Coffee price dynamics: An analysis of the retail-international price margin

Atanu Ghoshray [†*], Sushil Mohan [‡]

[†]Economics Group, Newcastle University Business School, Newcastle University, 5 Barrack Road, Newcastle-upon-Tyne, UK.

[‡]Brighton Business School, University of Brighton, Mithras House, Lewes Road, Brighton BN2 4AT, UK.

Abstract

We examine the dynamics of the margin between retail and international coffee prices from 1980 to 2018. We find no significant trend in the margin using a robust procedure for estimating a trend. Further, we establish that any deviations in the margin are transitory for the full sample as well as the periods prior to and after the demise of the ICA, but with asymmetric adjustment. One of the reasons for the observed asymmetry could be market concentration in the coffee supply chain at the coffee roasting level, which allows coffee roasters to keep a higher share of the profits.

JEL: C22; O13; Q13

Keywords: Coffee roasting industry, Coffee supply chain; Market power; Price dynamics; Price margin; Momentum Threshold Autoregression (M-TAR).

*Corresponding author. Email: Atanu.Ghoshray@newcastle.ac.uk; Tel: +44 (0) 191-2081654

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1. Introduction

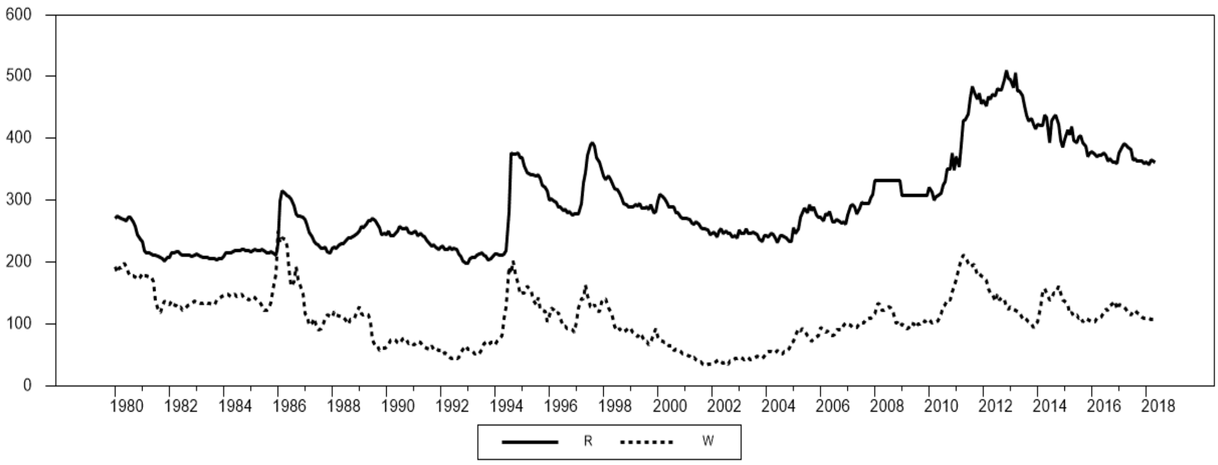
Most agricultural commodities move through a complex processing and distribution supply chain. Coffee is no exception. The focus of this study is on retail and international coffee prices, which appear at the downstream level along the coffee supply chain. We examine the dynamics of the gap between retail and international coffee prices or in other words, the price margin.[1] We address key questions about the dynamics of how this margin is evolving over time; whether this margin is constant or increasing, and if there are any significant deviations from this margin, how the retail and international prices adjust to correct such deviations. This research is of significance as it helps policy makers and practitioners to address concerns being raised on the effects of growth in market concentration (consolidations and increased market power of multinational roasting companies) in the coffee roasting industry, which gives coffee roasters relatively higher market power to capture a larger share in the coffee supply chain.

The coffee value chain relates to all revenues generated by activities carried out along the coffee supply chain. In the coffee supply chain, the producer price is the cash price received at the 'gate' by coffee producers (referred to as producers hereafter). The international price is the price of coffee delivered at the first point of entry in coffee consuming countries.[2] The international price therefore reflects the export price (c.i.f)[3] of coffee from coffee producing countries. The retail price is the national urban US price of roasted ground coffee.[4] For the

[1] Unless specified otherwise, 'coffee' means green (raw or unroasted) beans and coffee prices imply prices of green beans.
[2] The same definition of international coffee prices has been used in Shepherd (2005); Fafchamps and Hill (2008); Gómez et al., (2009); Subervie (2011); Lee and Gomez (2013).
[3] The export price includes the cost of insurance and freight, and any applicable custom duties. This is the price actually paid for physical deliveries of coffee (green bean) on the dock at port of destination.
[4] This definition of retail price has been used in Mehta and Chavas (2008).

purpose of this paper, the retail price is the price of the green coffee equivalent of the roasted ground coffee[5]. The focus of the paper is to analyse the dynamics of the margin between retail and international coffee prices. Figure 1 shows how the retail and international prices of coffee have been evolving over the last four decades. We can see from the graph that the gap between the prices – which is the margin – tends to vary over time.

Figure 1. Retail (R) and international (W) coffee prices (in US cents/lb)



Source: US Bureau of Labor Statistics (2020); ICO (2020).

Prior to the 1990s, unilateral and multilateral interventions in coffee markets were common, the primary objective being price support and price stabilisation for the specific welfare of producers. The interventions were implemented by the International Coffee Agreement (ICA) in 1962, through a quota system to stabilise/support international prices and attenuate competition. The demise of the ICA in 1989 over a disagreement on quotas and the initiation of economic reforms in developing countries in the late 1980s and early 1990s resulted in most countries liberalising their coffee sector. As a result, the world coffee market has become more competitive and subject to market forces. It is widely felt that the end of the ICA regime in

[5] More details about retail price is found in Section 4.1 where we describe the data.

1989 has resulted in a higher proportion of the income generated in the coffee supply chain retained in coffee consuming countries. In an analysis of value of the global coffee chain, the *World Coffee Producers Forum, Colombia July 10-12, 2017* declared that the share reaching coffee producing countries is very low, in contrast to that remaining in the hands of roasting companies in consuming countries. This is because of a shift in market dominance in favour of roasters over agents lower down in the coffee value chain. The paradox is while the coffee chain as a whole is profitable, the vast bulk of the profits are captured by the roasters, with adverse consequences for coffee producing countries dependent on earnings from coffee exports.[6]

However, not all agree that market concentration has contributed to the fall in the share in the value chain reaching coffee producing countries, arguing that it is more of a rhetoric against multinational roasting companies. For example, Bettendorf and Verhoven, 2000; Feuerstein, 2002; Koerner, 2002; and Durevall, (2003, 2017) conclude price transmission in the coffee market rejects the hypothesis that market power determines price transmission. These studies do accept that markets function imperfectly, and that price behaviour may not be an appropriate indicator of market power, given that a highly concentrated sector may be characterised by high price competition. Moreover, several studies argue that since the coffee markets have become more competitive, producers and exporters in the coffee supply chain, have actually increased their returns from more efficient markets, rather than being worse off.[7] A possible reason for the different conclusions of the above mentioned studies is the choice of the coffee

[6] For studies on implications of market concentration for producers see Ponte (2002); Daviron and Ponte (2005); Muradian and Pelupessy (2005); ActionAid South Centre (2008); Hoekman and Martin (2012); Sexton (2013); Igami (2015). For general studies on the low returns to producers and coffee producing countries see Calfat and Flores (2002); McCorriston et al. (2004); Shepherd (2005); Gibbon (2007); Levy (2008); and World Vison (2014).
[7] For studies on higher returns to economic agents in the coffee supply chain in coffee producing countries, see for example Raffaeli (1995); Bohman et al. (1996); Gilbert (1996); McIntire and Varangis (1999); Krivonos (2004); Jarvis (2005); Gemech et al. (2011); Russell et al. (2012).

price from the coffee supply chain employed in those studies. In this paper we focus on the retail and international price of coffee because retail price is a relevant measure of the price charged by final processors of coffee in the retail market; and international price is the export price of coffee from coffee producing countries after adding the cost of insurance and freight. The margin between the two prices includes the transfer costs as well as the profits of roasters. Therefore, the market power of roasters can be expected to be one of the factors affecting the dynamics of this margin.

