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Commodity market instability and asymmetries in developing countries:

Development impacts and policies

Les asymétries et l'instabilité du marché des matières premières dans les

pays en développement : politiques et impacts sur le développement

Asymmetries in Commodity Price Behaviour

(Work in Progress: Not to be quoted)

By

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Abstract. A number of developing 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 dynamic patterns of real commodity prices to help formulate their economic policies. Motivated by these considerations, this paper evaluates the time series properties of primary commodity prices by applying a range of new econometric techniques to a set of primary commodity prices using an updated version of the so-called Grilli-Yang Index over the period 1900-2010. A new set of powerful unit root tests allowing for asymmetric behaviour as described by Deaton and Laroque (1992) and by Prebisch (1950) is applied to determine whether the underlying commodity price series can be characterized as a stationary processes and whether asymmetry exists. Our results find that there is considerable evidence of asymmetry in commodity price dynamics.

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1. Introduction

Many developing countries are dependent on commodity prices as their main source of income. Additional income from commodity price booms can help the economies of low income countries that are reliant on a single or handful of commodities, while a slump in commodity prices can lead to a significant loss of income. Commodity prices are known to be highly volatile (Deaton and Laroque 1992) and therefore the large upswings and downswings in commodity prices leave countries dependent on such commodities highly vulnerable. For example, policy prescriptions can be potentially catastrophic if the income from a commodity boom is diagnosed as permanent when in actual fact turns out to be temporary. It is no surprise therefore, why the dynamic properties of commodity prices in relation to manufactures has been of great interest in the trade and development economics literature. Economists and policy makers have been interested to study the trends, volatility, cycles and persistence in commodity price data so that appropriate policy prescriptions can be made for developing countries dependent on commodities. Though there has been some progress in modelling some aspects of commodity price behaviour, the understanding and forecasting of commodity prices remains inadequate.

Commodity prices have been described by Deaton and Laroque (1992) to be dominated by long periods of doldrums punctuated by sharp upward spikes. This asymmetry, noted by Deaton (1999) shows that commodity prices record positive skewness, in that the prices show relatively more upward peaks than matching troughs. Further, Deaton (1999) argues that cycles in commodity prices are not persistent, that is in the long run, any shocks are ultimately transitory in nature and that real commodity prices revert to trend or a long run constant. The reason that commodity prices should be stationary, is due to the nature of production, storage and arbitrage (Wang and Tomek 2007). The empirical evidence for stationarity in commodity prices, provides mixed results. However, most studies conclude, that commodity prices tend to persist for a considerable length of time (Cashin et al., 2000) and that this persistence can change over time (Ghoshray 2013).

A large volume of studies have dealt with the issue of trends in commodity prices. That is, whether these prices have a tendency to decline over time in the long run as propounded in two separate studies by Prebisch (1950) and Singer (1950), popularly known as the Prebisch-Singer hypothesis. Prebisch (1950) offered a supply side theory. He argued that strong labour unions in countries that export manufactured 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 labour unions that are not able to increase wages during expansions and cannot prevent wages from falling during times of recessions. This causes primary commodity prices to increase by less than manufacturing goods prices during expansions but fall more during downturns. This type of asymmetry has not been extensively studied in the literature. The only empirical studies to my knowledge that has made a direct test for this is by Thirlwall and Bergevin (1985) and more recently Cashin et. al. (2002). Using a slope dummy variable regression, Thirlwall and Bergevin (1985) find that there is little evidence for the Prebisch (1950) hypothesis, that the terms of trade of primary commodities are subject to asymmetrical movements between period when prices of manufactured goods are rising and falling. Cashin et. al. (2002) adopt the Bry and Boschan (1971) algorithm to examine the duration and magnitude of cycles in international commodity prices. They find evidence of asymmetry in commodity prices as the duration of slumps exceeds the duration of booms. While the magnitude of price falls in the slump period is slightly larger than the magnitude of the price boom, the rate of change of prices during the prices during the boom phase is slightly faster than that of the slump phase.

Very recently it has been shown that the support for the Prebisch-Singer hypothesis has begun to wane, especially if one were to take into account structural breaks and calculate the 'prevalence' of a negative trend (see for example, Kellard and Wohar (2006); Ghoshray (2011), Ghoshray et. al. (2014)). In another related study, Deaton and Laroque (2003) set out a model that shows prices of commodities in developing countries can be characterised 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 labour 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 labour, prices will eventually revert to the constant subsistence level. However, in their paper, Deaton and Laroque (2003) find no empirical evidence for their results.

There are many studies that have highlighted the issue of cycles in commodity prices (Ertin and Ocampo (2012); Carter et. al. (2011); Deaton (1999)). It has been argued that inventory holding behaviour provides a clue to explaining the commodity price cycles along with supply and demand shocks, macroeconomic shocks and policy responses. For example, in the lead up to the spike, inventories get run down and prices begin to increase. When the stocks reach a critical point, the price spikes. Over the ensuing time period stocks get replenished and the price declines. This decline occurs more gradually than the spike (Carter et. al. 2011). The directional asymmetry in this case is opposite to the type of asymmetry described by Prebisch (1950).

