whether the noise component is stationary or not. The timing of structural breaks also plays a very important role in determining the exact nature of the trends within regimes that are demarcated by the estimated structural breaks. This result is in line with the view put forth by Bloch and Sapsford (2000) that when estimating the trend relationship, the choice of break dates can lead to different conclusions on the PSH. The break dates estimated in this paper coincide with a number of significant events that took place for primary commodities. A number of break dates are observed to have occurred in the 1940s which may be a result of the Great Depression of the 1930s which brought about a collapse of international trade and a surge in bilateral trade agreements and import controls (Ocampo and Parra, 2007). Some break dates occur after World War I, (tea, banana, tobacco) which can be explained partly as a result of the retreat towards autarky and partly because the era of low transportation costs gradually came to an end (Hadass and Williamson, 2003).

Table 5 summarizes the prevalence of trends. The prevalence of a negative trend is found to be marginally lower in comparison to recent studies by KW and Ghoshray. Our study finds a negative trend to exist only for rice, wheat, hides and sugar over the entire time span and a prevalent negative trend (for more than 70% of the time span) is found for banana. Given the relatively few commodities that experience a prevalent negative trend, the case for the PSH is weakened. Note however, apart from wheat, a negative trend is mainly found for commodities exported by developing countries. This finding would translate as an important point for policy makers to consider especially when a country is highly dependent on these commodities. When the real price for these countries' dominant export falls over a period of time, it may translate into a decline in the international purchasing power which in turn could lead to a deficit in the balance of payments. This could lead countries to borrow which in turn may exacerbate their debt problems. The World Bank has put forward policy recommendations for developing countries experiencing deteriorating terms of trade, which include diversification into other exports of primary commodities and also moving away from primary commodities that are in oversupply. For instance, some countries in the Pacific Rim, such as Hong Kong, Singapore, South Korea and Taiwan, have benefited from export diversification policies, whereas other developing countries, especially Sub-Saharan African countries, have been unable to break into global manufactures or services. Interestingly, three commodity prices (timber, beef and lamb) show a positive trend over the entire sample. This result contrasts sharply with that of KW and Ghoshray. According to the PYa estimate of the trend we find 8 prices that show no prevalent trend for the entire sample or a significant proportion (that is, 70%) of the time span. Using the HLTa method, there is more evidence (15 prices) of no significant trend. These results suggest that the Lewis (1954) model may be playing a part in the explanation of commodity price movement over time.

The prices of metals (copper, tin, lead, silver and zinc) show no evidence of a positive or negative trend. The rate of growth of consumption of zinc and lead has not been as high as that of aluminum, and technological advances and grade declines, which have both been modest, have almost exactly offset each other in determining lead and zinc production costs (Slade, 1982). This may explain the result of obtaining no trend. The only exception is aluminum which shows a negative trend for approximately 39% (that is, 1900 – 1941) of the sample. Growth rates for aluminum consumption have been high as new uses have been found, and technological advances, combined with economies of scale (Slade, 1982), may have lowered prices over a fraction of the period considered. In the 1970s, there was a serious concern raised by the Club of Rome (Meadows et al., 1972) that exhaustible resources would be depleted within thirty years. However, as documented by Radetski (2008), there has been a multi-fold increase in the production of copper and aluminum over the sample period considered in this study. This level of exploration and extraction of metals is likely to take place because as society exhausts existing mineral deposits, it is forced to exploit lower grade, more remote, and more difficult to process mineral resources. This activity tends to impart the upward trend of mineral commodity prices over time. However, at the same time, the costs and prices of non-renewables can be lowered with new discoveries and new technologies in exploration, and mining. If the cost-increasing effects of depletion are greater (lower) than the cost-reducing effects of new discoveries and technology, the real prices of mineral commodities tend to rise (decline) over time. Increases in the long run price of metals can reflect economic depletion, but the results from our study show no significant positive trends and so our evidence does not support the view that economic depletion has occurred.

The results in Tables 6a and 6b throw light on whether the primary commodities considered in this study are characterized as a difference stationary or a trend stationary process. The novelty of this method is that it allows for possible breaks if they exist, determined according to the Kejriwal and Perron (2010) and Harvey et al. (2010) sequential testing procedures. Out of the 24 commodity prices considered in this study, 8 commodity prices can be classified as difference stationary with or without breaks. The underlying price movements, whether they be trend stationary or difference stationary can seriously affect the income and consumption levels of developing countries. External shocks to commodity prices have important implications for the many developing countries that are dependent on commodity exports, as the persistence of upswings and downswings in prices can induce wide fluctuations in earnings from commodity exports (Cashin et al., 2002). Stabilization policies were introduced to smooth income flows. If the effect of any exogenous shock is short-lived, then the stabilization policies can be implemented to dampen the effect of such shocks and allow for external borrowing to smooth the path of national income and consumption. If commodity prices are shown to be difference stationary, then the cost of operating a price stabilization program would exceed the benefits of consumption or income smoothing. While it has been argued that stabilization policies are effective when the price series is trend stationary, they would be difficult to implement if the price series have a varying trend (Reinhart and Wickham, 1994). Our evidence indicates that 16 commodities display trend stationary behavior out of which 8 commodities (rice, wheat, sugar, beef, lamb, timber, lead and zinc) have no breaks in the trend. For these commodities, price stabilization policies are likely to be effective. However, as Cashin et al. (2000) and Cashin et al. (2002) have documented, the persistence of shocks to trend stationary commodity prices can vary to a large extent. For the remaining 8 commodities (coffee, tea, maize, palm oil, cotton, tobacco, jute and rubber) that do display trend stationary behavior but with a varying trend, such policies may be difficult to implement. In the case of the other 8 commodities (cocoa, banana, wool, hides, tin, copper, aluminum and silver) which exhibit difference stationary behavior, stabilization policies can prove to be ineffective. For instance, in the case of hides, we find a continuing relative decline coupled with an infinite persistence in prices. This may encourage exporting countries to take action to manage their supplies in order to keep the prices buoyant. Our results show that for a commodity such as hides, where there is no evidence of reversion to the long run trend, such policies should not be adopted. The reason is that the permanent effect of shocks to the price of hides renders such policies ineffective. In fact, price stabilization policies have been abandoned for many commodities which include cocoa where buffer stock operations ended in 1988; coffee, where regulated exports were abandoned in 1989; jute, where price stabilization ended with the 1989 agreement; and tin, where the International Tin Agreement collapsed in 1985 due to depletion of buffer stock. A commodity cartel for copper known as the Intergovernmental Council of Copper Exporting Countries, or CIPEC, was formed in 1967 by Chile, Peru, Zaire and Zambia with the purpose of raising prices through collective intervention in the copper market. A failed attempt to raise prices, by cutting back on production was made in the mid-1970s which was largely due to mistrust among members. Ultimately CIPEC was dissolved in 1988 due to withdrawal of several members. This study shows that for these commodities, we find evidence of difference stationary behavior or trend stationary behavior with varying trends. In the case of sugar, we find evidence of a negative stable trend with shocks that are transitory. The lapse of the International Sugar Agreement in 1984 could be the result of the attempt to stabilize prices at a high level without the flexibility to adjust downwards when the current level of stabilization turned out to be untenable. In recent years, commodity agreements have been transformed and their aims and objectives have been changed; the agreements are not concerned with price stabilization but are focused on promoting sustainability (Gilbert, 1996).