We analyse the dynamics of the price margin by using a robust econometric model to test whether the margin between retail and international prices has increased over time and whether deviations from the margin are asymmetric. We use robust tests for estimating the trend that allow us to be agnostic of the order of integration of the data; a common problem found in agricultural prices (see Ghoshray 2019). We further aim to determine whether retail and/or international prices respond to correct any deviation in the margin, and whether retail prices adjust at a different rate compared to international prices. Accordingly, we aim to answer two broad research questions:

Question I: While the margin between retail and international prices is likely to fluctuate, are these fluctuations around a constant or a trend? In other words, can the margin be described as a long run constant intertemporal equilibrium value? Or is it gradually increasing over time reflecting consolidation and increased market power over time in the roasting industry?

Question II: If such fluctuations were to occur, do they revert to this long run intertemporal constant equilibrium value or the underlying equilibrium trend? If so, is the adjustment asymmetric, thereby reflecting the presence of market power?

The rest of the paper is organised as follows. Section 2 presents an overview of the market concentration and price dynamics in the coffee market. Section 3 describes the econometric model, Section 4 describes the data and the empirical results and Section 5 concludes.

2. Market power and price dynamics in the coffee market

The general trend in the world coffee market has been the dominance and concentration of market power of multinational roasting companies in the coffee supply chain. In 1998, about two-thirds of the world's coffee was purchased by five multinational companies who controlled nearly two-thirds of the world market share for roasted and instant coffees: Philip Morris (Kraft Jacob Suchard), Nestle, Procter and Gamble (P&G), Sara Lee and Tchibo. Table 1 shows the companies' world market retail share for roasted and instant coffee in 1998 and 2014. The combined market share in 1998 of Nestle and Philip Morris was 49 percent. The trend of market dominance and consolidation of coffee market continued until 2002, with Nestle and Kraft Jacob Suchard further consolidating their share of the world market for roasted and instant coffee (Brown and Gibson, 2006; ActionAid and South Centre, 2008).

Thereafter (post-2002), the roasting market has witnessed a gradual trend of embracing greater diversity, evident from the emergence of new players and a gradual fall in market share of roasters compared to 1998 (see Table 1). The bigger size of the coffee sector and larger geographical dispersion of consumption has allowed for the emergence of a large number of small roasters. In addition, the US has seen the emergence of small specialty coffee roasters capturing a higher market share; this trend can also be seen in other European countries. Despite the gradual trend of market diversification, Nestle and Jacobs Douwe Egberts remain prominent players in the market, with 38 percent combined market share of global roast and instant coffee in 2014, though lower than their share of 49 percent in 1998 (Statistica, 2016).[8] Since 2014, there have been reconsolidation efforts by Jacobs Douwe Egberts, going on a buying spree in its quest to challenge the long standing market dominance of Nestle. Despite efforts of Nestle and Jacobs Douwe Egberts to hold on to their dominant position, there are signs of market diversification, albeit in a very gradual manner (Grabs, 2017).

Table 1. Share of global coffee market by roasters (1998 and 2014)

	1998			**2014**	
Roaster	Market share (%)	Cumulative share (%)	Roaster	Market share (%)	Cumulative share (%)
Philip Morris	25	25	Jacob Douwe Egberts	16	16
Nestle	24	49	Nestle	22	38
Sara Lee	7	56	Green Mountain	5	43
P&G	7	63	Strauss	3	46
Tchibo	6	69	Tchibo	2	48
Others	31	100	Others	52	100

Source: Ponte (2002); Statistica (2016)

The overview shows high levels of dominance and concentration of market power in multinational roasting companies until around 2002 followed by a gradual dilution of this

[8] Jacobs Douwe Egberts is a Dutch privately owned company that owns numerous beverage brands. It was formed in 2012 following the merger of Philip Morris (the coffee division of Mondelez International) with Douwe Egberts.

dominance due to some diversity and reorganisation in coffee roasting over the years. Despite the dilution, looking at Table 1, one can say that there still is continued concentration of market power in multinational roasting companies. Given the market power in coffee supply chain, it is not surprising that large roasters are regularly alleged for leveraging their power to capture high share of the rents that accrue in the coffee value chain.

Market concentration can alter the marketing systems and can have an impact on the dynamics of price transmission in the coffee value chain. Market concentration could potentially weaken the ability of coffee producing countries to influence international prices, while increasing the ability of coffee roasters in coffee importing countries to influence international prices and the extent to which changes in international prices are passed on to retail prices (and vice-versa), resulting in price transmission asymmetries in the coffee supply chain. As a case in point, there is evidence of price transmission asymmetries in supply chains for agricultural commodities. Various empirical studies focusing on food products find that increases in input (factor) prices are often transmitted more quickly to retail prices than decreases in these prices (Serra and Goodwin, 2003; Meyer and von Cramon-Taubadel, 2004; Lass, 2005; McLaughlin, 2006). The literature identifies market structure and the presence of non-competitive behaviour (i.e. market power) as the main cause for such asymmetry in price transmission (Ward, 1982; Bacon, 1991; Borenstein et al., 1997; Peltzman, 2000; Nakamura and Zerom, 2010). The literature identifies another explanation of imperfect transmission in the context of oligopolistic and monopsony markets. The risk of provoking a price war may make oligopolistic firms reluctant to lower their prices in response to fall in input prices; therefore price adjustment in response to the fall might be sluggish or take place only after time lags. In oligopolistic markets with unspoken collusion, oligopolistic firms will use price changes to signal the unspoken agreement (Balke et al., 1998; Brown and Yucel, 2000). When input prices rise, each oligopolistic firm will quickly adjust prices upwards to signal that collusion will be maintained, whereas the response of the oligopolistic firms will be slower when it comes to adjusting prices downwards when input prices fall to avoid undermining a tacit agreement. Further explanations by Kinnucan and Forker (1987) describe how government intervention can lead to asymmetric price transmission. Processors of agricultural commodities may believe that a reduction in input price may be temporary because it will trigger government intervention through support prices. In this context, processors will not react to a reduction in input prices, but they will quickly respond to increases in input prices because they will believe it is more likely to be long-lived.

