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# **Breaking Trends and the Prebisch-Singer Hypothesis: A Further Investigation**

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# Breaking Trends and the Prebisch-Singer Hypothesis: A Further Investigation\*

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## Abstract

This paper examines the Prebisch-Singer Hypothesis employing new time series procedures that are robust to the nature of persistence in the commodity price shocks, thereby obviating the need for unit root pretesting. Specifically, the procedures allow consistent estimation of the number of structural breaks in the trend function as well as facilitate the distinction between trend breaks and pure level shifts. In comparison with past studies, we find fewer cases of commodities that display negative trends thereby weakening the case for the Prebisch-Singer Hypothesis. Finally, a new set of powerful unit root tests allowing for structural breaks under both the null and alternative hypotheses is applied to determine whether the underlying commodity price series can be characterized as difference or trend stationary processes. Relative to the extant literature, we find more evidence in favor of trend stationarity suggesting that real commodity price shocks are mostly of a transitory nature.

**Keywords:** primary commodity prices, structural breaks, trend functions, Prebisch-Singer Hypothesis, unit roots

**JEL Classification:** 013, C22

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## 1 Introduction

An issue which has been of great interest in the trade and development literature is the possible existence of a long run trend in primary commodity prices relative to the price of manufactures. 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 following the independent studies by Prebisch (1950) and Singer (1950) which claimed that commodity prices should decline in relation to manufactured goods in the long run, which was labeled as the well known Prebisch-Singer Hypothesis.

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 (1999) argues that 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 states 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. As a result 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 base.

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 leading to a large volume of studies in the trade and development economics literature. The evidence has been mixed which leads to serious policy implications as to whether developing countries should specialize in primary commodity exports. The World Bank, for instance, has encouraged long term primary commodity projects in developing countries (Ardeni and Wright, 1992). Besides, free market solutions were provided to developing countries to deal with primary commodities instead of positive intervention (Maizels, 1994). The upshot is that countries which rely on the exports of primary commodities must understand the nature of commodity prices in order to devise their development macroeconomic policy (Deaton, 1999).

The continuing interest in the Prebisch-Singer Hypothesis stems from the fact that the central question is an empirical one. 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. If the price series is found to contain a unit root then the series is said to contain a stochastic trend such that the effect of shocks to the underlying series will be permanent. In other words, the noise component will be integrated of order one, that is,  $I(1)$ . If, however, the underlying price series is found to reject a unit root then the series is considered to be trend stationary and the effect of shocks on the series would be transitory in nature, that is, the noise will be integrated of order zero, that is,  $I(0)$ . 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 price series contains 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 price series is 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. These testing procedures are typically based on the minimum  $t$ -statistic corresponding to the unit root parameter over the set of permissible break dates or alternatively 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).

A problem with the application of these unit root tests is that while they allow for the possibility of structural breaks under the alternative hypothesis of (broken) trend stationarity, they do not allow for breaks under the null hypothesis of a unit root. Consequently, when breaks exist under the null, these tests tend to over-reject. The extent of size distortions associated with these tests has been amply demonstrated, for example, in Nunes et al. (1997), Lee and Strazicich (2001, 2003) and Kim and Perron (2009). Kim and Perron (2009) show that these tests can also suffer from low power as they do not exploit information about the presence of breaks. Given these size and power distortions, evaluating the significance of the trend coefficient based on these tests is likely to provide misleading conclusions regarding the empirical relevance of the Prebisch-Singer Hypothesis. Moreover, these tests 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. Finally, 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 Prebisch-Singer Hypothesis based on an updated Grilli-Yang index over 1900-2008 employing econometric procedures that allow consistent estimation of the number of structural breaks in the trend function while being completely agnostic about the nature of persistence in the commodity price shocks, thereby obviating the need for unit root pretesting. At a fundamental level, it seems intuitive to first determine if structural breaks exist at all before conducting unit root tests that allow for such breaks. The fact that earlier methods do not test for the number of structural breaks suggests that these methods may have low power in detecting deviations from the unit root if the researcher allows for more breaks than are actually present in the data generating process. Our analysis also facilitates a clear demarcation between slope breaks (accompanied by possible level breaks) and breaks that take the form of pure level shifts. Our procedures 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). In contrast to the existing literature, we find fewer cases of commodities that display negative trends thereby weakening the case for the Prebisch-Singer Hypothesis. Finally, 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. Relative to past studies, we find more evidence in favor of trend stationarity thereby suggesting that commodity price shocks are mostly of a transitory nature.

The rest of the paper is organized as follows. Section 2 briefly reviews the Prebisch-Singer Hypothesis and motivates the use of econometric techniques employed in this paper. Section 3 provides a description the econometric methodology. Section 4 presents the empirical results along with a discussion of the updated Grilli Yang Index. Section 5 discusses the policy implications of our analysis, and finally Section 6 concludes.

## 2 Literature Review

The original work on the Prebisch-Singer Hypothesis assumed that the underlying data series is trend stationary (Prebisch, 1950 and Singer, 1950). While Sapsford (1985), Grilli and Yang (1988), Helg (1991) and Ardeni and Wright (1992) among others have advocated a trend stationary model for commodity prices, Cuddington and Urzua (1989), Cuddington (1992), Bleaney and Greenaway (1993) and Newbold et al. (2005) recognized that commodity prices may be difference stationary. The evidence on the Prebisch-Singer hypothesis has been mixed. While Sapsford (1985) and Helg (1991) tend to support the hypothesis, Newbold and Vougas (1996) cannot provide compelling evidence as to whether the data is trend stationary or difference stationary, whereas Kim et al. (2003) suggest that commodity prices generally display unit root behavior and that there is limited evidence for the Prebisch-Singer Hypothesis.

The issue of structural breaks applied to primary commodity prices has been of great interest to researchers. Powell (1991), using a cointegration approach, narrowly rejects the Prebisch-Singer Hypothesis in favor of downward shifts in the real commodity price series for the years 1921, 1938 and 1975 which coincide with the sharp downturn in manufacturing output for the industrialized countries. Leon and Soto (1997) applied the single break Zivot–Andrews (Zivot and Andrews, 1992) test on primary commodity prices and found evidence of structural change. Their results seem to suggest that most commodity

prices can be described as trend stationary models and that the trend coefficients generally support the Prebisch-Singer Hypothesis.

Zanias (2005) and Kellard and Wohar (2006) have employed the Lumsdaine-Papell (Lumsdaine and Papell, 1997) test to allow for two structural breaks. Zanias (2005) applies the test to the extended aggregate Grilli Yang Index and concludes that the data can be adequately described by a trend stationary process with two intercept shifts. The breaks are identified in 1920 and 1984 which cumulatively account for a 62% drop. However, Zanias (2005) considered the aggregate index, which has been questioned by recent studies, which argue that there is considerable heterogeneity in the behavior of individual commodities. The study by Kellard and Wohar (2006, KW hereafter) is a case in point. KW conduct a study using the Grilli Yang Index of disaggregated commodity prices over the period 1900–98. Out of 24 commodity prices, their results indicate 14 are trend stationary allowing for 1 or 2 structural breaks. Overall, they show that the deterioration of commodity prices has been discontinuous. Following Leon and Soto (1997), KW argue that the existence of a single linear trend would be ‘strong evidence’ in favor of the Prebisch-Singer hypothesis. KW state that a single trend is a ‘summary measure’ of several trends which may be positive or negative. Arguing that reliance on a single trend may be misleading to policy makers, KW develop a measure to define the prevalence of a trend.

Balagtas and Holt (2009) consider tests of the linear unit root model against smooth transition and time varying alternatives. They note that while previous research focused on breaks in the intercept and trend, a more complete analysis would allow for breaks in the autoregressive and moving average terms. The main contribution of their paper is testing for and estimating nonlinear alternatives to a secular deterioration. Bootstrap procedures are employed to test the linear unit root model against different categories of smooth transition autoregression (STAR) models for the 24 disaggregated commodities of the Grilli Yang Index, from 1900-2003. Using the conventional significance levels, they show that in 19 of the 24 cases they can reject the null hypothesis of a linear model and are able to fit STAR type model to 16 of those 19 nonlinear models. Simulation experiments are conducted to confirm that there is little support for the Prebisch-Singer Hypothesis. However, though Balagtas and Holt (2009) argue that the nonlinearity in commodity price adjustment arises due to the impossibility of negative storage (Deaton and Laroque, 1995), they do not attempt to provide any economic intuition for such breaks. Further, no explanation is provided as to why such breaks should be smooth. For instance, Ocampo and Parra (2003) have argued that the decline in terms of trade have been discontinuous.

