Price Transmission in the Cocoa-Chocolate Chain

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Résumé

Les consommateurs ont souvent l'impression que les prix de détail répondent plus vite aux augmentations du prix de la matière première qu'aux baisses. Aussi, l'objectif de ce travail est de tester l'existence d'une transmission asymétrique des fluctuations de prix entre la matière première, la fève de cacao, et le produit fini, la tablette de chocolat sur le marché français. Deux formes d'asymétrie, ayant chacune une origine différente, sont recherchées : d'une part, une asymétrie dans la transmission des chocs positifs et négatifs, potentiellement liée à l'exercice d'un pouvoir de marché des industriels, et d'autre part, une asymétrie dans la transmission des grands et des petits chocs de prix liée à la présence de coûts d'ajustement. Les résultats, obtenus à partir d'une modélisation TAR du déséquilibre de prix par rapport à leur valeur de long terme, ne permettent pas de rejeter ces hypothèses. Sur la plus grande partie de la période couverte (1960-2003) le prix de la fève et le prix de la tablette évoluent indépendamment l'un de l'autre. Toutefois, au moment du boom du cacao (fin 70) le prix de la tablette répond rapidement à la hausse des cours de la fève tandis qu'à la fin des années 80, alors que le prix de la fève est retombé à un bas niveau, le prix de la tablette revient lentement vers l'équilibre.

Introduction

The aim of this paper is to explore the channels of transmission of the fluctuations in the world price of cocoa beans to the consumers of chocolate bars in France. More precisely, we look for asymmetric transmission of shocks between the price of cocoa beans and the price of the final good (chocolate bar). This case can be considered as an illustration of a more general pattern of asymmetric vertical price transmission in the commodity-end product chain. Indeed, there is an extensive literature on asymmetric price transmission particularly in the agricultural markets and food industry as well as in financial and gasoline markets (see Meyer and von Cramon-Taubadel, 2004, for a review). Generally speaking, there is a common perception among consumers that the retail prices respond faster to an increase in the price of raw material than to a decrease. For instance, Chen, Finney and Lai (2005) working on the US gasoline market have shown that the gasoline retail price responds faster to a crude oil increase than to a decrease.

The relative evolution of the prices of cocoa beans and of chocolate bar highlights two main phenomena. First, the cocoa price fluctuations are passed through consumers within one year. Second, transmission of price fluctuations appears to be asymmetrical. Sharp upward movements in the cocoa price seem to be more readily transmitted to the consumers than price falls. Numerous studies have highlighted the role played by imperfect competition and adjustment costs as potential causes for asymmetric price adjustment (Meyer and von Cramon Taubadel, 2004). Working on the chocolate marketing chain in France, Araujo and Brun (2007) showed that concentration in the processing and manufacturing industry during the 80s has resulted in transferring the price leadership to the large chocolate companies.

In this paper we focus on the nonlinear nature of the adjustment process between cocoa beans and chocolate bar prices that may result from market power and adjustment costs. Using threshold cointegration analysis, we consider two alternatives to the standard linear error correction model. First, we allow for two different short-run price dynamics depending on whether the deviation of the chocolate price from its long run equilibrium is above or below a critical threshold. This alternative which catches asymmetry in the transmission of positive and negative shocks may reflect a non-competitive behaviour in the chocolate industry. Second, a three-regime threshold autoregressive model allows taking into account the existence of adjustment costs. The underlying hypothesis is that small shocks are dampened by the processing industry while larger shocks are passed through consumers. As a consequence, price transmission is asymmetric not only according to the sign of the shock but also according to the size of the shock.