The upshot from these different studies is that commodity prices are likely to exhibit the following properties: (a) commodity prices are expected to be stationary. However, while there are a growing number of studies that propound this behaviour, the evidence on this still remains mixed. (b) Commodity prices may or may not be exhibit a long run secular trend. (c) The fluctuations of commodity prices over time around the underlying trend or long run constant are likely to be asymmetric. The asymmetry could be based on which 'state' the price is: that is, above or below the long run trend or constant. If price is below the constant it is likely to persist, and if above, then there is likely to be a quick reversion to the long run value. This follows from Deaton and Laroque (1992) that commodity prices exhibit periods of doldrums punctuated by harp spikes. Alternatively, the speed of momentum at which prices fluctuate around the long run trend may differ, depending on whether the prices are increasing or decreasing relative to the trend. This asymmetry based on 'direction' has been discussed by Prebisch (1950) and separately by Deaton (1999) and Carter et. al. (2011) where the underlying asymmetric adjustment is different.

There is unfortunately little empirical evidence relating directly to these asymmetric properties of commodity price cycles. As highlighted by Cashin et. al. (2002), which is one of the few studies that have empirically examined the nature of cycles in commodity prices, the consequences of rapid transitions from a period of price boom to a period of price slump are one of the most challenging issues facing policymakers in the many developing countries that are reliant on the exports of such commodities. In this paper we try to understand whether these cycles are indeed asymmetric as highlighted by the observations made by Carter et. al. (2011), Deaton (1999) and the seminal paper of Prebisch (1950). Knowledge of these features of cycles would further our understanding of the nature of commodity price booms and slumps.

Motivated by these considerations, this paper evaluates the time series properties of primary commodity prices by applying a range of new econometric techniques to a set of primary commodity prices using an updated version of the so-called Grilli-Yang Index over the period 1900-2010. A new set of powerful unit root tests allowing for the nature of asymmetric behaviour as described by Deaton and Laroque (1992) and by Prebisch (1950) is applied to determine whether the underlying commodity price series can be characterised as a stationary processes and whether asymmetry exists.

The rest of the paper is organized as follows: Section 2 provides a description the econometric methodology; Section 3 presents the data; Section 4 describes the empirical results and finally Section 5 concludes.

2. Econometric Model

The two types of asymmetry discussed above can be captured by employing econometric procedures due to Enders and Granger (1998) and the more recently developed and powerful method of Lee et. al. (2011). In order to address the problem that of the possible presence of a unit root that may exist in commodity prices we test for unit roots allowing for such asymmetry. As a starting point we assume the benchmark linear model by specifying the commodity price to exhibit trend reverting behaviour by estimating the equation below:

$$P_t = \alpha + \beta t + \varepsilon_t \tag{1}$$

where P_t is the commodity price, α is an arbitrary constant, β denotes the trend parameter, and ε_t is the error term which may be serially correlated. To test whether the price is stationary, we estimate equation (1) using ordinary least squares and then apply an Augmented Dickey-Fuller (ADF) test, on the estimated residuals of equation (1) of the following kind:

$$\Delta\varepsilon_t = \gamma\varepsilon_{t-1} + \sum_{i=1}^{p} \psi_i \Delta\varepsilon_{t-i} + \omega_t \tag{2}$$

where ω_t is a white noise error term and p denotes the number of lags to whiten the residuals. The number of lags is selected using the Schwartz Bayesian Criteria (SBC). Rejecting the null hypothesis $H_0\colon(\gamma = 0)$ of a unit root implies that the residuals of (1) are stationary and that any shocks to the commodity price would have a transitory effect. Enders and Granger (1998), however, argue that the test for unit roots and its extensions are mis-specified if the adjustment is asymmetric. They consider two alternative specifications, called the threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) models which allow for asymmetric adjustment. Under the TAR specification, the estimated residuals ε_t from (1) can be written as:

$$\Delta\varepsilon_t = I_t\gamma_1\varepsilon_{t-1} + (1 - I_t)\gamma_2\varepsilon_{t-1} + \omega_t \tag{3}$$

where I_t is the Heaviside Indicator function such that:

$$I_t = \begin{cases} 1 \text{ if } \varepsilon_{t-1} \geq \tau \\ 1 \text{ if } \varepsilon_{t-1} < \tau \end{cases} \tag{4}$$

If the system is convergent, then the long run equilibrium value of the sequence is given by $\varepsilon_t = \tau$. To estimate the threshold τ we use the methodology proposed by Chan (1993)[1]. The sufficient conditions for the stationarity of ε_t are $\gamma_1 < 0$, $\gamma_2 < 0$, and $(1 + \gamma_1)(1 + \gamma_2) < 1$ (Petrucelli and Woolford 1984). In this case, the TAR model shows different rates of adjustments depending on the state of the disequilibrium.

In light of the asymmetry described by Deaton and Laroque (1992), that commodity prices are characterised by long periods of doldrums, punctuated by sharp spikes, the TAR model would show in this case, $-1 < \gamma_1 < \gamma_2 < 0$, so that the negative phase of the series ε_t will tend to be more persistent than the positive phase.

Under the M-TAR specification the Heaviside Indicator function is written as:

$$I_t = \begin{cases} 1 \text{ if } \Delta\varepsilon_{t-1} \geq \tau \\ 1 \text{ if } \Delta\varepsilon_{t-1} < \tau \end{cases} \tag{5}$$

In this case, the series ε_t exhibits more momentum in one direction than the other.

[1] The estimated residual series was sorted in ascending order. The largest and smallest 15% of the residual series were eliminated and each of the remaining 70% of the values were considered as possible thresholds. For each of these thresholds the equation was estimated using (3) and (4). The estimated threshold yielding the lowest residual sum of squares was chosen as the estimate of the optimal threshold.