A common theme in most of these studies is that they typically employ unit root



tests which allow for breaks under the alternative of trend stationarity but not under the null hypothesis of a unit root. These tests have been shown to suffer from serious size distortions due to the asymmetric treatment of breaks under the null and alternative hypotheses. Moreover, if a break is indeed present, this information is not exploited to improve the power of the testing procedure (a detailed discussion of this issue together with Monte Carlo evidence demonstrating the finite sample problems associated with this type of tests can be found in Kim and Perron, 2009, Nunes et al., 1997 and Lee and Strazicich, 2001 and 2003). Further, in most cases, the estimates of the break dates are obtained by minimizing/maximizing these unit root tests over all possible break dates which, in general, do not provide consistent estimates of the true break dates (Vogelsang and Perron, 1998). In fact, these 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.

Ghoshray (in press, Ghoshray hereafter) addresses some of the drawbacks of these studies particularly in relation to the choice of the null hypothesis that only allows for a linear unit root. Ghoshray attempts to address this gap by allowing for breaks under both the null hypothesis of a unit root as well as the trend stationarity alternative by employing the LM unit root test developed by Lee and Strazicich (2003, 2004). The test 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. Besides, the LM test possesses greater power than the Lumsdaine-Papell test. The main findings of this paper are that 11 out of 24 commodity prices are found to be difference stationary implying that shocks to these commodities tend to be permanent in nature. The remaining thirteen prices are found to exhibit trend stationary behavior with either one or two structural breaks. Most of the commodities that do not exhibit difference stationary behavior seem to contain no significant trends. There are fewer cases, in relation to past studies, of commodities that display negative trends thereby weakening the case for the Prebisch-Singer hypothesis. 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 of 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. In other words, the empirical methodology estimates the trend function without any requirements for unit root pretesting. 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 Prebisch-Singer Hypothesis. However, evidence of the Prebisch-Singer Hypothesis 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. An important limitation of their procedure is that it only allows for a single break in slope and does not attempt to distinguish between breaks in slope (possibly accompanied by breaks in level) and pure level breaks.

### 3 Econometric Methodology

Our econometric methodology is aimed at addressing each of the limitations associated with existing procedures as discussed in the previous section. The most general model considered can be described as:

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

$$u_t = \alpha u_{t-1} + v_t, \quad t = 2, \dots, T, \quad u_1 = v_1 \quad (2)$$

where  $DU_{it} = I(t > T_i)$ ,  $DT_{it} = (t - T_i)I(t > T_i)$ ,  $i = 1, \dots, K$ . A break in the trend occurs at time  $T_i = [T\lambda_i]$  when  $\beta_i \neq 0$ . The dates 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  $H_1: \beta_i \neq 0$ .<sup>1</sup>

The first step tests for one structural break (that is  $K = 1$  in (1)) in the slope of the trend function using procedures that are robust to the stationarity/non-stationarity

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<sup>1</sup>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 HLT). This, however, does not mean that the tests are incorrectly sized in the presence of pure level shifts (see the simulation experiments in section 3).

properties of the data. We employ two such procedures proposed by PYb and HLTb respectively. 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.<sup>2</sup> 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.

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 (1)) 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 in our empirical analysis (109), we allow for a maximum of two breaks in our empirical analysis.<sup>3</sup> 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 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.<sup>4</sup>

Conditional on the presence of a stable slope at the initial step (that is  $\beta_i = 0$  in (1) for  $i = 1, \dots, K$ ), 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

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<sup>2</sup>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.

<sup>3</sup>This assumption is common to the majority of existing empirical studies.

<sup>4</sup>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).

root/stationarity properties of the data.<sup>5</sup> 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 (1). 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 sample. In models with pure level shifts, trend significance is 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 (1) for at least one  $i = 1, \dots, K$ ), 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).<sup>6</sup> 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).

To ensure brevity of the main text as well as to enhance readability, we have relegated the discussion of the various testing procedures including the notation for the different tests and estimates to the Appendix.

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<sup>5</sup>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.

<sup>6</sup>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.

## 4 Data and Empirical Results

Prior to Grilli and Yang’s (1988, GY hereafter) seminal paper, there were many inadequacies in the availability of consistent commodity price data in the economics profession. The *Economist* Index and the W. A. Lewis Index were then two of the main indices which provided a large backlog of data on commodity prices, even though their level of data accuracy had often been called into question. The former had been subject to frequent revisions and was weighted by “the relative values of commodities in the import trade of industrial countries” (GY, p.3), resulting in a one-sided focus on the import patterns of developed countries, and it also excluded fuel-based commodities. On the other hand, the W. A. Lewis Index only ran until 1938 and used the export unit values of certain countries rather than international market quotations.

In terms of the available data on manufactured goods series, again the W. A. Lewis Index started in 1870 but had gaps in the data owing to both World Wars. Maizels (1970, cited by GY, p.3) also constructed a series but this only comprised average prices of certain periods. Finally, although the United Nations’ data on manufacturing unit values (MUV) spanned the entire twentieth century, again it contained gaps in the data during the war years.

Given this situation regarding data inadequacy, GY opted to construct “a U.S. dollar index of prices of twenty-four internationally-traded nonfuel commodities, beginning in 1900” (GY, p.3). The Grilli Yang Commodity Price Index (GYCPI) is “base-weighted, with 1977-79 values of world exports of each commodity used as weights” (GY, p.3) and is a means of capturing the evolution of international prices of a basket of primaries.

As for the updated version of the MUV, the gaps in the data are corrected for, via interpolations using U.S. and U.K. data on export and import unit values. This updated series “reflects the unit values of exports of manufactures of a number of industrial countries” (GY, p.5), with weights which vary to show the changing importance of different manufactured goods over time – such changes are shown in several updates, which occur every few years until 1938 and then again in: 1959, 1963, 1970, 1975, and 1980 (United Nations 1969, 1972, 1976, 1982, and 1987, cited in GY, p.5). GY then go on to use the series GYCPI/MUV which “measures the evolution of the purchasing power of a basket on nonfuel primary commodities in terms of traded manufactures, valued at international prices” (GY, p.7).

One caveat to this construction process is that neither the manufactured goods price series nor the primary commodity price series can be complete proxies for the components of the net barter terms of trade, as developing countries’ level of total imports comprise

more than manufactured goods, whilst developed countries' total imports comprise more than just primary commodities. Furthermore, GY place emphasis on the fact that a declining trend in relative prices should not be taken solely as a declining real income effect - the income effect relies not only on movements in relative prices but also on the evolution of the purchasing power of exports; what's more, the authors highlight that "one has to account simultaneously for the movements in the relative prices of exports and for the quantity of exports" (GY, p.7) - an expression which reflects this is the income terms of trade which "reflects the purchasing power of total exports in terms of imports."

An extended data set of the original GYCPI is employed in this study. The data was updated according to the method documented in the paper 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 MUV index. Figure 1 plots the 24 deflated commodity price series.

[Figure 1 about here]

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  $t_\eta$  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.

[Table 1 about here]

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. A further 2 commodity prices were found to contain two breaks in the intercept. No structural breaks were found for the remaining 11 commodity prices.

Having determined that structural breaks are present in 13 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.

[Tables 2 and 3 about here]

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.

As described earlier, 2 commodities were found to exhibit pure level shifts. The estimated trend coefficients for these 2 commodities are calculated and reported in Table 4.

[Table 4 about here]

For both commodities (hides, tin) we find that the trend coefficients are positive and significant across the different regimes delineated by the structural breaks. The estimated break dates in this case may be interpreted as prices experiencing a sharp jump or collapse at the break points. Powell (1991) provides an explanation as to why relative commodity prices that experience a positive trend over a certain interval of time may lead to a negative shift in commodity prices. During periods of high commodity prices innovation in production methods or the use of substitute commodities are promoted. When the boom period is over, the correction is greater than expected contributing to a sharp drop

in commodity prices. This explanation may be offered for the first structural break which occurred after a period when commodity prices had risen substantially after the First World War boom. Institutional factors can be used to interpret pure shifts, such as the collapse of the International Tin Agreement coincides with the second structural break for tin.

Figure 2 below plots the 13 commodity prices that experience structural breaks. 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 seem to be more plausible.

[Figure 2 about here]

For the 11 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 5 below.

[Table 5 about here]

Except for rice and wheat, both the HLTa and PYa methods produce estimates of the trend function of the same sign. Out of the 11 commodities, 5 commodities do not show any evidence of a significant positive or negative trend. Only sugar shows support of the Prebisch-Singer hypothesis, whereas for beef, lamb and timber, the sign of the trend function is positive. While there is no evidence of a significant trend in rice and wheat using the HLTa method, we find a significant negative trend according to the PYa method.