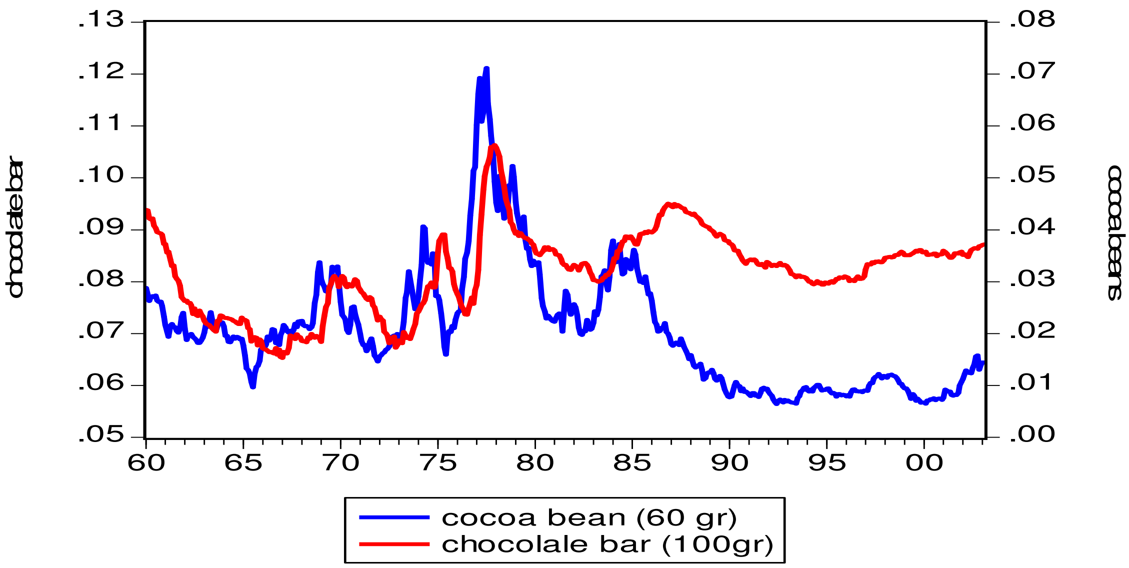
The paper is organised as follows. Section one presents the evolution of the monthly prices of cocoa beans and chocolate bar and briefly highlights the main features of the chocolate marketing chain. The dataset covers the period from January 1960 through February 2003. Section two exposes the TAR models of vertical price transmission under consideration. Section three develops the empirical testing strategy. Following numerous authors (*e.g.* Balke and Fomby 1997, Enders and Siklos, 2001, Goodwin and Piggott, 1999), we use the two-step approach to cointegration of Engle and Granger extended to encompass possible asymmetric adjustment to disequilibrium. Estimates indicate that a three-regime error correction model is the most appropriate. On the one hand, the chocolate price does not adjust to small shocks in the cocoa market. On the other hand, the speed of adjustment is larger for negative deviations than for positive ones.

1. The cocoa – chocolate chain

The relative evolution, over the period 1960 – 2003, of the world price of cocoa beans[1] and the retail price of the chocolate bar in the French market[2] highlights two phenomena (figure 1). First, cocoa beans price fluctuations are passed through chocolate bar price with a lag of roughly one year. Second, positive shocks in the cocoa price appear to be fully transmitted to the retail chocolate price. This phenomenon can be especially observed during the 70s. On the opposite, cocoa price decreases are passed through chocolate price in a lessen way or are not transmitted at all. For instance, cocoa price experienced a transitory fall during the years 1965-66 which is not reflected in the chocolate price. The same observation is valid for the period 1980-84. Moreover, from 1985 until the end of the 90s, the cocoa price experienced a long lasting phase of decline while the chocolate price stayed at a rather steady level.

These features of price movements in the cocoa chain suggest the existence of important lags and asymmetries in price transmission. Increases in the input price seem to be more fully transmitted to the output price than equivalent decreases.

Figure 1. Cocoa bean and chocolate bar prices (constant euros)



Among the different possible causes for asymmetric price transmission, two main causes have received a special attention in the literature: non-competitive market structure and adjustment costs which may result in two different types of asymmetry in the adjustment process.

[1] Average of London and New York Stock Exchange prices. Source: International Financial Statistics of the IMF.

[2] This is the retail price of a 100 gr. bar of black chocolate sold on the French market since 1960 (source INSEE). A 100 gr bar of chocolate contains approximately 60 gr. of cocoa bean. Both prices are in euros.