In light of the asymmetry described by Prebisch (1950), we would expect, $|\gamma_1| < |\gamma_2|$, so that the M-TAR model exhibits little adjustment for $\Delta\varepsilon_{t-1} > \tau$ but substantial decay for $\Delta\varepsilon_{t-1} < \tau$. In other words, increases tend to be slow, but decreases tend to revert quickly back to the attractor irrespective of where disequilibrium is relative to the attractor. Alternatively, if we expect the asymmetry to be of the type described by Deaton (1999), Carter et. al. (2011), then we would expect, $|\gamma_1| > |\gamma_2|$, increases tend to be revert quickly to the attractor, but decreases tend to be relatively slow.

The Φ-statistic for the null hypothesis of stationarity of ε_t, that is, $H_0\colon (\gamma_1 = \gamma_2 = 0)$ is obtained from estimating equation (3) and compared to the critical values computed by Enders and Granger (1998). If we can reject the null hypothesis, it is possible to test for asymmetric adjustment since γ_1 and γ_2 converge to multivariate normal distributions (Tong 1990). The F statistic is used to test for the null hypothesis of symmetric adjustment, that is, $H_0\colon (\gamma_1 = \gamma_2)$. Diagnostic checking of the residuals is undertaken to ascertain whether the ω_t series is a white noise process using the Ljung-Box Q tests. If the residuals are correlated, equation (3) needs to be re-estimated in the following form, and the SBC is used to determine the lag length:

$$\Delta\varepsilon_t = I_t\gamma_1\varepsilon_{t-1} + (1 - I_t)\gamma_2\varepsilon_{t-1} + \sum_{i=1}^{p}\phi_i\Delta\varepsilon_{t-i} + \omega_t \tag{5}$$

A drawback of this procedure is that the TAR model suffers from low power in comparison to the ADF test. However, the M-TAR model has increased power in comparison to the ADF tests when the underlying data series is asymmetric. In this paper, we obviate this problem by relying on the LM based TAR and M-TAR method proposed by Lee et. al. (2011) which builds on the TAR and M-TAR models of Enders and Granger (1998) based on the LM unit root tests of Schmidt and Lee (1991) and Schmidt and Phillips (1992) with increased power. The advantage of this procedure is that the LM based tests are more powerful than the ADF tests regardless of whether the underlying model is symmetric or asymmetric (Lee et. al. 2011).

To briefly describe the procedure of Lee et. al. (2011), we consider the commodity price to be given by the following data generating process:

$$P_t = \delta' Z_t + e_t; \, e_t = \rho e_{t-1} + u_t \tag{6}$$

where Z_t contains the deterministic terms. In first differences (6) becomes:

$$\Delta P_t = \delta' \Delta Z_t + u_t \tag{7}$$

Let $\tilde{S}_t = P_t - \tilde{\upsilon} - Z_t \tilde{\delta}$ be the detrended price series where $\tilde{\delta}$ is estimated from (7) and $\tilde{\upsilon}$ is the restricted MLE $\left(\tilde{\upsilon} = P_1 - Z_1 \tilde{\delta}\right)$. In the spirit of the model put forward by Enders and Granger (1998), Lee et. al (2011) specify the LM TAR model as:

$$\Delta P_t = I_t \gamma_1 \tilde{S}_{t-1} + (1 - I_t)\gamma_2 \tilde{S}_{t-1} + \sum_{i=1}^{p} \phi_i \Delta \tilde{S}_{t-i} + \omega_t \tag{8}$$

Where I_t is the Heaviside Indicator function such that:

$$I_t = \begin{cases} 1 \text{ if } \tilde{S}_{t-1} \geq \tau \\ 1 \text{ if } \tilde{S}_{t-1} < \tau \end{cases} \tag{9}$$

and for the LM M-TAR model the Heaviside Indicator function is:

$$I_t = \begin{cases} 1 \text{ if } \Delta\tilde{S}_{t-1} \geq \tau \\ 1 \text{ if } \Delta\tilde{S}_{t-1} < \tau \end{cases} \tag{10}$$

In the spirit of Enders and Granger (1998), the statistic for the null hypothesis of stationarity is, $H_0: (\gamma_1 = \gamma_2 = 0)$ is obtained from estimating equation (8) and compared to the critical values computed by Lee et. al. (2011). If we can reject the null hypothesis, it is possible to test for asymmetric adjustment. The F statistic is used to test for the null hypothesis of symmetric adjustment, that is, $H_0: (\gamma_1 = \gamma_2)$.

3. Data

The primary commodity price series used in this paper have been taken from the well-known and much-used Grilli and Yang (1988) index of commodity prices, which is an index of 24 internationally-traded non-fuel commodity prices, ranging from 1900 to 1987. We use data until 2010 which has been extended by Pfaffenzeller[2]. Pfaffenzeller *et al.*'s (2007) technical note provides a full description of the calculation and extension of the Grilli-Yang data set. The data set that we choose to analyse consists of 13 primary commodity prices measured annually over the period 1900 – 2010 and deflated by the Manufacturing Unit Value (MUV) index. The reason for discarding 11 out of the 24 commodity prices is that these 11 out of 24 commodities contain either one or two breaks in the slope. Since the analysis of unit roots requires a reasonably long data set we conduct the remaining analysis on the 13 commodity prices that do not contain any structural breaks thereby making use of the full 111 observation points. The details of the empirical results of the structural break tests are described in the next section.