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  $\Psi(.)$  [defined as  $\Psi(.) = 1 - \Psi(-) - \Psi(+)$ ]. Table 6 displays the relative measure results for all 24 commodities.

[Table 6 about here]

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. If one were to consider the PYa method, then 13 of the 24 commodities contain at least a single significant



negative trend. According to the HLTa method, only 1 commodity (sugar) shows a significant negative trend for the entire sample, whereas for the PYa method we find two further commodities, being rice and wheat.

Using the HLTa approach, no other commodity shows a negative trend for 70% of the sample period. If we were to consider at least 50% of the sample period, then the number of commodities rises by two (cotton, jute). Contrasting the result with the PYa method, for 70% of the sample period a negative trend is prevalent in 4 commodities, and the number rises to 8 commodities in total if we were to consider 50% of the sample.

Comparing the results with KW and Ghoshray, we find that there is less evidence of a prevalent negative trend. While KW find 8 commodities out of 24 to contain a negative trend (over 70% of the sample period), and Ghoshray find 6 out of 24 commodities, our results show 1 commodity using the HLTa and 4 commodities using the PYa method. The upshot is that our results further weaken the case for the Prebisch-Singer Hypothesis.

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 Table 7.

[Table 7 about here]

The results from Table 7 show that 17 of the 24 commodity prices can be classified as a trend stationary process. For the remaining 7 prices (cocoa, banana, wool, tobacco, copper, aluminum and silver) the null hypothesis of a unit root was not rejected, concluding that the prices contained a unit root. Comparing the results with Ghoshray, where 11 commodities are found to be difference stationary, and KW where 10 prices are difference stationary, we find that there are fewer commodities that are classified as difference stationary. Comparing with Ghoshray a similar match is found for only 2 commodities (being cocoa and aluminum) and with KW only 4 commodities (being cocoa, banana, tobacco and copper). However, one must note that the sample size chosen in this study is slightly longer than Ghoshray and more so in the case of KW. Our results show that for the 17 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

7 commodity prices found to be difference stationary, one can expect that the effects of shocks to these commodities are likely to be permanent.

## 5 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 apriori 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 Prebisch-Singer Hypothesis.

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 6 summarizes the prevalence of trends. The prevalence of a negative trend is found to be present in fewer commodities in comparison to recent studies by KW, Ghoshray and Harvey et al. (2010). Our study finds a negative trend to exist only for rice, wheat and sugar over the entire time span and a prevalent negative trend (for more than 70% of the time span) is found for banana. As a result, the case for the Prebisch-Singer hypothesis is considerably weakened by our results. Apart from wheat, a negative trend is mainly found for commodities exported by developing countries. Urgent solutions are needed when a country is highly dependent on one or a few commodity exports. For example, St. Vincent and Honduras derive 20-49% of their earnings from banana. Mauritius and Guyana earn between 20-49% from their exports of sugar. It has been argued by institutions such as the World Bank that countries which experience a deteriorating terms of trade should diversify into new exports of primary commodities and away from primary commodities that are in oversupply. Some Asian countries, such as Hong Kong, Singapore, South Korea and Taiwan, have benefited from successful export

diversification policies, whereas other developing countries, especially in Africa have been left behind. Diversification of exports into other primary commodities would depend on the existing resource availability and potential export destinations.

Interestingly, 5 commodity prices (hides, tin, timber, beef and lamb) show a positive trend over the entire sample. This result contrasts sharply with that of KW and Ghoshray. According to PYa estimate of the trend we find 7 prices that show no significant trend for the entire sample or a significant proportion of the time span. Using the HLTa method, there is more evidence (13 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.

Lutz and Singer (1994) indicate that policies by Bretton-Woods institutions have, intentionally or not, promoted the production and expansion of primary commodities by developing countries, contributing to the declining trend of commodities. A natural question that arises is whether one can make a case for promoting industrialization in such countries. Some studies have hinted that the trends in prices of manufactures from developing countries may be on the decline (see Sarkar and Singer, 1991; Maizel et al., 1998). The exponents of inward looking development strategies tend to make a case for import substitution. However, such protectionist policies have been criticized, particularly on grounds of misallocation of resources, inefficiency and corruption (Ocampo and Parra, 2007). The evidence suggests that such policy measures can be detrimental given the mixed results obtained on price behavior for different commodities. For example, when considering major exports of developing countries, we find that sugar and rice display a trend stationary process with a negative trend for the entire period. This is in sharp contrast to cocoa, copper, silver and banana, which exhibit a driftless random walk.

The heterogeneity of the results for the estimated trends obtained for individual prices confirms the evidence obtained by Leon and Soto (1997), KW and Ghoshray that the use of aggregate measures [see for example, Zanias (2005)] may be misleading. Besides, import substitution has become unpopular with countries such as Brazil and India. Both countries had initially embraced this policy but subsequently rejected them in favor of liberalized market policies with particular reference to exports (Sapsford and Balasubramanyam, 1994).

The results in Table 7 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, 7 commodity prices

(cocoa, banana, wool, tobacco, copper, aluminum and silver) can be classified as difference stationary with or without breaks. The other 17 commodities can be classified as trend stationary with or without breaks. This result contrasts sharply with recent studies made by KW, where they find 10 commodity prices to be classified as trend stationary and Ghoshray's study which finds 13 commodity prices to be trend stationary.

The underlying price movements, whether they be trend stationary or difference stationary can seriously affect the income and consumption levels of developing countries. Stabilization policies were introduced to smooth income flows. While it has been argued that stabilization policies are effective when the price series is trend stationary, they may be difficult to implement if the price series have a varying trend (Reinhart and Wickham, 1994).

Our evidence indicates that 17 commodities display trend stationary process out of which 10 commodities (rice, wheat, sugar, beef, lamb, hides, timber, tin, lead and zinc) have no breaks in the trend. For these commodities price stabilization policies are likely to be effective. However, for the remaining 7 (coffee, tea, maize, palm oil, cotton, jute and rubber) commodities that do display trend stationary behavior with a varying trend, such policies may be difficult to implement. In the case of a further 7 commodities (cocoa, banana, wool, tobacco, copper, aluminum and silver) which exhibit difference stationary process, stabilization policies can prove to be 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; sugar, where price stabilization measures were removed in 1992; 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. This study shows that for these commodities, (except sugar and tin) we find evidence of difference stationary behavior or trend stationary process with varying trends. Nowadays, these commodity agreements are not concerned with price stabilization but are focused on promoting sustainability.

The compensatory financing scheme is a mechanism designed to smooth the effects of shortfalls in export revenues. This setback may arise due to a negative trend in primary commodity prices over a certain interval of time. An example of such a financing scheme is the IMF's CFF which commenced in 1963 and replaced by the CCFF in 1988. The EU maintained three schemes (STABEX, SYSMIN and COMPEX). The CCFF provides countries that lack reserves or the capacity to borrow, with the necessary finance for consumption smoothing. However this scheme is not effective for 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). On the other hand, the SYSMIN appeared to benefit the industrialized countries, COMPEX was ineffective and largely symbolic, while STABEX was abolished as it was deemed to be inequitable and counter-productive (Page and Hewitt, 2001). While most of these compensation schemes are defunct, one could consider the effectiveness of replacements such as the CRMG of the World Bank. These policies should be based on the underlying prevalence of the trend, and evidence from this study suggests that the design and form of assistance would be difficult to implement given the mixed and varying trend results that are found for various commodities.

It is difficult to draw conclusions regarding future terms of trade for most commodities where we find evidence that segments of a downward trend are interspersed by periods of approximate stability. However, when comparing with recent studies such as KW and Ghoshray, there are relatively more commodities that experience a stable linear trend over the entire sample. For commodities that experience one or two breaks, forecasting of prices may be difficult since the break points can be unpredictable.

## 6 Conclusion

This paper employs a range of novel tests to determine breaks in commodity prices, measure the underlying trends within the regimes defined by the 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 Prebisch-Singer Hypothesis is carried out without taking an apriori 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. Our findings indicate that there are fewer cases, in relation to past studies, of commodities that display negative trends thereby weakening the case for the Prebisch-Singer hypothesis.