The upshot from this discussion leads one to observe that asymmetric price adjustment behaviour can be relevant in the context of coffee. Where market power is in the hands of roasters, this would mean that increases in international prices would trigger a prompt increase in retail prices to ensure no reduction in roasters margins. However, decreases in international prices may not elicit the same response (i.e. prompt decrease in retail prices) as roasters are in a position to exploit their market power by keeping prices above the competitive level. There have been studies that analyse the impact of ICA termination on price transmission at various levels in the coffee supply chain (Bohman et al., 1996; Buccola and McCandlish, 1999; Krivonos, 2004; Mehta and Chavas, 2008; Fafchamps and Vargas, 2008; Gomez et al., 2009; Lee and Gómez, 2013). Shepherd (2005) examined the impact of the end of ICA on price transmission from producer to international prices and from international to retail prices employing a vector autoregression (VAR) model. The results suggested that the ICA termination did not improve price transmission because of market power exerted by coffee roasters. Moreover, asymmetries in price transmission at all levels of the supply chain were identified, particularly during the post-ICA period. A set of studies (for example Feuerstein, 2002; Shepherd, 2005) are concerned with short term price transmission issues and seek evidence regarding the allegations of growing market power in the coffee roasting sector through the 1990s. They find that price transmission to the retail sector is asymmetric, with retail prices more responsive to increases than decreases.

Mehta and Chavas (2008) studied the price effects of the ICA termination and found that the short-run retail price response was greater for increases than for decreases in international prices during the post-ICA period. Lee and Gómez (2013) examine price transmission from international to retail coffee prices and find evidence of short run asymmetries with differences among importing countries because of dissimilarities in market structures across countries. For example, in the US, retail prices rise faster than they fall in response to changes in international prices while in Germany and France retail prices respond faster when international prices are falling. Leibtag et al., (2007) use price data over the period 1997 to 2004 to study the path of raw material cost pass-through in the US coffee industry to gain insights on how changes in the input cost of coffee affect consumer (supermarket) prices of coffee. They do not find robust evidence that coffee prices respond more to increases than to decreases in coffee input costs. Nakamura and Zerom (2010) point out that firms have to pay a 'menu cost' to adjust the consumer prices which could result in price rigidity behaviour in terms of delayed and sluggish response to changes in input prices. However, the role of 'menu cost' is more valid in the short-run and for relatively small changes in input prices, in comparison to the long run and for substantial changes in input prices where the role of menu costs is negligible. Subervie (2011) apply threshold cointegration to analyse the dynamics of international coffee price transmission to producers over the pre and post-ICA period. They find asymmetric price adjustments in the post-ICA period (large decreases in world prices being transmitted relatively quickly to producers as against increases) that can be seen as expressions of an unfavourable pricing influence over the post-ICA period. This could be because of emergence of new market structure over the post-ICA period, meaning that termination of the ICA may have failed to create competitive market structures in some cases.

In general, we cannot assume that high levels of roaster buyer power will necessarily lead to high international and retail price margin or even high profit margins for roasters. This is because even if the higher buyer power allows roasters to exert pressure on keeping lower international prices, in a free market it can be expected that roasters would compete with each other from the seller-side of the roasted coffee market. Furthermore, it is reasonable to expect that the margin between international and retail prices is likely to fluctuate, given the volatile nature of coffee prices in general.[9] The following econometric model is designed to capture these dynamic properties of the coffee price margin.

3. Econometric model

In this section, we lay out the framework for conducting the empirical test of coffee price margin adjustment. We particularly focus on the possibility that the cyclical adjustment of the coffee price margin around its long term mean or trend might be asymmetric in the light of the arguments made in the previous sections. The speed or momentum at which price margins fluctuate around the long run mean or trend may differ, depending on whether the prices are increasing or decreasing relative to the mean.

We test whether the price margin has been increasing over time as has been suggested in some past studies (e.g., Talbot 1997, Calfat and Flores 2002; Ponte 2002). This is a testable hypothesis using the following trend function:

[9] The high volatility of coffee prices is due to the susceptibility of output to frosts, disease, and droughts, magnified by the inelasticity of demand with respect to prices and income, and price inelasticity of supply (Mehta and Chavas, 2008).

$$P_t^R - P_t^W = \alpha + \beta t + u_t; \qquad u_t = \sum_{i=1}^{k} \varsigma_i u_{t-i} + e_t \tag{1}$$

where P_t^R and P_t^W denote the retail and international prices for coffee respectively, α is an arbitrary constant and $P_t^R - P_t^W$ denotes the margin between the prices; the term t denotes the time trend, u_t measures the deviation from trend, which may be serially correlated and is described in this case as an AR(k) process. The lag length k is selected using the Modified Akaike Information Criterion (MAIC) following Ng and Perron (2001) with k allowing to be in the range $Int\left[0, 12(T/100)^{1/4}\right]$, where T denotes the sample size, and Int denotes the integer value[10]. The parameter β denotes whether the margin is changing over time. For example, if $\beta > 0$, then the margin will increase over time, as opposed to decreasing with time if $\beta < 0$; and there is no change in the margin if $\beta = 0$. However, a problem with estimating the trend function is the unknown underlying order of integration of the price margin, $P_t^R - P_t^W$. If the order of integration is non-stationary I(1), then the standard method of least squares to estimate the trend will suffer from severe size distortions (Perron 1988). Alternatively, if the data is stationary I(0) but modelled as I(1), the trend estimation will be inefficient and lacking power (see Perron and Yabu, 2009a).

To obviate this problem of possible non-stationarity of the price margin, we employ the robust procedure of trend estimation due to Perron and Yabu (2009a). This method allows one to be agnostic to the underlying order of integration of the data. A quasi–feasible generalised least squares (q-FGLS) procedure is applied to obtain the estimate of β the trend parameter, denoted $\hat{\beta}$, and construct the robust q-FGLS t–statistic for the unbiased and median unbiased estimate,

[10] This is based on a rule suggested by Schwert (1989).

that is, $t_{\beta}^{RQF}(UB)$ and $t_{\beta}^{RQF}(MU)$ respectively (see Perron and Yabu 2009a). For completeness, we calculate both the unbiased and the median unbiased estimates.

To estimate the dynamics of the price margin, we employ the momentum threshold autoregressive (MTAR hereafter), which is a procedure proposed by Enders and Granger (1998). This procedure is particularly attractive as it neatly fits in with the estimation and hypothesis testing as in this case of coffee price adjustment. As a prelude to determining whether adjustments of the margin are asymmetric or not, we need to establish that deviation of the margin are transitory in nature. As Enders and Granger (1998) stipulate, that rejection of the unit root null in the margin, would allow us to test for asymmetric adjustment. This is reinforced by their view that testing for unit roots are mis-specified if the underlying adjustment is asymmetric.