Figure 1 below provides plots of a selection of the commodity prices we analyse in this study. As described by Deaton (1999) it is possible to observe the prevalence of sharp upward spikes that tend to characterise these commodity prices.

[Figure 1 about here]

We find that the trend seems to be insignificant for commodities such as copper, cocoa and zinc. The long periods of doldrums punctuated by sharp upward spikes can be noticed in the price series. For example the spike in coffee prices in the late 1970s can be traced to the bad weather conditions that cut coffee output in Brazil. We can see from Table 1 the descriptive statistics of the price series. Almost all commodity series are shown to exhibit significant positive skewness and excess kurtosis, which is a feature of these prices.

[Table 1 about here]

The sign of positive skewness in the data implies that the positive spikes are more relative to the negative spikes. Excess kurtosis implies that outliers are present in the data. These features of commodity prices are what we would usually expect from the observations made by Deaton and Laroque (1992).

4. Empirical Results

[2] Accessed from http://www.stephan-pfaffenzeller.com/cpi.html

Since it has been widely documented in the literature that commodity prices may be characterised by structural breaks, we proceed at this stage by testing for the presence and the number of breaks in the trend function making use of several novel procedures allowing us to be agnostic to the persistence of the errors. 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 shown in Table 1. The test statistics $ExpW$ and t_η are the Kejriwal and Perron (2010) and Harvey et. al. (2007) tests for the null hypothesis of no slope break respectively. The U test is the Harvey et al. (2010) procedure for the null of no level breaks. Out of the 24 commodity prices, 11 commodities are found to contain at least one structural break. The sequential test statistic $ExpW(2|1)$ due Kejriwal and Perron (2010) of one versus two slope breaks is applied to those commodity prices where we find evidence of a single break (results not reported here). These results are not unique to this study; they have been compiled as Table 2 in Ghoshray et. al. (2014) and partly reproduced in this paper as Table 2.

[Table 2 about here]

Consequently no structural breaks were found for the remaining 13 commodity prices. These prices include beef, cocoa, copper, hides, lamb, lead, rice, silver, sugar, timber, tin, wheat and zinc. The subsequent analysis is carried out on these 13 prices so that we can make use of the full sample of 111 observations.

To facilitate comparison with the linear symmetric model, we conduct a simple ADF test on the commodity prices to determine whether the commodity prices are trend stationary, and if so, to find evidence as to whether the adjustment to trend is symmetric. The results are shown in the last column of Table 2. The null hypothesis of a unit root as given by (2) can be rejected in 6 out of the 13 commodity prices. These are rice, sugar, wheat, hides, timber and zinc. This would imply, that for these commodities any exogenous shocks would be transitory in nature.

However, following Enders and Granger (1998), the test for unit roots and its extensions are mis-specified if the underlying adjustment process is asymmetric. To this end, we apply the TAR model to capture any signs of asymmetry that characterise commodity prices as dominated by long periods of doldrums punctuated by sharp upward spikes (Deaton and Laroque, 1992). This would mean that when prices are below their long run trend or equilibrium, they tend to be persistent, while when they are above the trend, then those periods are relatively short-lived. The results are shown in Table 3 below.

[Table 3 about here]

First of all, the signs and magnitude of the adjustment process show that the negative phase of commodity prices tends to be more persistent than the positive phase in only 5 (copper, lamb, lead, tin, timber) out of the 13 commodities. However, when testing for a unit root, only lamb and timber reject the null hypothesis of a unit root, thereby permitting us to test for the null of symmetry. Moving on to the test for symmetry we find that for both commodities (lamb, timber), we cannot reject the null thereby concluding there is no significant difference between the two adjustment parameters are γ_1 and γ_2. Interestingly, we find that hides and rice display stationary process but with adjustment coefficients showing magnitudes that are contrary to those propounded by Deaton and Laroque (1992). However, in the case of hides, there is no asymmetry, but we do find asymmetry to exist for rice. In all cases the Ljung Box Q statistic shows that there is no evidence of serial correlation in the residuals. In summary we conclude there is no evidence (except for rice) of asymmetry using the TAR model due to Enders and Granger (1998) and therefore the dynamic behaviour of commodity prices described by Deaton and Laroque (1992) does not hold in this case. Further, (except for rice) since no asymmetry is found to exist in the data, the power of the TAR model again the linear symmetric ADF test would be lower.

Moving on to the M-TAR model we aim to capture the asymmetry that has been put forward by Prebisch (1950), that commodity price downswings are likely to be more pronounced than upswings, which would contribute to the deterioration of the long run trend in commodity prices. The results are shown in Table 4 below.

[Table 4 about here]

The signs and magnitude of the adjustment process show that the negative momentum of commodity prices tends to be greater than the positive momentum in only 5 (copper, lamb, lead, sugar, timber) out of the 13 commodities. However, when testing for a unit root, only lamb, sugar and timber reject the null hypothesis of a unit root, thereby permitting us to test for the null of symmetry. Moving on to the test for symmetry we find that for two commodities (lamb, sugar), we cannot reject the null thereby concluding there is no significant difference between the two adjustment parameters, γ_1 and γ_2. However, we find evidence of asymmetry as argued by Prebisch (1950) in the case of timber. Interestingly we find beef, cocoa, rice, silver, wheat and zinc to reject the null hypothesis of a unit root and then to conclude that the underlying process of adjustment is asymmetric. For these 5 commodities, the adjustment process is asymmetric but contrary to the type of asymmetry propounded by Prebisch (1950). Rather, the nature of asymmetry relates to the type documented in Carter et. al. (2011). The Ljung Box Q statistic shows that there is no evidence of serial correlation in the residuals. In summary we conclude there is evidence of asymmetry using the M-TAR model due to Enders and Granger (1998) for 8 out of the 13 commodities chosen in this study. However, the dynamic behaviour put forward by Prebisch (1950) only holds for timber and zinc. For those eight commodities where we find asymmetry to exist in the data, the power of the M-TAR model again the linear symmetric ADF test would be higher. This is not true however, for the remaining 5 commodities.