We also 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 procedures which only allow for breaks under the alternative of (broken) trend stationarity. We find that 7 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 17 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 proves to be difficult. 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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## Appendix: Description of Testing Procedures

### A.1 Robust Tests for Breaks in Trend

#### A.1.1 The Harvey et al. (2009) Test for a Break in Slope

Harvey et al. (2009) propose test statistics that are constructed by taking a weighted average of the regression  $t$ -statistics from a regression in levels and a regression in differences. The weighting function is based on the KPSS stationarity statistics applied to the levels and differenced data. First differencing (1) [for  $K = 1$ ] yields

$$\Delta y_t = \beta_0 + \mu_1 D_{1t} + \beta_1 DU_{1t} + \varepsilon_t, \quad t = 2, \dots, T \quad (\text{A.1})$$

where  $\varepsilon_t = \Delta u_t$ ,  $D_{1t} = I(t = T_1 + 1)$  and  $DU_{1t} = I(t > T_1)$ . Consider the  $t$ -statistics

$$t_0(\lambda_1) = \frac{\hat{\beta}_1(\lambda_1)}{\sqrt{\hat{\omega}_1^2(\lambda_1) [\{\sum_{t=1} x_{L1,t}(\lambda_1) x_{L1,t}(\lambda_1)'\}^{-1}]_{44}}} \quad (\text{A.2})$$

$$t_1(\lambda_1) = \frac{\tilde{\beta}_1(\lambda_1)}{\sqrt{\tilde{\omega}_1^2(\lambda_1) [\{\sum_{t=1} x_{D1,t}(\lambda_1) x_{D1,t}(\lambda_1)'\}^{-1}]_{33}}} \quad (\text{A.3})$$

In (A.2),  $x_{L1,t}(\lambda_1) = \{1, t, DU_{1t}, DT_{1t}\}$ ,  $DT_{1t} = (t - T_1)I(t > T_1)$ ,  $\hat{\beta}_1(\lambda_1)$  is the OLS estimate of  $\beta_1$  from (1) and  $\hat{\omega}_1^2(\lambda_1)$  is an estimate of the long-run variance based on the OLS residuals  $\hat{u}_t(\lambda_1) = y_t - \hat{\mu}_0(\lambda_1) - \hat{\beta}_0(\lambda_1)t - \hat{\mu}_1(\lambda_1)DU_{1t} - \hat{\beta}_1(\lambda_1)DT_{1t}$ . In (A.3),  $x_{D1,t}(\lambda_1) = \{1, D_{1t}, DU_{1t}\}$  and  $\tilde{\beta}_1(\lambda_1)$  is the OLS estimate of  $\beta_1$  from (A.1) and  $\tilde{\omega}_1^2(\lambda_1)$  is an estimate of the long-run variance based on the residuals  $\tilde{\varepsilon}_t(\lambda_1) = y_t - \tilde{\beta}_0(\lambda_1) - \tilde{\mu}_1(\lambda_1)D_{1t} - \tilde{\beta}_1(\lambda_1)DT_{1t}$ . The following long-run variance estimators are used:

$$\begin{aligned} \hat{\omega}_1^2(\lambda_1) &= T^{-1} \sum_{t=1}^T \hat{u}_t^2(\lambda_1) + 2T^{-1} \sum_{j=1}^{T-1} (1 - j/(l+1)) \sum_{t=j+1}^T \hat{u}_t(\lambda_1) \hat{u}_{t-j}(\lambda_1) \\ \tilde{\omega}_1^2(\lambda_1) &= (T-1)^{-1} \sum_{t=1}^T \tilde{\varepsilon}_t^2(\lambda_1) + 2(T-1)^{-1} \sum_{j=1}^{T-2} (1 - j/(l+1)) \sum_{t=j+2}^T \tilde{\varepsilon}_t(\lambda_1) \tilde{\varepsilon}_{t-j}(\lambda_1) \end{aligned}$$

with  $l = [4(T/100)^{1/4}]$ . Next, consider stationarity statistics  $S_0(\lambda_1)$  and  $S_1(\lambda_1)$  calculated from the residuals  $\{\hat{u}_t(\lambda_1)\}_{t=1}^T$  and  $\{\tilde{\varepsilon}_t(\lambda_1)\}_{t=2}^T$  respectively:

$$\begin{aligned} S_0(\lambda_1) &= \frac{\sum_{t=1}^T \left( \sum_{i=1}^t \hat{u}_i(\lambda_1) \right)^2}{T^2 \hat{\omega}_1^2(\lambda_1)} \\ S_1(\lambda_1) &= \frac{\sum_{t=2}^T \left( \sum_{i=2}^t \tilde{\varepsilon}_i(\lambda_1) \right)^2}{(T-1)^2 \tilde{\omega}_1^2(\lambda_1)} \end{aligned}$$

The next step is to choose a weight function which converges to unity when  $u_t$  is  $I(0)$  and to zero when  $u_t$  is  $I(1)$ . Based on the properties of the stationarity statistics, the weight function  $\eta(S_0(\lambda_1), S_1(\lambda_1)) = \exp[-\{g_1 S_0(\lambda_1) S_1(\lambda_1)\}^{g_2}]$  is recommended. Finally, the proposed test statistic is

$$t_\eta = \left\{ \eta(S_0(\hat{\lambda}_1), S_1(\hat{\lambda}_1)) \right\} t_0(\hat{\lambda}_1) + m_\xi \left\{ \left[ 1 - \eta(S_0(\hat{\lambda}_1), S_1(\hat{\lambda}_1)) \right] \right\} t_1(\tilde{\lambda}_1) \quad (\text{A.4})$$

where  $\hat{\lambda}_1 = \arg \sup_{s \in \Lambda_1} |t_0(\frac{s}{T})|$ ,  $\tilde{\lambda}_1 = \arg \sup_{s \in \Lambda_1} |t_1(\frac{s}{T})|$  with  $\Lambda_1 = [\epsilon T, (1 - \epsilon)T]$ . The parameter  $\epsilon$  determines the level of trimming used. The positive constant  $m_\xi$  is chosen such that, for a significance level  $\xi$  under  $H_0$ , the asymptotic critical value in the  $I(0)$  and  $I(1)$  cases coincide. This ensures that the asymptotic null critical values of  $t_\eta$  are the same regardless of whether  $u_t$  is  $I(0)$  or  $I(1)$ . Based on a range of Monte Carlo simulations on the finite sample size and power of the tests, they recommend choosing  $g_1 = 500$  and  $g_2 = 2$  for the construction of the weight function  $\eta(\cdot)$ . Note that both stationarity statistics are evaluated at the breakpoint estimator  $\hat{\lambda}_1$ , this being a consistent estimator of the true break fraction irrespective of whether  $u_t$  is stationary or not.

### A.1.2 The Perron and Yabu (2009b) Test for a Break in Slope

Perron and Yabu (2009b) propose an alternative approach to testing the stability of the trend function based on a Feasible Quasi Generalized Least Squares procedure. First, the OLS estimate of  $\alpha$  is obtained from the autoregression

$$\hat{u}_t = \alpha \hat{u}_{t-1} + \sum_{i=1}^k \zeta_i \Delta \hat{u}_{t-i} + e_{tk} \quad (\text{A.5})$$

where  $k$  is chosen using the Bayesian Information Criterion (BIC) ( $k$  is allowed to be in the range  $[0, [12(T/100)^{1/4}]]$ ). The corresponding estimate is denoted  $\tilde{\alpha}$ . To improve the finite sample properties of the tests, Perron and Yabu use a bias-corrected version of  $\tilde{\alpha}$ , denoted  $\tilde{\alpha}_M$ , proposed by Roy and Fuller (2001) (See Perron and Yabu, 2009b for details of the bias correction procedure). Next, Perron and Yabu propose the use of the following super-efficient estimate of  $\alpha$ :

$$\tilde{\alpha}_{MS} = \begin{cases} \tilde{\alpha}_M & \text{if } |\tilde{\alpha}_M - 1| > T^{-1/2} \\ 1 & \text{if } |\tilde{\alpha}_M - 1| \leq T^{-1/2} \end{cases}$$