First, asymmetry in the transmission of positive and negative shocks may be due to imperfect competition in the processing/distribution chain. Agro-food industry is highly concentrated and processing companies and/or distributors are often accused to exert a market power. Processing industry may be prompted to pass, more rapidly and more fully, the increases in the input prices than the decreases, on to the consumers.

This hypothesis of price leadership in the cocoa processing industry has been tested by Araujo and Brun (2007) in a game theory framework. The authors show that Cote d'Ivoire, the main cocoa producing country, has lost its market power during the 80s to the benefit of the cocoa industry which now plays a leading role in price formation.

Indeed the cocoa industry has experienced important changes during the period under consideration, leading to a significant increase in industry concentration. Every stage of the cocoa chain -from cocoa bean production to chocolate products distribution- is highly concentrated. The three largest cocoa producing countries (Cote d'Ivoire, Ghana, Indonesia) represent more than 70 % of cocoa world production while six multinational companies control 90 % of the chocolate processing, manufacturing and distribution.

Cocoa processing[3] requires large investments. An important movement of concentration has taken place in the processing industry since the second half of the 80s. At the world level three companies realize more than 70 % of cocoa processing: Barry Callebaut (51 %), ADM (11 %), Blommer (9%). Cargill and Cémoi play also an important role, each representing 6 % of market share[4].

At the same time, numerous mergers and acquisitions took place in the manufacturing industry[5]. Leading international companies in the chocolate industry are: Barry Callebaut (24%) Nestlé (21%), Mars (13%), Cadbury (12%), Kraft Foods (10%), Hershey (10%).[6] In France, the chocolate market is dominated by Barry Callebaut (60-70%), Cargill (10-20%) and Cémoi (10-20%). These companies are vertically integrated and control each stage of the processing and manufacturing of chocolate.

The distribution sector is also highly concentrated. Hypermarkets and supermarkets are the major sellers of chocolate products (80%). They sell the quasi totality of the chocolate bars consumed in France. This sector also experienced an increased concentration during the last decade.

The high concentration in cocoa processing, manufacturing and distribution may thus result in non-competitive situations and asymmetric price transmission.

Second, adjustment costs in the packaging and distribution stage of the marketing process are also a possible cause of asymmetry in price transmission according to the size of the shocks. Changing prices generate so-called "menu costs" (Barro, 1972). For instance the costs of reprinting price lists or catalogues may lead to late and asymmetric adjustment of prices. Fixed adjustment costs are expected to create a price band inside which the retail price does not adjust to fluctuations in the raw material price as the adjustment cost would exceed

[3] Process during which cocoa beans are transformed in nib, liquor, butter, cake and powder.
[4] Data concerning 2003.
[5] Process that consists in the blending and refining of cocoa liquor, cocoa butter and other ingredients such as milk and sugar.
[6] Source : CNUCED, Infocomm. See also aussi Dorin (2003).

the benefit. As a consequence, processors and/or distributors respond to "small" input price fluctuations by increasing or reducing their margins. The output price will adjust only if the fluctuations in the input price exceed a critical level. Moreover, the adjustment costs are not necessarily symmetric with respect to price increases or decreases (Meyer and von Cramon-Taubael, 2004, Peltzman, 2000, etc).

Non-competitive markets and adjustment costs may thus result in nonlinear price dynamic. Threshold effects occur on the one side, when large shocks bring about a different response than small shocks due for instance to the presence of adjustment cost, and on the other side, when positive shocks (increases in the input price) trigger a different response of the output price than negative shocks. In the next section we develop a nonlinear model of price transmission that takes into account these two types of asymmetry.

2. Models of vertical price transmission

As the price series of the cocoa beans and of the chocolate bar are I(1) (see table 1 below,), we examine the relationship between the two series in a cointegration framework. Standard models of cointegration assume linear and symmetric adjustment towards the long run equilibrium relationship. We consider here the non linear model developed by Tong (1983) namely a threshold autoregressive model.