As discussed earlier, the updated TAR and M-TAR tests due to Lee et. al. (2011) have more power irrespective of whether the underlying data series is asymmetric or not. We therefore apply these LM type tests to the data. The results of the LM TAR model are given in Table 5 below:

[Table 5 about here]

The signs and magnitude of the adjustment coefficients are broadly similar, and as with the Enders and Granger (1998) TAR model the null hypothesis of a unit root can be rejected for the same commodities with only rice showing a trend stationary process with asymmetric adjustment which is contrary to the type of asymmetry put forward by Deaton and Laroque (1992). By employing the more powerful LM TAR model due to Lee et. al. (2011) we do not find any difference in our conclusions when compared to the Enders and Granger (1998) TAR model.

The results however change when we consider the LM M-TAR model due to Lee et. al. (2011). The results are shown in Table 6 below:

[Table 6 about here]

With the M-TAR model due to Enders and Granger (1998) we found 8 out of the 13 commodities to show a trend stationary process with asymmetric adjustment. However, with the LM M-TAR model we find 10 out of the 13 commodity prices to display a trend stationary process with asymmetric adjustment. The extra two commodities that we find to exhibit this behaviour are tin and copper. Out of these two commodities, only copper shows the support that was propounded by Prebisch (1950). In summary, we can conclude that while 10 commodities show support for the M-TAR type asymmetric adjustment, only three of these commodities (timber, copper, zinc) show support for the view of Prebisch (1950). The remaining seven commodities show support for the reasons documented in Carter et. al. (2011). However, there is significant evidence of asymmetry in general. The substantial evidence of asymmetry that is found where increases in commodity prices tend to revert quickly to the attractor and decreases tend to be slower, warrants further research as to what might explain this dynamic behaviour.

5. Conclusion

While a great deal of attention has been paid to the issue of trends in primary commodity prices, there is little empirical evidence relating directly to the possible asymmetric properties of commodity price cycles. Deaton and Laroque (1992) in an influential paper have described commodity prices to have long periods of doldrums punctuated by sharp upward spikes. This would imply that in relation to a threshold, commodity prices would be persistently below the threshold and when prices move above it then the movement would be sharp, in the sense that it would revert quickly back to the attractor or long run intertemporal equilibrium. To address this behaviour we adopt the TAR model assuming that commodity prices will be stationary. Further, as propounded by Prebisch (1950) in his influential study, primary commodity prices relative to the prices of manufactured goods would fall during cyclical downturns by more than they would rise during cyclical upturns. To test this type of asymmetry related to speed of adjustment, we make use of the M-TAR model, whereby asymmetry is addressed by suggesting that there is more momentum in price adjustment depending on whether prices are increasing or decreasing. Applying the Enders and Granger method (1998) and the more recently developed and powerful method of Lee et. al. 2011, we find commodity prices to broadly exhibit stationary behaviour with considerable evidence of asymmetries. However, while asymmetries exist, the type of asymmetric behaviour seems to be contrary to what Deaton and Laroque (1992) and Prebisch (1950) propound. In case of the momentum type asymmetry there seems to be some support to the findings of Cashin et. al. (2002). Overall, asymmetries do exist and their effect on developing countries can have non-trivial effects which merit further attention.

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TABLES

Table 1. Descriptive Statistics

	Skewness	Kurtosis
Beef	0.50**	-0.87*
Cocoa	1.17***	2.17***
Copper	1.24***	1.67***
Hides	1.38***	1.92***
Lamb	0.37	–0.97**
Lead	0.32	0.49
Rice	0.26	–0.55
Silver	2.86***	11.76***
Sugar	1.88***	5.48***
Timber	–0.09	–0.93*
Tin	1.20***	0.98**
Wheat	0.62**	–0.09
Zinc	2.84***	10.52***

***, and ** denote significance at the 1% and 5% levels respectively

Table 2. Slope and Level Break Test Results

	ExpW	ETA	U	# Breaks	ADF-t
Beef	0.22	1.47	0.48	0	–3.04
Cocoa	–0.07	1.39	0.51	0	–2.44
Copper	0.02	1.76	0.34	0	–2.36
Hides	0.56	2.16	0.50	0	–5.23***
Lamb	–0.26	1.16	0.45	0	–3.12
Lead	–0.17	1.77	0.37	0	–2.62
Rice	–0.25	2.07	0.32	0	–4.38***
Silver	–0.03	1.86	0.38	0	–2.06
Sugar	–0.28	1.97	0.41	0	–4.64***
Timber	–0.19	2.41	0.45	0	–3.76**
Tin	–0.26	1.47	0.51	0	–2.54
Wheat	0.37	1.51	0.33	0	–5.07***
Zinc	0.11	2.47	0.41	0	–4.76***