Accordingly, in this study, we use the MTAR method due to Enders and Granger (1998).
The MTAR model along with the indicator function I_t and estimated threshold τ can be written as:

$$\Delta Z_t = I_t \gamma_1 Z_{t-1} + (1 - I_t)\gamma_2 Z_{t-1} + \sum_{i=1}^{p} \phi_i \Delta Z_{t-i} + \omega_t \qquad (2)$$

$$I_t = \begin{cases} 1 \text{ if } \Delta Z_{t-1} \geq \tau \\ 0 \text{ if } \Delta Z_{t-1} < \tau \end{cases} \qquad (3)$$

where Z_t is the detrended margin. The MTAR model allows the margin to exhibit more momentum in one direction than the other. If we find, $|\gamma_1| < |\gamma_2|$, that implies we find the MTAR model exhibits little adjustment for $\Delta Z_{t-1} > \tau$ but substantial decay for $\Delta Z_{t-1} < \tau$. In other words, during the phase where $\Delta P_t^R > \Delta P_t^W$ creates a deviation in the margin, this deviation tends to be corrected relatively slowly, in comparison to the phase where a deviation arising from $\Delta P_t^R < \Delta P_t^W$ which is corrected at a relatively faster rate. Alternatively, if the asymmetry were to be contrary to that proposed, then we would find, $|\gamma_1| > |\gamma_2|$; then the increasing margin due to $\Delta P_t^R > \Delta P_t^W$ tends to dissipate at a faster rate compared to the phase where $\Delta P_t^R < \Delta P_t^W$.

We use the methodology proposed by Chan (1993) to estimate the threshold denoted τ[11]. The estimated residual series are sorted in ascending order, that is, $\Delta Z_1 < \Delta Z_2 < \cdots < \Delta Z_T$ where T denotes the number of usable observations. The largest and smallest 15 percent of the ΔZ_t series are eliminated and each of the remaining 70 percent of the values were considered as possible thresholds. For each of the possible thresholds the equation was estimated using (2) and (3). The estimated threshold yielding the lowest residual sum of squares was deemed to be the appropriate estimate of the threshold.

The null hypothesis of a unit root in the coffee price margin is given by the following testable hypothesis, that is, $H_0: (\gamma_1 = \gamma_2 = 0)$ which is obtained from estimating equation (2) and comparing to the critical values computed by Enders and Granger (1998), against the alternative H_A: $\gamma_1 < 0$ and/or $\gamma_2 < 0$. Note that under the null hypothesis, the margin is symmetric and there is a unit root in both regimes; while under the alternative hypothesis there

[11] Enders (2001) points out that the demean series will be a biased estimator of the threshold τ as γ_1 and γ_2 differ which motivates the estimation of the consistent threshold which is super-consistent using the grid search approach by Chan (1993). Note, while calculating the critical value Enders (2001) contains an error which is pointed out by Cook and Manning (2003) where they show that the power of consistent threshold MTAR model is found to be higher than all forms of plausible alternatives using the newly designed critical values which we consider in this paper. Actually Enders (2001) concedes in his paper that the results he obtains for the MTAR are counter-intuitive as the consistent threshold M-TAR model employs a consistent estimator of the threshold and therefore should have increased power.

is a stationary process in at least one regime (see Enders and Granger 1998). If we can reject the null hypothesis, it is possible to test for asymmetric adjustment, that is, $H_0: (\gamma_1 = \gamma_2)$, using the F statistic. If the test statistic is greater than the critical value from the F table, we can reject that the adjustment to any deviation is symmetric, enabling us to conclude that there is asymmetric adjustment. Diagnostic checking of the residuals is undertaken using the Ljung-Box Q tests to ascertain whether the ω_t series in (2) is a white noise process; and to ensure the residuals are white noise, the right hand side of (2) is augmented by lagged variables given by $\sum_{i=1}^{p} \phi_i \Delta Z_{t-i}$. The lag length p in (2) is determined by the 'General to Specific' criterion.

The positive finding of stationarity with MTAR adjustment for the price margin allows us to estimate an asymmetric error correction model (AECM) as follows:

$$\Delta P_t^R = \psi_1 V_{t-1}^+ + \psi_2 V_{t-1}^- + \sum_{i=1}^{p} \theta_i \Delta P_{t-i}^R + \sum_{i=1}^{p} \xi_i \Delta P_{t-i}^W + \upsilon_t \tag{4}$$

$$\Delta P_t^W = \lambda_1 V_{t-1}^+ + \lambda_2 V_{t-1}^- + \sum_{i=1}^{p} \vartheta_i \Delta P_{t-i}^R + \sum_{i=1}^{p} \mu_i \Delta P_{t-i}^W + \eta_t \tag{5}$$

where $V_{t-1}^+ = I_t(P_t^R - P_t^W + \tau)$ and $V_{t-1}^- = I_t(P_t^R - P_t^W + \tau)$ where I_t is the indicator function as shown in (3); υ_t and η_t are white noise error terms. The parameters ψ_1 and ψ_2 as shown in (4) show the speed of adjustment coefficients of retail prices in response to a positive and negative deviation respectively, from the long run equilibrium margin. The lagged differenced prices that augment equation (4) are to ensure that the error term υ_t is white noise. Similarly, the parameters λ_1 and in (5) show the speed of adjustment coefficients of international prices in response to a positive and negative deviation respectively from the long run equilibrium margin. The AECM allows us to test for Granger causality thereby allowing us to determine the predictive power of the change in retail (international) prices on international (retail) prices. For example, in (4) we can test the null hypothesis $H_0: (\xi_1 = \xi_2 = \cdots = \xi_p = 0)$. Under the null we are testing that international prices do not Granger cause retail prices. Rejecting the null, implies that international prices Granger cause retail prices indicating predictive ability. Similarly, we can set up the null hypothesis H_0: $(\vartheta_1 = \vartheta_2 = \cdots = \vartheta_p = 0)$ to test whether retail prices Granger cause international prices. Rejecting the null would imply causality exists.

Finally, we carry out innovation accounting by tracing out the responses to exogenous shocks to the MTAR model. Following Coakley et al. (2001), we use the recursive nature of the model to create a history $h_{t-1} = \{Z_{t-1}, Z_{t-2}, \dots,\}$ and time t shock ω_t, which serves as an initial condition, and n randomly selected shocks $V_t = \{v_{t+1}, v_{t+2}, \dots, v_{t+n}\}$. Using the parameter estimates of the MTAR model, we generate k sets of forecasts for the shocked model $\{Z_{t+i}(h_{t-1}, \omega_t)\}_{i=0}^{n}$ and k sets of baseline forecasts $\{Z_{t+i}(h_{t-1})\}_{i=0}^{n}$ using the same history and random future shocks in all the forecasts (see Coakley et al., 2001). The generalised impulse response function can be defined as:

$$IRF(i, h_{t-1}, \omega_t) = E[Z_{t+i}|h_{t-1}, \omega_t] - E[Z_{t+i}|h_{t-1}], \qquad i = 0,1,2, \dots, n$$

The difference of the two forecasts is averaged over the k replications and repeated 100 times for different combinations of history and shocks.

4. Empirical Analysis

This section is structured into three sub-sections, we first describe the data, followed by the robust estimation of the trend of the margin and then the estimation of the asymmetric price adjustment of the margin.