It is shown that using such a super-efficient estimate is crucial for obtaining procedures with nearly identical limit properties in the  $I(0)$  and  $I(1)$  cases. This estimate is then used to construct the quasi-differenced regression

$$\begin{aligned} (1 - \tilde{\alpha}_{MS}L)y_t &= (1 - \tilde{\alpha}_{MS}L)x'_{L1,t}\Psi + (1 - \tilde{\alpha}_{MS}L)u_t, \quad t = 2, \dots, T \\ y_1 &= x'_{L1,1}\Psi + u_1 \end{aligned} \quad (\text{A.6})$$

where  $\Psi = (\mu_0, \beta_0, \mu_1, \beta_1)'$ . Denote the resulting estimates by  $\tilde{\Psi}^{FG} = (\tilde{\mu}_0^{FG}, \tilde{\beta}_0^{FG}, \tilde{\mu}_1^{FG}, \tilde{\beta}_1^{FG})'$ . The Wald test  $W_{QF}(\lambda_1)$  for a particular break fraction  $\lambda_1$ , where the subscript  $QF$  stands for Quasi Feasible GLS, is given by

$$W_{QF}(\lambda_1) = \frac{\left(\tilde{\beta}_1^{FG}(\lambda_1)\right)^2}{\sqrt{\tilde{h}_v(\lambda_1) [(X^{\alpha'} X^{\alpha})^{-1}]_{44}}}$$

where  $X^{\alpha} = \{x_{L1,1}, (1 - \tilde{\alpha}_{MS}L)x_{L1,2}, \dots, (1 - \tilde{\alpha}_{MS}L)x_{L1,T}\}'$ . The quantity  $\tilde{h}_v(\lambda_1)$  is an estimate of  $(2\pi \text{ times})$  the spectral density function of  $v_t = (1 - \alpha L)u_t$  at frequency zero. When  $|\tilde{\alpha}_{MS}| < 1$ , a kernel-based estimator

$$\tilde{h}_v(\lambda_1) = T^{-1} \sum_{t=1}^T \hat{v}_t^2(\lambda_1) + 2T^{-1} \sum_{j=1}^{T-1} k(j, \tilde{l}) \sum_{t=j+1}^T \hat{v}_t(\lambda_1) \hat{v}_{t-j}(\lambda_1)$$

is used where  $\hat{v}_t(\lambda_1)$  are the OLS residuals from (A.6). The function  $k(j, \tilde{l})$  is the quadratic spectral kernel and the bandwidth  $\tilde{l}$  is selected according to the plug-in method advocated by Andrews (1991) using an AR(1) approximation. They also consider an alternative choice based on an autoregressive spectral density estimator (at frequency zero). Both estimators yielded very similar results in our context, hence we report results based on the kernel-based estimator only. When  $\tilde{\alpha}_{MS} = 1$ , the estimate suggested is an autoregressive spectral density estimate that can be obtained from the regression

$$\hat{v}_t = \sum_{i=1}^k \zeta_i \hat{v}_{t-i} + e_{tk} \quad (\text{A.7})$$

Denoting the estimate by  $\hat{\zeta}(L) = 1 - \hat{\zeta}_1 L - \dots - \hat{\zeta}_k L^k$  and  $\hat{\sigma}_{ek}^2 = (T - k)^{-1} \sum_{t=k+1}^T \hat{e}_{tk}^2$ ,  $\tilde{h}_v = \hat{\sigma}_{ek}^2 / \hat{\zeta}(1)^2$ . The order of the autoregression (A.7) is again selected using the BIC.

Following Andrews (1993) and Andrews and Ploberger (1994), Perron and Yabu considered the Mean, Exp and Sup functionals of the Wald test for different break dates. They found that with the Exp functional, the limit distributions in the  $I(0)$  and  $I(1)$  cases are nearly identical. They thus recommend the test statistic

$$ExpW = \log \left[ T^{-1} \sum_{\lambda_1 \in \Lambda_1} \exp \left( \frac{1}{2} W_{QF}(\lambda_1) \right) \right]$$

### A.1.3 The Harvey et al. (2010) Test for Breaks in Level

Harvey et al. (2010) propose a robust procedure for detecting multiple level breaks while accommodating a linear trend in the underlying data generating process. The model considered is

$$y_t = \mu_0 + \sum_{i=1}^n \mu_i I(t > T_i) + \beta_0 t + u_t$$

The null hypothesis is  $H_0: \mu_i = 0$  for  $i = 1, \dots, n$  while the alternative is that of at least one break in level; that is  $H_1: \mu_i \neq 0$  for at least one  $i \in \{1, \dots, n\}$ . Let  $\hat{\beta}_0$  denote the estimator of the trend coefficient,  $\beta_0$ , from the OLS regression of  $y_t$  on  $\{1, t\}, t = 1, \dots, T$ . The proposed test statistic is based on the quantities

$$\begin{aligned} M &= \max_{t \in \Lambda_1} \left| M_{t, [mT]} - \hat{\beta}_0 \left[ \frac{m}{2} T \right] \right| \\ S_0 &= (\hat{\omega}_v)^{-1} T^{-1/2} M \\ S_1 &= (\hat{\omega}_u)^{-1} T^{1/2} M \end{aligned}$$

where

$$M_{t, [mT]} = \frac{\sum_{i=1}^{\lfloor \frac{m}{2} T \rfloor} y_{t+i} - \sum_{i=1}^{\lfloor \frac{m}{2} T \rfloor} y_{t-i+1}}{\lfloor \frac{m}{2} T \rfloor}$$

and  $\hat{\omega}_v, \hat{\omega}_u$  denoting long-run variance estimates appropriate for the case of  $I(1)$  and  $I(0)$  shocks, respectively (see Harvey et al., 2009b for details on the construction of these estimates). Based on the finite sample properties of the procedure, the choice  $m = 0.10$  is recommended for practice. The proposed test is

$$U = \max \left\{ S_1, \left( \frac{cv_\xi^1}{cv_\xi^0} \right) S_0 \right\}$$

where  $cv_\xi^1$  and  $cv_\xi^0$  denote the  $\xi$ -level asymptotic critical values of  $S_1$  under  $I(1)$  errors and  $S_0$  under  $I(0)$  errors, respectively. The computed value of  $U$  is then compared with  $\kappa_\xi cv_\xi^1$ , where  $\kappa_\xi = cv_\xi^{\max} / cv_\xi^1$ , where  $cv_\xi^{\max}$  is the  $\xi$ -level critical value from the limit distribution of  $\max \left\{ S_1, \left( \frac{cv_\xi^1}{cv_\xi^0} \right) S_0 \right\}$ .

## A.2 Procedures for Selecting the Number of Breaks

### A.2.1 The Kejriwal and Perron (2010) Sequential Procedure for Slope Breaks

Building on the work of Perron and Yabu (2009a), Kejriwal and Perron (2010) propose a sequential procedure that allows one to obtain a consistent estimate of the true number of breaks irrespective of whether the errors are  $I(1)$  or  $I(0)$ . The first step is to conduct a test for no break versus one break. Conditional on a rejection, the estimated break date is obtained by a global minimization of the sum of squared residuals. The strategy proceeds by testing each of the two segments (obtained using the estimated partition) for the presence of an additional break and assessing whether the maximum of the tests is significant. Formally, the test of one versus two breaks is expressed as

$$ExpW(2|1) = \max_{1 \leq i \leq 2} \{ ExpW^{(i)} \}$$

where  $ExpW^{(i)}$  is the one break test in segment  $i$ . We conclude in favor of a model with two breaks if  $ExpW(2|1)$  is sufficiently large.<sup>16</sup>

### A.2.2 The Harvey et al. (2010) Sequential Procedure for Level Breaks

Harvey et al. (2010) also propose the following sequential procedure for selecting the number of level breaks in addition to the  $U$  test discussed above. First, if  $S_1 > \kappa_\xi cv_\xi^1$ , denote  $\tilde{t}_1 = \operatorname{argmax}_{t \in \Lambda_1} (\hat{\omega}_v)^{-1} T^{-1/2} \left| M_{t, [mT]} - \hat{\beta}_0 [\frac{m}{2} T] \right|$ . Then, denoting  $\Lambda_2 = [\tilde{t}_1 - [mT] + 1, \tilde{t}_1 + [mT] - 1]$ , if  $\max_{t \in \Lambda_1 - \Lambda_2} (\hat{\omega}_v)^{-1} T^{-1/2} \left| M_{t, [mT]} - \hat{\beta}_0 [\frac{m}{2} T] \right| \leq cv_\xi^1$ , we conclude that the procedure based on  $S_1$  selects one break; otherwise, two breaks are selected. The number of breaks is denoted  $n'_1$ . A similar procedure based on  $S_0$  gives  $n'_0$  breaks. The number of breaks selected by the sequential procedure based on  $U$  is then  $n_U = \max(n'_1, n'_0)$ . For a given number of breaks, consistent estimates of the break dates in the presence of  $I(1)$  errors are also suggested (See Harvey et al., 2010 for details).