***, and ** denote significance at the 1% and 5% levels respectively

Table 3. TAR Model

	γ_1	γ_2	$H_0: (\gamma_1 = \gamma_2 = 0)$	$H_0: (\gamma_1 = \gamma_2)$	LB-Q
Beef	–0.21 (2.49)	–0.12 (1.91)	4.95	NA	2.82 [0.58]
Cocoa	–0.23 (2.79)	–0.06 (1.12)	4.32	NA	0.47 [0.97]
Copper	–0.07 (1.15)	–0.18 (2.29)	3.30	NA	3.38 [0.49]
Hides	–0.47 (3.98)	–0.32 (2.83)	10.84***	0.96 [0.32]	7.33 [0.12]
Lamb	–0.18 (2.19)	–0.25 (3.36)	7.10***	0.41 [0.52]	0.37 [0.98]
Lead	–0.07 (0.98)	–0.21 (2.77)	4.33	NA	2.64 [0.62]
Rice	–0.41 (4.72)	–0.15 (2.10)	12.86***	5.67 [0.02]	1.97 [0.74]
Silver	–0.12 (0.88)	–0.06 (1.05)	2.32	NA	6.75 [0.15]
Sugar	–0.43 (4.70)	–0.18 (1.55)	12.29***	2.72 [0.10]	6.74 [0.15]
Timber	–0.18 (2.33)	–0.28 (3.05)	7.37***	0.63 [0.42]	4.24 [0.37]
Tin	–0.08 (1.25)	–0.15 (2.32)	3.48	NA	1.87 [0.76]
Wheat	–0.47 (4.15)	–0.38 (3.66)	12.02***	0.47 [0.49]	2.66 [0.61]
Zinc	–0.43 (4.56)	–0.24 (2.01)	12.42***	1.41 [0.23]	5.77 [0.21]

The numbers in parentheses are the absolute values of the t ratios, and the numbers in square brackets denote probability values. *** denote significance at the 1% level.

Table 4. M-TAR Model

	γ_1	γ_2	$H_0: (\gamma_1 = \gamma_2 = 0)$	$H_0: (\gamma_1 = \gamma_2)$	LB-Q
Beef	–0.25 (3.55)	–0.05 (0.76)	6.60*	3.69 [0.05]	4.39 [0.35]
Cocoa	–0.30 (3.30)	–0.06 (1.09)	5.84*	5.43 [0.02]	0.68 [0.95]
Copper	0.06 (0.52)	–0.16 (2.94)	4.47	NA	3.35 [0.50]
Hides	–0.68 (3.24)	–0.27 (2.86)	8.03***	3.57 [0.06]	5.60 [0.23]
Lamb	–0.19 (2.98)	–0.32 (2.89)	7.52***	1.16 [0.28]	0.34 [0.98]
Lead	–0.10 (1.41)	–0.19 (2.34)	3.75	NA	2.96 [0.56]
Rice	–0.38 (4.22)	–0.18 (2.42)	11.39***	3.19 [0.07]	2.36 [0.67]

Silver	–0.29 (3.78)	–0.01 (0.21)	7.21*	8.79 [0.00]	6.64 [0.15]
Sugar	–0.23 (2.37)	–0.46 (4.33)	12.22***	2.59 [0.11]	5.01 [0.28]
Timber	–0.12 (0.95)	–0.32 (4.53)	11.51***	10.25 [0.00]	3.14 [0.53]
Tin	–0.27 (2.90)	–0.07 (1.33)	5.09	NA	2.89 [0.57]
Wheat	–0.56 (5.14)	–0.29 (2.80)	14.33***	4.21 [0.04]	3.23 [0.52]
Zinc	–0.19 (1.90)	–0.54 (5.07)	14.70***	5.74 [0.02]	4.57 [0.33]

The numbers in parentheses are the absolute values of the t ratios, and the numbers in square brackets denote probability values. ***, and ** denote significance at the 1% and 5% levels respectively

Table 5. LM TAR Model

	γ_1	γ_2	$H_0: (\gamma_1 = \gamma_2 = 0)$	$H_0: (\gamma_1 = \gamma_2)$	LB-Q
Beef	–0.21 (2.49)	–0.12 (1.91)	4.95	NA	2.82 [0.58]
Cocoa	–0.23 (2.79)	–0.06 (1.12)	4.32	NA	0.47 [0.97]
Copper	–0.07 (1.15)	–0.18 (2.29)	3.30	NA	3.38 [0.49]
Hides	–0.47 (3.98)	–0.30 (2.46)	8.43***	1.24 [0.26]	0.61 [0.96]
Lamb	–0.18 (2.22)	–0.24 (3.34)	7.07**	0.36 [0.54]	0.37 [0.98]
Lead	–0.07 (0.98)	–0.21 (2.77)	4.33	1.75 [0.18]	2.64 [0.62]
Rice	–0.41 (4.72)	–0.15 (2.10)	12.86***	5.67 [0.02]	1.97 [0.74]
Silver	–0.09 (1.37)	–0.06 (1.07)	1.47	NA	2.43 [0.65]
Sugar	–0.42 (3.94)	–0.19 (1.50)	8.10**	2.43 [0.12]	1.08 [0.89]
Timber	–0.23 (2.52)	–0.33 (3.37)	7.79**	0.66 [0.41]	0.11 [0.99]
Tin	–0.09 (1.39)	–0.14 (2.17)	3.34	NA	1.86 [0.76]
Wheat	–0.36 (2.87)	–0.32 (2.95)	6.14*	0.07 [0.78]	0.50 [0.97]
Zinc	–0.50 (5.03)	–0.35 (2.92)	15.67***	1.02 [0.31]	1.55 [0.81]