4.1 Data

All price data are in nominal terms and measure the monthly average price in US cents per pound for the period January 1980 to May 2018 (see Figure 1). As a measure of retail price we use the monthly averages of the national urban US price of roasted ground coffee obtained from the US Bureau of Labor Statistics, 2020.[12] The data series for the period 1980-2018 has missing values for the monthly averages of September and October 2007, and of January 2008 to November 2009. However, the annual averages for the year 2007, 2008 and 2009 are available from the International Coffee Organization (ICO); we have interpolated the missing values using annual averages for the year. The retail price reflects only the price of roasted ground coffee; prices of whole bean gourmet coffee and coffee drinks are not reflected in this price (see Mehta and Chavas, 2008). We focus on roasted ground coffee since green beans is the input in the production process and the product has been homogenous over the years. By input we mean raw material input; the conversion process to roasted ground coffee does include other inputs such as labour and machinery. However, the other input costs are much lower for roasted ground coffee compared to instant coffee or other forms of coffee, which explains our choice of retail price to reflect the price of roasted ground coffee. In accordance with internationally accepted practices, the ICO has stipulated conversion factors to convert different types of coffee to green bean equivalent (GBE); for converting roasted coffee to GBE requires multiplying the weight of roasted coffee by 1.19. Therefore, for the retail price to reflect the price of same quantity of coffee we convert the retail price of ground coffee to GBE price by dividing the retail price by 1.19.[13]

The bulk of the coffees used in roasted ground coffee blends are Brazilian Natural grade Arabica and Robusta (both Asian and Brazilian Robusta); rough estimates place the share of Arabica and Robusta coffee in the roasted ground coffee blends at 60 and 40 percent

[12] Data can be accessed at US Bureau of Labor Statistics, Consumer Price Index, Average Price data, coffee, 100%, ground roast, all sizes, per lb; Series Id: APU0000717311.

[13] This adjustment is made simply for comparison purposes. The price only changes in scalar terms and is not affected when conducting the econometric analysis of the price dynamics.

respectively. For international price to be comparable with the retail price of ground coffee, we use a weighted average price of Arabica (60 percent) and Robusta (40 percent) coffee. As a measure of international price, we use the weighted average of the ICO monthly Indicator Price for Brazilian Natural Arabica (60 percent weight) and Robusta (40 percent weight) for the period 1980 to 2018. The prices are available on a monthly average basis from the ICO database and are calculated by weighting the ex-dock prices on the international markets in New York, Bremen/Hamburg and Le Havre/Marseilles markets (ICO, 2020).[14]

4.2 Estimating the trend of the price margin

In the first instance we test whether the margin between retail and international prices has been increasing over time. To this end, we apply the robust procedure due to Perron and Yabu (2009a) to determine whether the trend is significant or not. This amounts to testing the null hypothesis that $H_0: (\beta = 0)$ in equation (1). As discussed before, this test allows us to be agnostic to the underlying order of integration of the data. The results of the trend function are given by the following regression:

$$P_t^R - P_t^W = \underset{(8.51)}{90.27} + \underset{(0.397)}{0.355}t$$

The trend estimate $\hat{\beta} = 0.355$ is positive in magnitude but statistically not different from zero as shown by the standard error 0.397 in parentheses. The associated t-statistic for the trend estimate is calculated to be 0.894, which falls below the conventional levels of significance. We repeat the robust trend estimation procedure using the median unbiased statistic and the results are identical.[15] We can conclude from this result that the trend in the coffee price margin is not significant. While the sign is positive, there is too much variability around this trend to

[14] Data can be accessed at ICO, composite & group indicator prices - monthly averages; data prior to 1990 is available on request from the ICO.
[15] The results are not reported here but are available on request.

concur that it is significantly positive. This result is not surprising given the amount of variation we observe in the graph of the margin in Figure 2. Besides international prices can lead to higher volatility in retail prices with a lag (see Mehta and Chavas, 2008), that contributes to the large variation of the margin over time. Using the robust trend estimate of Perron and Yabu (2009a) we calculate the 90% confidence interval of the trend estimate to be (–0.299, 1.01), and the 95% confidence interval is (–0.423, 1.13). Both confidence intervals cannot exclude zero thereby rendering the trend to be insignificant.

Figure 2. Price margin of retail-international coffee prices.

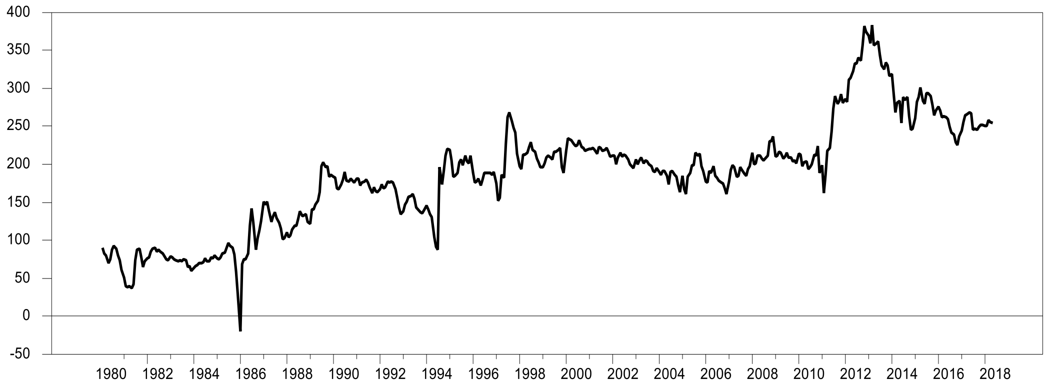


Figure 2 tends to reflect that there may be a slight positive trend but the huge amount of variability around the trend renders the trend estimate to be statistically insignificant. Our robust procedure is in line with the finding by Mehta and Chavas (2008) where they too find an insignificant trend estimate of the retail-international price margin using a 'delta method'. This result departs from those studies that concluded an increasing margin such as Talbot (1997), Calfat and Flores (2002) and Ponte (2002). However, these latter studies do not use robust econometric methods.

However, one may argue that there could be one or more structural breaks in the trend thereby causing the estimate to be insignificant. To address this, we make use of robust structural break tests. First, we apply the Perron and Yabu (2009b) quasi-feasible Wald (W-QF) test for a single structural break in the trend. Given the length of the data sample, we also apply the sequential test for multiple structural breaks using the robust Supremum F test (sup-F) procedure of Sobriera and Nunes (2016). Both tests are robust, allowing us to be agnostic of the underlying order of integration of the data series. The results of the structural break tests are given in Table 2 below.

Table 2. Robust tests for structural breaks

Single Break (W-QF)				Multiple Breaks (sup-F)			
Null	W-QF	5% c.v.	10% c.v.	Null	Sup-F	5% c.v.	10% c.v.
$(0\|1)$	–0.29	1.67	1.13	$(0\|1)$	3.04	9.41	7.68
				$(0\|2)$	2.99	8.44	7.17
				$(0\|3)$	2.24	7.17	6.40

Notes: The single break test due to Perron and Yabu (2009b) is given by the quasi-Feasible Wald (W-QF) test. The test is carried out using the null hypothesis of no break against a single break or $(0|1)$. The critical values (c.v.) at the 5% and 10% significance levels are reported alongside the test statistics. The multiple breaks due to Sobriera and Nunes (2016) are given in the second column of results using the sup-F test. These are sequential tests that test for no break against one (0|1), two (0|2) and three (0|3) breaks separately.