## A.3 Robust Procedures for Trend Significance

For testing the significance of the trend coefficient in a stable linear trend model, we again use procedures that are robust to whether the errors are  $I(0)$  or  $I(1)$ . These are proposed in Harvey et al. (2007) and Perron and Yabu (2009a) and are similar to the robust procedures discussed earlier with regard to tests for the stability of the trend function. There are, however, certain differences and it is thus useful to provide a brief description of these procedures.

The model is

$$y_t = \mu_0 + \beta_0 t + u_t \quad (\text{A.8})$$

with  $u_t$  generated as  $u_t = \alpha u_{t-1} + v_t$ ,  $t = 2, \dots, T$ ,  $u_1 = v_1$ . The goal is to obtain confidence intervals for  $\beta_0$  that are asymptotically valid whether  $\alpha = 1$  or  $|\alpha| < 1$ .

### A.3.1 The Harvey et al. (2007) Procedure

Harvey et al. (2007) propose the estimate

$$\hat{\beta}_0(\text{HLTa}) = \frac{\{1 - \varphi(U, S)\} \hat{\beta} \tilde{\omega} + \varphi(U, S) \tilde{\beta} \hat{\omega}}{\{1 - \varphi(U, S)\} \tilde{\omega} + \varphi(U, S) \hat{\omega}} \quad (\text{A.9})$$

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<sup>16</sup>For the general model with  $k$  breaks, the estimated break points are obtained by a global minimization of the sum of squared residuals. The strategy proceeds by testing each  $k + 1$  segment (obtained using the estimated partition) for the presence of an additional break. The test thus amounts to the application of  $k + 1$  tests of the null hypothesis of no change versus the alternative hypothesis of a single change and assessing whether the maximum is significant. See Kejriwal and Perron (2010) for more details.

with the associated  $100(1 - \alpha)\%$  confidence interval

$$\hat{\beta}_0(\text{HLTa}) \pm c_{\alpha/2} \frac{\hat{\omega}\tilde{\omega}}{\{1 - \varphi(U, S)\}\tilde{\omega} + \varphi(U, S)\hat{\omega}} \quad (\text{A.10})$$

In (A.9) and (A.10),  $\hat{\beta}$  and  $\tilde{\beta}$  are the OLS estimates from the levels regression (A.8) and its first-differenced version respectively, the quantity  $c_{\alpha/2}$  is such that  $P(x > c_{\alpha/2}) = \alpha/2$  for  $x \sim N(0, 1)$ ,  $\hat{\omega}^2$  and  $\tilde{\omega}^2$  are long-run variance estimators using the quadratic spectral kernel with Newey and West (1994) automatic bandwidth selection adopting a non-stochastic prior bandwidth of  $[4(T/100)^{2/25}]$  and the weight function  $\varphi(U, S) = \exp(-.00025(U/S)^2)$ . The quantity  $U$  is the local GLS-detrended augmented Dickey-Fuller  $t$ -ratio of Elliott et al. (1996) i.e. the usual  $t$ -ratio for testing  $\alpha^* = 0$  in the regression equation

$$\Delta\tilde{u}_t = \alpha^*\tilde{u}_{t-1} + \sum_{j=1}^p \phi_j \Delta\tilde{u}_{t-j} + \tilde{e}_t$$

where  $\tilde{u}_t$  are the local GLS detrended residuals obtained from the regression of  $(y_1, y_2 - \bar{\alpha}y_1, \dots, y_T - \bar{\alpha}y_{T-1})'$  on  $(z_1, z_2 - \bar{\alpha}z_1, \dots, z_T - \bar{\alpha}z_{T-1})'$ , where  $z_t = (1, t)'$  and  $\bar{\alpha} = 1 - \bar{c}/T$  with  $\bar{c} = -13.5$  and  $S$  is the standard KPSS stationarity test statistic

$$S = \frac{\sum_{t=1}^T \left( \sum_{i=1}^t \hat{u}_i \right)^2}{T^2 \hat{\omega}^2}$$

with  $\hat{u}_t$  being the residuals from OLS estimation of (A.8). Harvey et al. (2007) show that the confidence interval (A.10) is asymptotically valid regardless of whether the errors are  $I(1)$  or  $I(0)$ .

### A.3.2 The Perron and Yabu (2009a) Procedure

The procedure is quite similar in spirit to the Perron and Yabu (2009b) procedure so we omit the details and outline the main differences. The reader is referred to the Perron and Yabu (2009a) paper for further details. First, the residuals  $\hat{u}_t$  in (A.5) are now obtained from a regression of  $y_t$  on  $x_t = (1, t)'$ . Next, the super-efficient estimate  $\tilde{\alpha}_{MS}$  (obtained as discussed earlier) is used to estimate the quasi-differenced regression

$$\begin{aligned} (1 - \tilde{\alpha}_{MS}L)y_t &= (1 - \tilde{\alpha}_{MS}L)x_t'\Psi^0 + (1 - \tilde{\alpha}_{MS}L)u_t, \quad t = 2, \dots, T \\ y_1 &= x_1'\Psi^0 + u_1 \end{aligned}$$

where  $\Psi^0 = (\mu_0, \beta_0)'$ . Denote the estimate of  $\beta_0$  from this regression by  $\hat{\beta}_0(\text{PYa})$ . The, using the notation  $X^{FG} = (x_1^{FG}, x_2^{FG}, \dots, x_T^{FG})'$  with  $x_1^{FG} = (1, 1)'$ ,  $x_t^{FG} = [1 - \tilde{\alpha}_{MS}, t - \tilde{\alpha}_{MS}(t - 1)]$  for  $t = 2, \dots, T$ , a  $100(1 - \alpha)\%$  confidence interval for  $\beta_0$ , again valid for both  $I(1)$  and  $I(0)$  errors, is obtained as

$$\hat{\beta}_0(\text{PYa}) \pm c_{\alpha/2} \sqrt{\tilde{h}_v [(X^{FG'}X^{FG})^{-1}]_{22}} \quad (\text{A.11})$$

where the quantity  $c_{\alpha/2}$  is as defined in (A.10) and the estimate  $\tilde{h}_v$  is constructed in the same way as that for the Perron and Yabu (2009b) procedure discussed in section A.1.2.



## A.4 Unit Root Tests

### A.4.1 The Harris et al. (2009) Test

Harris et al. (2009) propose a test for a unit root in the presence of a possible trend break based on a GLS detrending procedure similar to that used by Elliott et al. (1996) in the stable trend case. Consider the model given by (1) and (2). The first step is to obtain an estimate of the break fraction by minimizing the sum of squared residuals from OLS estimation of the first differenced regression (A.1). This is denoted  $\tilde{\lambda}_1$ . Applying a quasi-differenced transformation to (1) yields

$$(1 - \alpha(\tilde{\lambda}_1)L)y_t = (1 - \alpha(\tilde{\lambda}_1)L)x'_{L1,t}(\tilde{\lambda}_1)\Psi + (1 - \alpha(\tilde{\lambda}_1)L)u_t, \quad \alpha(\tilde{\lambda}_1) = 1 - \frac{c(\tilde{\lambda}_1)}{T} \quad (\text{A.12})$$

where  $c(\tilde{\lambda}_1)$  denotes the value at which the asymptotic Gaussian local power envelope for a break fraction  $\tilde{\lambda}_1$  at a given significance level has power equal to .50. Letting  $\tilde{\Psi}_{c(\tilde{\lambda}_1)}$  and  $\tilde{u}_{t,c(\tilde{\lambda}_1)}$  denote the OLS estimate and residuals from (A.12), the next step is to estimate the Augmented Dickey -Fuller type regression

$$\Delta \tilde{u}_{t,c(\tilde{\lambda}_1)} = \phi \tilde{u}_{t-1,c(\tilde{\lambda}_1)} + \sum_{j=1}^{k_1} \delta_j \Delta \tilde{u}_{t-j,c(\tilde{\lambda}_1)} + e_{k_1,t}, \quad t = k_1 + 2, \dots, T \quad (\text{A.13})$$

The unit root statistic, denoted  $H$ , is then the  $t$ -statistic for  $\phi = 0$  in (A.13). The lag length  $k_1$  is selected using the modified Akaike Information Criterion (MAIC) proposed in Ng and Perron (2001).