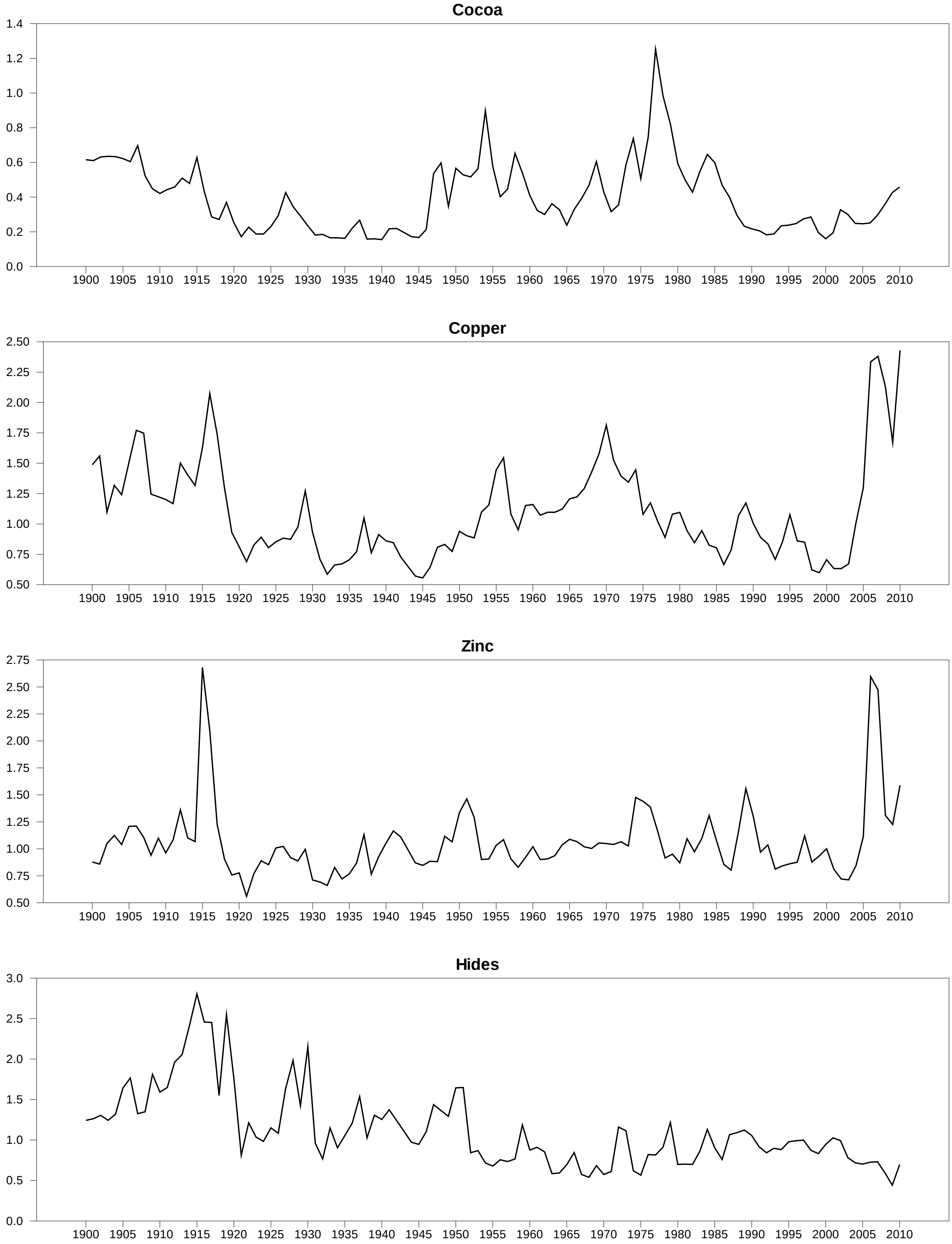
The numbers in parentheses are the absolute values of the t ratios, and the numbers in square brackets denote probability values. ***, **, and * denote significance at the 1%, 5% and 10% levels respectively.