Using the single structural break test due to Perron and Yabu (2009b) we find the estimated W-QF test statistic to be below the critical value at standard conventional levels. We therefore cannot reject the null hypothesis of no break against the alternative of a single break [that is, $(0|1)$] in the trend. Based on the robust single break test, our conclusion is that there is no evidence of breaking trends. Further, as a confirmatory test, we apply the robust multiple break test procedure due to Sobriera and Nunes (2016) allowing for up to 3 breaks based on the sample size. In each of the cases we cannot reject the null hypothesis of no break against one break [that is,$(0|1)$], followed by no break against two breaks [that is, $(0|2)$], and finally, no break against three breaks [$(0|3)$]. The upshot is that there is no evidence of any breaking trends, therefore trend estimation is secular, but we find the estimate to be insignificant. Hereafter, the econometric analysis of the dynamics of the margin excludes the presence of the trend.

4.3 Estimating the dynamics of the price margin

We test for the dynamic behaviour of the price margin using the MTAR model. The delay parameter for the models is set as $d = 1$. The MTAR model is estimated using equations (3) and (4) and the non-zero threshold is calculated using the method by Chan (1993).[16] The results are presented in Table 3:

Table 3. Results of the MTAR model for the full sample

Null hypothesis	Parameter/Test	MTAR
Positive deviation	γ_1	–0.006 (–0.77)
Negative deviation	γ_2	–0.289 (–5.78)***
Non-stationarity	$H: (\gamma_1 = \gamma_2 = 0)$	16.97[a]
No asymmetry	$H: (\gamma_1 = \gamma_2)$	31.28 [0.00]**
No serial correlation	Ljung-Box Q	0.43 [0.97]

[16] We employ a non-parametric approach due to Tsay (1989) to identify the existence of threshold effects in the AR component of the model. This test returns an F statistic to test the null hypothesis of no changes in the parameter estimates of the AR representation of the data. We find that the null can be rejected suggesting the existence of a threshold, prompting us to use a threshold model instead of a simple AR. For brevity, we do not report the details but the results are available on request.

Threshold estimate	–27.05

Notes: [a] denote rejection of the null at the 1% significance level. The numbers in parentheses denote t-statistics and the numbers in square brackets denote p-values. The critical values of the MTAR test at the 1% significance level is 6.99. The results in Table 3 show that we can reject the null at the 1% significance level.

In the first column of Table 3 we set out the hypothesis of interest. First, we report the estimate of the γ_1 parameter which provides the rate of adjustment when the change in retail prices is greater than the change in international prices thereby creating a 'positive deviation' in the margin. The opposite deviation, labelled as a 'negative deviation' in the margin, returns the parameter estimate γ_2, which provides the rate of adjustment when the retail prices are changing at a slower rate than international prices making the deviation shrink, which we will call 'negative'. The null hypothesis of non-stationarity is given by $H_0: (\gamma_1 = \gamma_2 = 0)$, which simply states that the price margin is a random walk. Subject to rejecting the null of non-stationarity, the null hypothesis of symmetry is given by $H_0: (\gamma_1 = \gamma_2)$, which states that there is no significant difference for the rates of adjustment given by γ_1 and γ_2. This is followed by a diagnostic test to determine whether there is no serial correlation in the residuals of the regression equations given by (3).

We can reject the null of non-stationarity, that is, $H_0: (\gamma_1 = \gamma_2 = 0)$ (given by the test statistic equal to 16.97, significant at the 1% level) in the margin. Note that under the alternative hypothesis (that is, H_A: $\gamma_1 < 0$ and/or $\gamma_2 < 0$) there is a stationary process in at least one regime (see Enders and Granger 1998); this is found to be when the there is a negative deviation. In particular, we find the negative discrepancy is corrected at the rate of 28.9 percent every month based on the parameter estimate of $\gamma_2 = -0.289$. However, for a positive discrepancy we find no signs of adjustment. The parameter estimate of $\gamma_1 = -0.006$ is found to be statistically insignificant. The null hypothesis of symmetry, given by $H_0: (\gamma_1 = \gamma_2)$ is rejected as shown by the p-value of 0.004, thereby concluding that there is asymmetric adjustment. To sum up, we can conclude from the MTAR model, that a positive deviation in the margin tends to persist, whereas a negative deviation in the margin is rapidly corrected, thereby underscoring the case for asymmetric adjustment.

Though there is no evidence of a single or multiple structural breaks in the margin, we conduct further estimations to determine whether the dynamics of price adjustment are the same prior and after the elimination of the export quota system in 1989 following the collapse of the ICA. Accordingly, we divide the full sample into two regimes: Regime I, from January 1980 to August 1989; and Regime II from September 1989 to May 2018. We apply the MTAR model to both regimes to check whether the underlying price dynamics remain the same or change. The results are presented in Table 4.

Table 4. Results of the MTAR model for the two regimes.

Null hypothesis	Parameter/Test	Regime I	Regime II
Positive deviation	γ_1	0.003 (0.16)	0.002 (0.34)
Negative deviation	γ_2	–0.89 (–7.73)[a]	–0.11 (–4.44)[a]
Non-stationarity	$H: (\gamma_1 = \gamma_2 = 0)$	30.26[a]	10.15[a]
No asymmetry	$H: (\gamma_1 = \gamma_2)$	59.58 [0.00]	15.88 [0.00]
No serial correlation	Ljung-Box Q	7.29 [0.12]	0.40 [0.98]
Threshold estimate		–22.35	–5.91

Notes: [a] denotes rejection of the null at the 1% significance levels respectively. The numbers in parentheses denote t-statistics and the numbers in square brackets denote p-values. The critical values of the MTAR test at the 1% significance level is 6.99. The results in Table 4 show that we can reject the null at the 1% significance level.

In some respects the empirical results lead to the same broad conclusion for both regimes. For both regimes, we can reject the null of non-stationarity, that is, H_0: $(\gamma_1 = \gamma_2 = 0)$ in the margin (which is given by the test statistic equal to 30.26, significant at the 1% level in Regime I, and equal to 10.15, significant at the 1% level in Regime II). Note that under the alternative hypothesis (that is, H_A: $\gamma_1 < 0$ and/or $\gamma_2 < 0$) there is a stationary process in at least one regime, which can be expected in the MTAR model (see Enders and Granger 1998). In both cases the adjustment takes place for the negative deviation given by the parameter estimate γ_2. In Regime I, any deviation in the steady state margin, is corrected at the rate 89 percent every month and in Regime II at the rate of 11 percent every month. In this case, the discrepancy occurs when the margin shows a negative discrepancy. These conclusions on adjustment of price margins for each separate regime are no different from that of the full sample. When considering the parameter estimate γ_1, we find that the rates of adjustment are statistically insignificant. That is, the parameter estimates of 0.003 in Regime I and 0.002 in Regime II are reported to be statistically insignificant, so that a positive deviation in the steady state margin, is allowed to persist. We find that prior to the elimination of export quotas, adjustment to correct any negative discrepancy is relatively fast (i.e., 89 percent of the deviation is corrected every month in Regime I), in comparison to the period when quotas were eliminated (where 11 percent of the deviation is corrected every month in Regime II). We can confirm that there is asymmetric adjustment, by rejecting the null hypothesis of symmetry, given by H_0: $(\gamma_1 = \gamma_2)$ as shown by the p-value of 0.06 in Regime I and 0.09 in Regime II, showing rejection at the 10% significance level. Apart from the differences in the speed of adjustment, the dynamics for the full sample reflects that of the sub-samples corresponding to the regimes prior to and after the collapse of the ICA.