### A.4.2 The Carrion-i-Silvestre et al. (2009) Tests

Carrion-i-Silvestre et al. (2009) propose an alternative testing procedure which allows for multiple structural breaks in the level and/or slope of the trend function under both the null and alternative hypotheses. The tests are extensions of the  $M$  class of tests analyzed in Ng and Perron (2001) and the feasible point optimal statistic of Elliott et al. (1996). We will provide a brief description of the tests for the two breaks model. The model is

$$y_t = \mu_0 + \beta_0 t + \mu_1 DU_{1t} + \beta_1 DT_{1t} + \mu_2 DU_{2t} + \beta_2 DT_{2t} + u_t$$

where  $DU_{it} = I(t > T_i)$ ,  $DT_{it} = (t - T_i)I(t > T_i)$  ( $i = 1, 2$ ) and the errors  $u_t$  are generated as in (2). First, the estimates of the break fractions  $\lambda = (\lambda_1, \lambda_2)$  and the regression parameters are obtained by minimizing the sum of squared residuals from the quasi-differenced regression analogous to (A.12). The sum of squared residuals evaluated at these estimates is denoted  $S(\alpha(\hat{\lambda}), \hat{\lambda})$  with  $\alpha(\hat{\lambda}) = 1 - \frac{c(\hat{\lambda})}{T}$ . The feasible point optimal statistic is then

$$P_T^{gls}(\hat{\lambda}) = \frac{S(\alpha(\hat{\lambda}), \hat{\lambda}) - \alpha(\hat{\lambda})S(1, \hat{\lambda})}{s^2(\hat{\lambda})}$$

where  $s^2(\hat{\lambda})$  is an autoregressive estimate of the spectral density of  $v_t$  at frequency zero:

$$s^2(\hat{\lambda}) = s_{ek}^2 / (1 - \hat{b}(1))^2 \quad (\text{A.14})$$

where  $s_{ek}^2 = (T - k)^{-1} \sum_{t=k+1}^T \hat{e}_{tk}^2$ ,  $\hat{b}(1) = \sum_{j=1}^k \hat{b}_j$ , with  $\hat{b}_j$  and  $\hat{e}_{tk}$  obtained from the OLS estimation of

$$\Delta \tilde{y}_t = b_0 \tilde{y}_{t-1} + \sum_{j=1}^k b_j \Delta \tilde{y}_{t-j} + e_{tk}$$

with

$$\tilde{y}_t = y_t - \hat{\Psi}_2' x_{L2,t}(\hat{\lambda}), \quad x_{L2,t}(\hat{\lambda}) = \left\{ 1, t, DU_{1t}(\hat{\lambda}), DU_{2t}(\hat{\lambda}), DT_{1t}(\hat{\lambda}), DT_{2t}(\hat{\lambda}) \right\} \quad (\text{A.15})$$

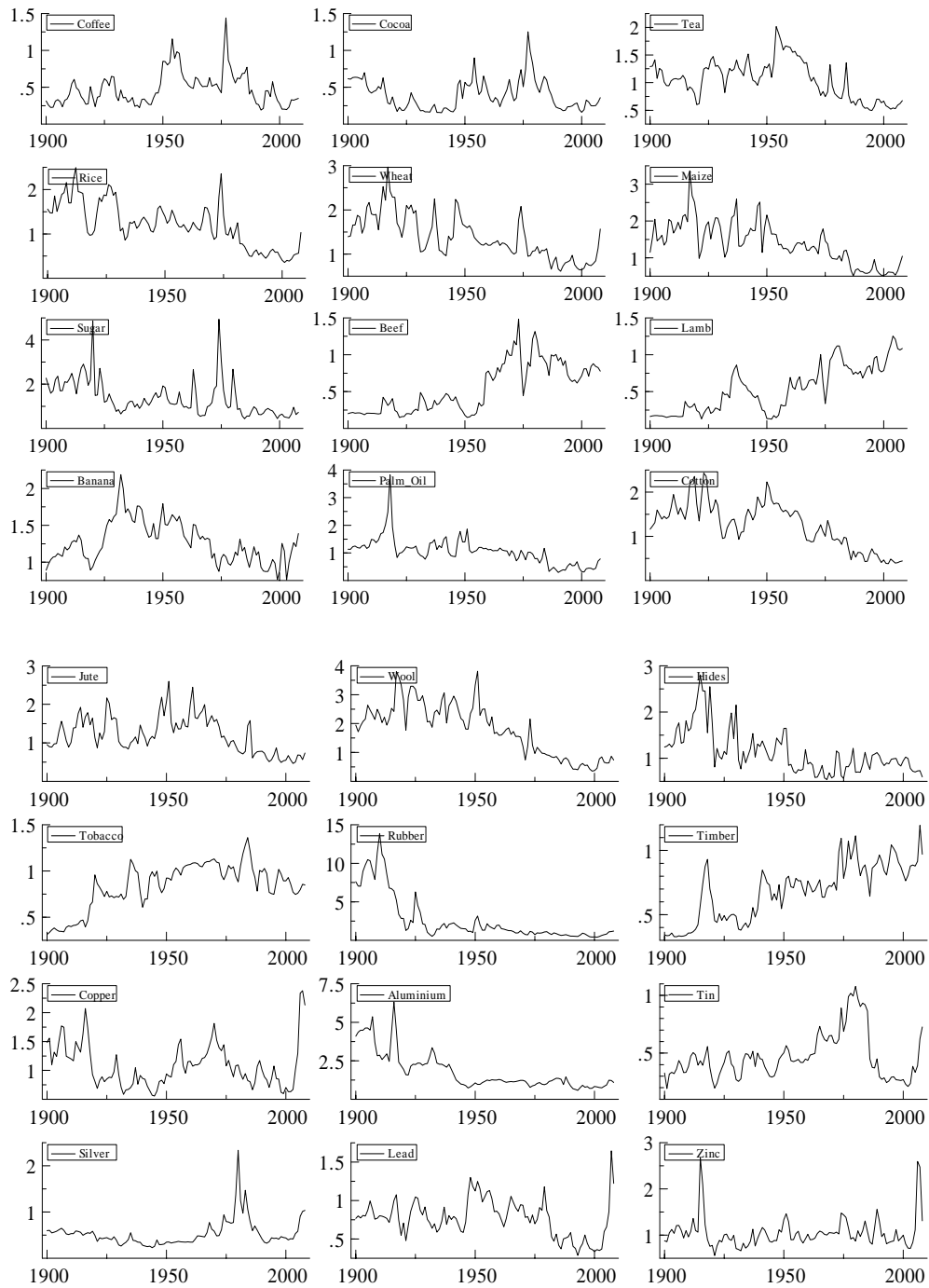
and  $\hat{\Psi}_2$  being the OLS estimate obtained from the quasi-differenced regression.

Carrion-i-Silvestre et al. (2009) also consider extensions of the  $M$ -class of tests analyzed in Ng and Perron (2001). These are given by