The variability in primary commodity prices may lead to variable export revenues, and also variable producer surpluses, variable consumption and variability in government fiscal position (Cashin et al., 2000). International compensatory finance schemes were established in the 1960s and 1970s with an aim to compensate for shortfalls in the export revenue of individual commodity exporting countries. If a country were to experience a shortfall in export revenues, then contributions were made from the schemes, and when export revenues recovered the country was expected to make the repayment. The Compensatory Finance Fund (CFF) was established by the IMF in 1963 and the STABEX in the mid-1970s by the European Union with these objectives in mind. However, these schemes did not last as these measures required the determination of an equilibrium price around which stabilization could be centered. It has been argued that the effectiveness of policies such as the compensatory financing scheme, should be based on the underlying nature of the trend and persistence of shocks to commodity prices. However, this scheme would not be effective for trends that are varying and shocks that are long lasting; rather in these cases the structural adjustment of the economy should be brought up to its new long run level of national income and consumption (Kaibni, 1986). The mixed evidence obtained on price trends and persistence for different commodities suggests that certain policy measures may not be effective. The heterogeneity of the results obtained for individual prices confirms the evidence obtained by Leon and Soto (1997), KW and Ghoshray that the use of aggregate measures may be misleading. For commodities that experience one or two breaks, forecasting of prices can prove to be difficult since the break points would be unpredictable.

Though it is generally believed that commodities in finite supply can be set apart from those commodities that are renewable, there are other factors that have been discussed earlier, such as exploration and capital investment that play an important role in the economics of resource depletion. These factors affect the underlying price trends making it difficult to form any general predictions about these trends. Real prices of non-renewable commodities have been roughly trendless over time with no general evidence of stationarity around deterministic trends or structural breaks, whereas the extraction of non-renewable resources has strongly increased (see Krautkraemer, 1998; Livernois, 2009; Cynthia-Lin and Wagner, 2007). If historical trends continue, innovation in the extraction technology will offset the depletion of easily accessible deposits. Even if non-renewable resource use and production increase exponentially, resource prices might stay constant in the long run. Extraction costs increase with cumulative extraction, but then remain constant as a “backstop” supply is reached.

7 Conclusion

This paper employs a range of novel econometric procedures to determine breaks in commodity prices, measure the underlying trends within the regimes delineated by the estimated break points and determine whether real primary commodity prices contain stochastic trends. An important methodological aspect of our analysis is that our evaluation of the direction and magnitude of trends is carried out without taking an a priori stand on the persistence of the noise component or on whether the breaks occur purely in level or in both level and slope. This is relevant from a practical standpoint since such persistence is usually known in practice and unit root pretesting has been shown to suffer from serious econometric problems. Moreover, in contrast to existing studies, we are able to distinguish between the case of pure level shifts and that of slope shifts accompanied by possible shifts in level. Further, we employ a new class of unit root tests in order to provide reliable evidence regarding the persistence of commodity price shocks. This class of tests allows for structural breaks under both the null and alternative hypotheses thereby alleviating these tests of size and power distortions that plague most existing procedures which only allow for breaks under the alternative of (broken) trend stationarity as well as ignore information regarding the presence or absence of breaks. We find that 8 out of 24 commodity prices can be characterized as difference stationary implying that shocks to these commodities tend to be permanent in nature. The remaining 16 prices are found to exhibit trend stationary behavior. For both types of trending behavior we find evidence of one or two structural breaks. The changes in economic conditions and environment over the length of time chosen for this study justify the case to allow for structural breaks. With the different commodities analyzed in this study, we observe different patterns of trends. Given that we find evidence that some commodities experience segments of a downward trend interspersed by periods of approximate stability, forecasting of commodity prices is likely to be difficult. Dependence on commodities with uncertain price trends and persistence can seriously destabilize the economy and as a result an appropriate policy response would be income smoothing. But these stabilization policies incur a cost; and given that average prices and trends are extremely hard to determine, actions which initially may seem to be purposeful may turn out to have diametrically opposite effects. The evidence from this study suggests that policy recommendations would be difficult to implement given the mixed and varying trend results.

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