Since we find the margin to be stationary with MTAR adjustment, we proceed to estimate an asymmetric error correction model (AECM) as described by equation (5) and (6) in the previous section. The results of the AECM are given in Table 5 below. The results include the full sample and the two regimes.

Table 5. Results of the ECM for the full sample and the two regimes

	Full sample		Regime I		Regime II	
	ΔP_t^R	ΔP_t^W	ΔP_t^R	ΔP_t^W	ΔP_t^R	ΔP_t^W
Positive deviation	–0.010	–0.003	0.010	–0.017	0.011	–0.021
	(–1.73)	(–0.60)	(0.94)	(–0.62)	(0.80)	(–2.01)
Negative deviation	–0.087	0.166	–0.35	0.47	–0.089	0.002
	(–2.69)	(4.69)	(–8.60)	(4.44)	(–4.22)	(0.11)
Granger Causality	9.15	2.67	3.73	2.19	5.81	2.31
	[0.00]	[0.00]	[0.00]	[0.02]	[0.00]	[0.00]
Ljung Box Q	0.29	0.51	4.32	1.40	0.22	0.196
	[0.99]	[0.97]	[0.11]	[0.84]	[0.99]	[0.99]

The numbers in parentheses denote the t-statistics, while the numbers in square brackets are probability values.

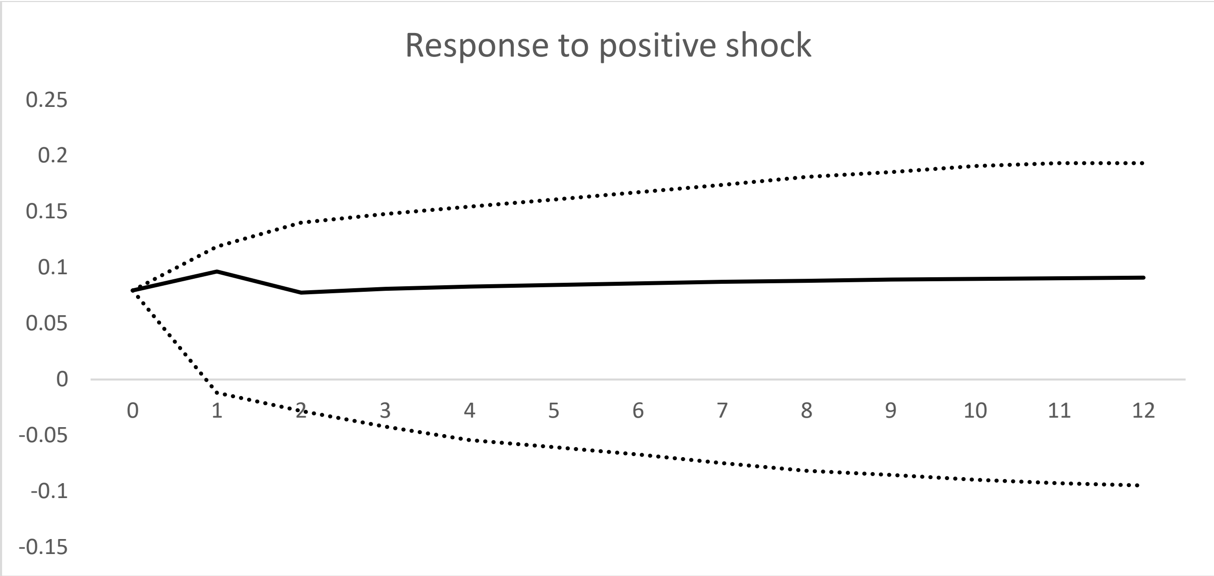
We first consider the full sample results. During the phase, when there is a positive deviation in the margin, we find that such a deviation is corrected by the retail prices, albeit at a very slow rate of 1% every month. International prices do not adjust to correct this deviation, as the speed of adjustment parameter (–0.003) is statistically insignificant. However, if there is negative discrepancy in the margin, then we find that such a deviation is corrected by both retail and international prices. International prices adjust at the rate of 16.6 percent every month. In comparison, retail prices adjust at the rate of 8.7 percent every month. These results allow us to determine how retail and international prices adjust, and we find the former adjusts at a sluggish rate to the latter, and both only correct any deviation during the phase when the price margin shows a negative discrepancy. In the short run we find bidirectional causality, where a feedback effect is found to exist between international and retail prices. This implies changes in retail prices lead to a change in international prices and vice versa in the short run. The model is free from serial correlation as shown by the Ljung-Box Q test statistics.

When considering Regime I, we find the dynamics to be different. When the price margin shows a positive discrepancy, neither the retail nor international prices adjust to correct the deviation. When a negative discrepancy occurs, we find that such a deviation is corrected by both prices. We find the retail prices adjusts at a relatively slower rate (35 percent every month) compared to the international prices (47 percent) and this adjustment occurs during the phase when the price margin shows a negative discrepancy. In this period prior to the collapse of the ICA, we find that the absolute value of the speed of adjustment to be faster when compared to the speed of adjustment coefficients for the full sample. The short run shows that the Granger causality is bidirectional, similar to the dynamics of the full sample. No problems with serial correlation are reported. In the case of Regime II, the correction to any deviation is distinct to the full sample results and Regime I. If there is a negative deviation in the margin, then only the retail prices adjust at the rate of 8.9 percent every month to close the deviation. On the contrary, when there is a positive deviation in the margin, then only international prices adjust at the rate of 2.1 percent every month to eliminate the deviation. In the short run we find the Granger causality results to be no different to the other regimes, where we find feedback effects. The Ljung-Box Q test statistics show that there are no issues with serial correlation.

It has been well documented in the literature that supply shocks, such as adverse weather, can cause large variations in price, which can be further exacerbated if the demand functions for a commodity such as coffee is price inelastic, causing innovations in the price margin. To analyse the short run adjustment in the margin, we conduct an innovation accounting exercise by making use of generalised impulse response analysis. As noted by Potter (1996), the response to a price shock in models that show symmetric adjustment are independent of the history of the time series and the sign and magnitude of the given shock. However, for models that display asymmetric adjustment, which implies nonlinearity, the impulse response functions are functions of the history of the price series and the sign and magnitude of the shock. Since we are conducting impulse responses on a MTAR model, we adopt the nonlinear approach proposed by Koop et al., (1996) where they make use of a generalised impulse response function. We follow the procedure as described in the earlier section by Coakley et al., (2001) where they highlight the superiority of the generalised impulse response function in a MTAR model.