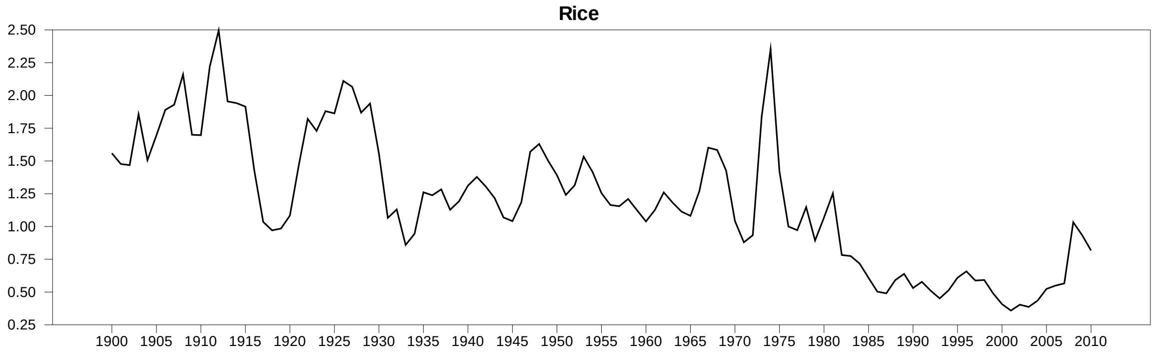
Table 6. LM M-TAR Model

	γ_1	γ_2	$H_0: (\gamma_1 = \gamma_2 = 0)$	$H_0: (\gamma_1 = \gamma_2)$	LB-Q
Beef	–0.25 (3.53)	–0.06 (0.79)	6.57**	3.63 [0.06]	4.11 [0.39]
Cocoa	–0.26 (3.85)	–0.06 (1.20)	4.60*	3.35 [0.07]	0.52 [0.97]
Copper	0.06 (0.52)	–0.16 (2.92)	4.48*	2.94 [0.08]	3.29 [0.51]
Hides	–0.79 (3.71)	–0.33 (3.40)	10.36***	4.58 [0.03]	0.84 [0.93]
Lamb	–0.20 (3.09)	–0.29 (2.66)	7.22***	0.62 [0.43]	0.34 [0.98]
Lead	–0.13 (1.78)	–0.15 (1.91)	3.42	NA	2.64 [0.62]
Rice	–0.38 (4.22)	–0.17 (2.42)	11.40***	3.20 [0.08]	2.36 [0.67]
Silver	–0.29 (3.75)	–0.01 (0.18)	6.31**	8.37 [0.00]	0.44 [0.98]
Sugar	–0.25 (2.21)	–0.40 (3.41)	7.26**	0.95 [0.33]	1.07 [0.89]
Timber	0.07 (0.53)	–0.36 (4.84)	12.81***	9.43 [0.00]	1.03 [0.90]
Tin	–0.35 (3.93)	–0.07 (1.32)	8.08***	7.78 [0.00]	1.42 [0.84]
Wheat	–0.48 (3.98)	–0.23 (2.05)	8.15***	3.65 [0.06]	1.10 [0.89]
Zinc	–0.20 (2.01)	–0.53 (4.94)	14.23***	4.95 [0.03]	4.70 [0.32]

The numbers in parentheses are the absolute values of the t ratios, and the numbers in square brackets denote probability values. ***, **, and * denote significance at the 1%, 5% and 10% levels respectively.

Figure 1. Selected Commodity Prices

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2.00
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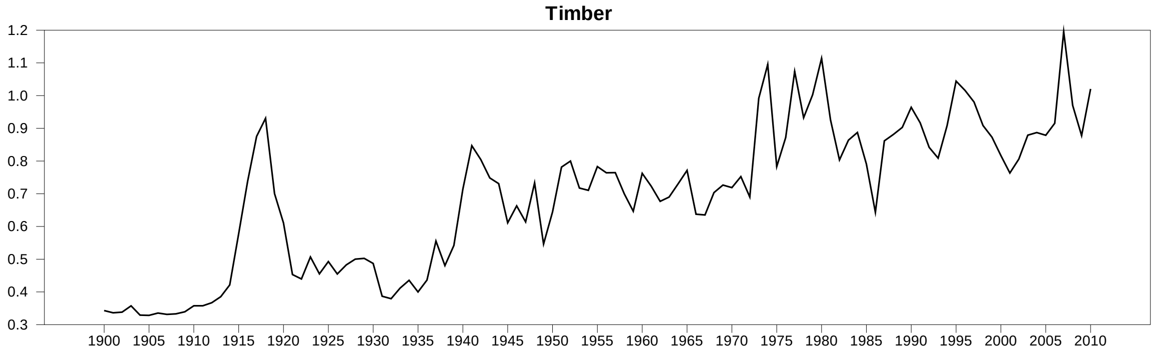
1.50
1.25
1.00
0.75
0.50
0.25
1900 1905 1910 1915 1920 1925 1930 1935 1940 1945 1950 1955 1960 1965 1970 1975 1980 1985 1990 1995 2000 2005 2010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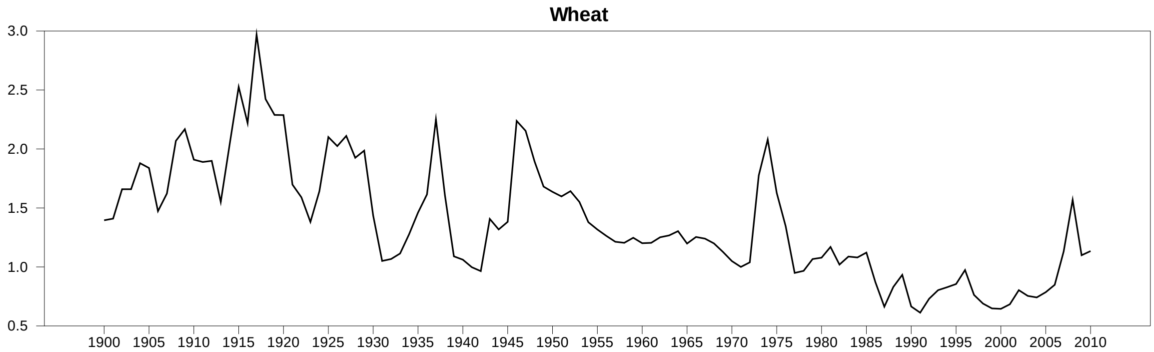
Wheat
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2.0
1.5
1.0
0.5
1900 1905 1910 1915 1920 1925 1930 1935 1940 1945 1950 1955 1960 1965 1970 1975 1980 1985 1990 1995 2000 2005 2010