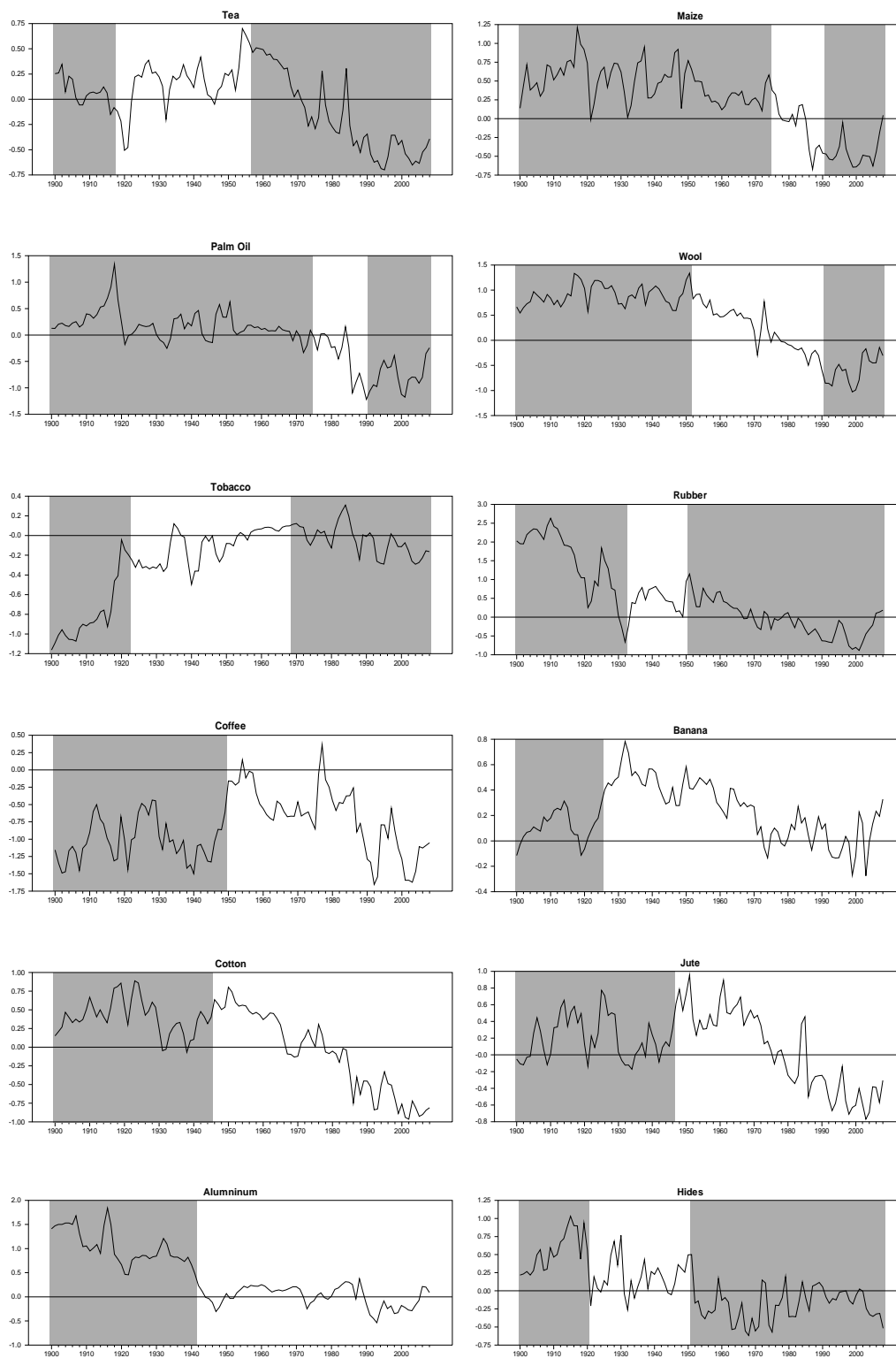
$$\begin{aligned} MZ_{\alpha}^{gls}(\hat{\lambda}) &= (T^{-1} \tilde{y}_T^2 - s^2(\hat{\lambda})) (2T^{-2} \sum_{t=2}^T \tilde{y}_{t-1}^2)^{-1} \\ MSB^{gls}(\hat{\lambda}) &= (T^{-2} \sum_{t=2}^T \tilde{y}_{t-1}^2)^{1/2} / s^2(\hat{\lambda}) \\ MZ_t^{gls}(\hat{\lambda}) &= (T^{-1} \tilde{y}_T^2 - s^2(\hat{\lambda})) (4s^2(\hat{\lambda}) T^{-2} \sum_{t=2}^T \tilde{y}_{t-1}^2)^{-1/2} \\ MP_T^{gls}(\hat{\lambda}) &= [c^2(\hat{\lambda}) T^{-2} \sum_{t=2}^T \tilde{y}_{t-1}^2 + (1 - c(\hat{\lambda})) T^{-1} \tilde{y}_T^2] / s^2(\hat{\lambda}) \end{aligned} \quad (\text{A.16})$$

where  $s^2(\hat{\lambda})$  and  $\tilde{y}_t$  are as defined in (A.14) and (A.15). These test statistics (with obvious modifications) are also used to detect pure level breaks with a stable slope parameter. See Carrion-i-Silvestre et al. (2009) for details.

**Figure 1.** Grilli Yang Index: 1900-2008.



**Figure 2.** Regimes for Commodity Prices that contain Structural Breaks.



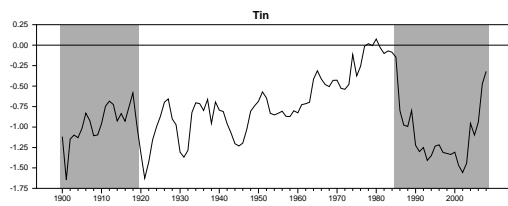


Table 1: Robust Tests for Breaks in Trend

Commodity\Test	Slope Breaks				Level Breaks	
	$ExpW$	$ExpW(2 1)$	$t_\eta$	#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	0.42	0
Wheat	0.37	-	1.51	0	0.31	0
Maize	4.25*	3.89*	3.52*	2	-	-
Sugar	-0.28	-	1.97	0	0.35	0
Beef	-0.22	-0.22	1.47	0	0.49	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	0.56*	2
Tobacco	3.28*	634.40*	2.73	2	-	-
Rubber	0.21	200.89*	3.35*	2	-	-
Timber	-0.19	-	2.41	0	0.44	0
Copper	0.02	-	1.76	0	0.36	0
Aluminium	0.05	-0.14	3.65*	1	-	-
Tin	-0.26	-	1.47	0	0.54*	2
Silver	-0.03	0.14	1.86	0	0.37	0
Lead	-0.17	-	1.77	0	0.35	0
Zinc	0.11	-	2.47	0	0.37	0

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

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

Commodity\Estimate	One Break			Two Breaks				
	$\hat{\beta}_0$	$\sum_{i=0}^1 \hat{\beta}_i$	Date	$\hat{\beta}_0$	$\sum_{i=0}^1 \hat{\beta}_i$	$\sum_{i=0}^2 \hat{\beta}_i$	Date 1	Date 2
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: Trend Coefficient Estimates (PYa) - Commodities with Breaks in Slope (in Percentage)

Commodity\Estimate	One Break			Two Breaks				
	$\hat{\beta}_0$	$\sum_{i=0}^1 \hat{\beta}_i$	Date	$\hat{\beta}_0$	$\sum_{i=0}^1 \hat{\beta}_i$	$\sum_{i=0}^2 \hat{\beta}_i$	Date 1	Date 2
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: Trend Coefficient Estimates - Commodities with Pure Level Shifts (in Percentage)

Commodity\Estimate	$\hat{\beta}_0$	Date 1	Date 2
Hides	0.43 [0.03,0.83]	1920	1951
Tin	1.70 [1.29,2.11]	1919	1985

Table 5: Trend Coefficient Estimates - Commodities with a Stable Linear Trend (in Percentage)

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]
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]
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]

Table 6: Prevalence of Trends

Commodity	HLTa			PYa		
	$\Psi(-)$	$\Psi(+)$	$\Psi(.)$	$\Psi(-)$	$\Psi(+)$	$\Psi(.)$
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	1.000	0.000	0.000	1.000	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	1.000	0.000	0.000	1.000	0.000
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

Table 7: Unit Root Tests

Commodity\Test	Slope Breaks						No Slope Breaks				
	$MZ_{\alpha}^{gls}$	$MSB^{gls}$	$MZ_t^{gls}$	$P_T^{gls}$	$MP_T^{gls}$	$H$	$MZ_{\alpha}^{gls}$	$MSB^{gls}$	$MZ_t^{gls}$	$P_T^{gls}$	$MP_T^{gls}$
Coffee	-19.70	0.16*	-3.11*	4.45	4.34	-3.44*	-	-	-	-	-
Cocoa	-	-	-	-	-	-	-8.82	0.24	-2.09	10.90	10.37
Tea	-34.46*	0.12*	-4.09*	6.48*	6.29*	-	-	-	-	-	-
Rice	-	-	-	-	-	-	-19.61*	0.16*	-3.07*	5.17*	5.03*
Wheat	-	-	-	-	-	-	-17.05*	0.16*	-2.82*	6.38*	5.94*
Maize	-24.82*	0.14*	-3.46*	9.40	8.60	-	-	-	-	-	-
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*
Banana	-5.85	0.29	-1.71	19.36	16.80	-1.67	-	-	-	-	-
Palmoil	-32.82*	0.12*	-3.98*	6.93*	6.62*	-	-	-	-	-	-
Cotton	-24.87*	0.14*	-3.51*	4.22*	4.10*	-1.83	-	-	-	-	-
Jute	-18.59	0.16*	-3.04*	4.35	4.28	-2.87	-	-	-	-	-
Wool	-18.60	0.16	-2.94	11.62	11.45	-	-	-	-	-	-
Hides	-	-	-	-	-	-	-14.27*	0.18*	-2.51*	7.88	7.30
Tobacco	-12.40	0.20	-2.48	16.45	15.73	-	-	-	-	-	-
Rubber	-24.44*	0.14*	-3.41*	9.42	8.91	-	-	-	-	-	-
Timber	-	-	-	-	-	-	-20.70*	0.15*	-3.20*	4.62*	4.49*
Copper	-	-	-	-	-	-	-11.53	0.19	-2.18	9.52	9.05
Aluminium	-15.33	0.18	-2.71	6.42	6.49	-2.86	-	-	-	-	-
Tin	-	-	-	-	-	-	-19.72*	0.15*	-3.06*	6.38	5.11
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.