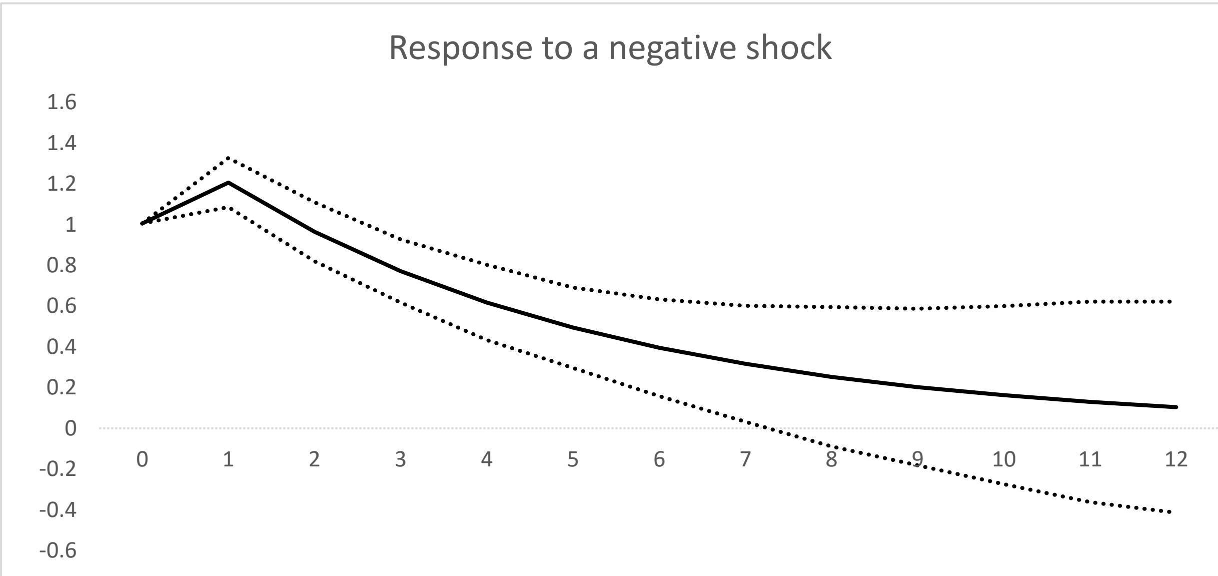
The impulse response function is estimated by averaging the 100 individual draws. The response to a positive and negative shock are given in Figures 3 and 4 respectively.

Figure 3. Generalised Impulse Response function to a positive change $\Delta Z_{t-1} \geq 0$ shock



The solid line shows the response. The dotted lines are the standard error bands. The horizontal axis shows the time horizon of 12 months.

Figure 4. Generalised Impulse Response function to a negative change $\Delta Z_{t-1} < 0$ shock



The solid line shows the response. The dotted lines are the standard error bands. The horizontal axis shows the time horizon of 12 months. The shock is normalised to a unit change.

A response is computed for each phase, one phase being where the change in retail prices is greater than international prices (labelled as a positive deviation) shown in Figure 3, and the other phase being where the change in retail prices is less than international prices (labelled as a negative deviation) shown in Figure 4. In both cases the responses for each phase are computed by randomly selecting histories from the data, such that $\Delta Z_{t-1} \geq 0$ (positive change) and $\Delta Z_{t-1} < 0$ (negative change). The horizon is set to be 12 months and all the responses are normalised so that the initial effect of the shock is unity for all histories. Figure 3 shows the response of the price margin to a positive unit shock. As expected, the response of a shock is insignificant as shown by the standard error bands containing the value zero over the entire time horizon. The response after a unit shock is generally flat, after a very slight increase in the month following the shock. This is expected, given the insignificant estimate that we obtain for γ_1 from the MTAR model (see Table 3). The impulse response that we obtain for a negative shock is shown in Figure 4. We find that in response to a unit shock, the margin increases slightly in the next month, but there after starts to gradually dissipate. The response is no longer significant after 7 months which underscores the rate of adjustment being significant in the MTAR model for negative deviations in the margin.

5. Conclusion

This paper examines the price dynamics of the margin between retail and international coffee prices. Analysing monthly data from 1980 to 2018, we make several contributions. First, we find no evidence of a statistically significant increasing trend in the margin between retail and international prices. Our findings lend support to Mehta and Chavas (2008) and are backed by robust tests for estimating the trend in the margin. We analyse this issue further by conducting tests for structural breaks, and we find no evidence of the possibility of any breaking trends. We conclude the trend estimate is statistically insignificant, and the variability in the margin dominates any possible underlying trend in the margin. Secondly, we establish that any deviations in the margin are transitory for the full sample as well as the periods prior to and after the demise of the ICA, but with asymmetric adjustment. In particular, positive discrepancies tend to persist, whereas negative discrepancies are corrected through time. Thirdly, we uncover the short run and long run dynamics from the AECM. The long run analysis shows that adjustments to the increase or decrease in the margin are asymmetric. In the case of the full sample, during the phase where the margin shows a positive discrepancy, only retail price adjusts at a very slow but significant rate to correct the discrepancy. During the phase when the margin has a negative discrepancy, retail and international prices adjust to correct the deviation. In this phase, when the adjustment takes place, the retail prices adjust at a relatively slower rate compared to the international prices. This pattern changes in the regime prior to the collapse of the ICA as well as the regime after the demise of the ICA. In the regime when the ICA was in place, retail and international prices only adjust to a negative deviation. In the post-ICA regime, retail prices adjust to correct a negative deviation while international prices adjust to correct a positive deviation. In each case we find when the margin shows a negative discrepancy, the retail prices adjust to correct the deviation, thereby attempting to maintain their margin. One of the reasons for the observed asymmetry we find for the entire sample could be market concentration (i.e. oligopsony power) in the coffee supply chain at the roasting level, which allows roasters to keep a higher share of the rents/profits by keeping the retail prices higher compared to the international prices. Our results lend support to the idea that power might be concentrated in the hands of large roasters and policies may be needed to be devised that help promote competition. For example, they could include promoting greater market diversification in the coffee roasting industry to increase competition in the coffee roasting sector; promoting coffee roasting facilities in large coffee producing countries to reduce the distance between coffee suppliers and coffee roasters for greater coffee market integration; and providing easier access to credit finance and price risk management (financial) instruments to economic agents in the coffee supply chain to improve competition in the coffee market.

A limitation of the study is that the margin between retail and international prices is not simply a markup over marginal cost but includes costs of conversion of coffee to roasted ground coffee, and therefore changes in the conversion costs will result in changes in the margin. As stated earlier, we choose roasted ground coffee since the product has been generally homogenous over the years. Although the conversion process from coffee to roasted ground coffee does include other inputs such as labour and machinery, these inputs are relatively lower for roasted ground coffee compared to other forms of coffee. Moreover, one can expect that increases in labour (wage) cost over the years is to some extent compensated by technological improvements in the conversion process. We should therefore keep this in mind in drawing conclusions relating to the trend in the margin and acknowledge that the price dynamics of the margin can be influenced by factors in addition to market power of roasters.

It may be worth mentioning that our study is restricted to mainstream coffee, while an important trend in the coffee market is the growth of niche and specialty markets (including sustainable coffee), making the coffee market highly differentiated.[17] Fitter and Kaplinsky (2001) argue that the benefits from the differentiated coffee market do not trickle down to coffee producing countries because roasters are buying a more homogenous coffee in the mainstream market (at more or less the same price) and differentiate their offering through product proliferation to increase their returns. Not everyone accepts this view. The other view is that the niche and specialty market demands differentiated coffee (higher quality or coffee that meets particular

[17] There is no universally accepted agreement on what constitutes as specialty coffee, so it is difficult to exactly quantify the share of specialty coffee of total coffee sold in retail outlets, though industry reports estimate the share to be over 20 percent in the US and European market.

production standards), and such coffee is usually in limited supply, which allows their suppliers to capture higher prices. We feel this is an area that calls for further research; for example, how coffee product differentiation affects coffee market competition and coffee price dynamics.

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