Standard linear and symmetric adjustment

The long run relationship between the two prices can be written:

$$Pc_t = \alpha_0 + \alpha_1 Pb_t + \mu_t \qquad (1)$$

Pc is the price of the chocolate bar. *Pb* is the price of the cocoa bean.

Cointegration of *Pc* and *Pb* depends upon the nature of the autoregressive process for μ_t. In the standard model of cointegration μ_t is stationary with zero mean and may follow a linear AR(p) process in the form:

$$\mu_t = \gamma_0 + \sum_{i=1}^{p} \gamma_i \mu_{t-i} + e_t \qquad (2)$$

e_t is a white noise disturbance.

If the series are cointegrated, the Granger representation theorem guarantees the existence of an error-correction representation of the variables in the form:

$$\Delta P_{it} = \lambda_i \mu_{t-1} + \sum_{j=1}^{p} \theta_{ij} \Delta P_{it-j} + \varsigma_{it} \qquad (3)$$

With $P_{it} = Pc_t, Pb_t$ and ζ_{it} is a white noise disturbance.

μ_{t-1} is the error correction term.
λ_i measures the speed of adjustment of ΔP_{it} to a deviation of P_{it} from its long-run equilibrium level.

In this standard error correction model (ECM), adjustment is linear and symmetric: λ_i is constant and negative. At every period, a constant proportion of any deviation from the long run equilibrium is corrected, regardless of the size or the sign of the deviation and the system moves back towards the equilibrium.

However, as discussed above, different types of market failures may prevent a continuous and linear adjustment of prices. Threshold cointegration model are thus considered.

A general threshold adjustment model

If *Pc* and *Pb* are characterised by threshold cointegration and asymmetric adjustment, then the equilibrium error, μ_t, will follow a self-exciting threshold autoregression (TAR). In such a model the autoregressive decay depends on the state of the variable of interest, μ_{t-d}.

The general form of a TAR(k; p, d) process is given by:

$$\mu_t = \gamma_0^{(j)} + \sum_{i=1}^{p} \gamma_i^{(j)} \mu_{t-i} + e_t^{(j)}, \qquad r_{j-1} \leq \mu_{t-d} < r_j \tag{3}$$

where k is the number of regimes that are separated by k-1 thresholds r_j (j = 1 to k-1). In each regime, μ_t follows a different linear autoregressive process depending on the value of μ_{t-d}. d is the threshold lag or delay parameter. It represents the delay in the error correction process. $e_t^{(j)}$ are zero-mean, constant-variance i.i.d random variables.

The stationarity of the threshold autoregressive process depends crucially on the behaviour of μ_t in the outer regimes. The equilibrium error may behave like a random walk inside the threshold range, but as long as it is mean-reverting in the outer regimes it is a stationary stochastic process (Balke and Fomby, 1997).

In the remainder of the section, we consider two special cases:

- a two regimes model allowing asymmetric adjustment to deviations in the positive and negative directions,
- a three regimes model with two asymmetric thresholds which, in addition, takes into account asymmetry in the correction of large and small deviations.

A TAR model with two regimes

A threshold autoregressive model with one threshold allows asymmetry in price adjustment to positive and negative deviations from the long run equilibrium. The TAR(2; p, d) is given by:

$$\mu_t = \begin{cases} \rho_0^{(1)} + \sum_{i=1}^{p} \rho_i^{(1)} \mu_{t-i} + e_t^{(1)} & \text{if } \mu_{t-d} \geq \tau \\ \rho_0^{(2)} + \sum_{i=1}^{p} \rho_i^{(2)} \mu_{t-i} + e_t^{(2)} & \text{if } \mu_{t-d} < \tau \end{cases} \tag{4}$$

τ is the unknown threshold value.

The corresponding error correcting model can be written:

$$\Delta P_{it} = \rho_{i1} I_t \mu_{t-1} + \rho_{i2}(1 - I_t)\mu_{t-1} + \sum_{j=1}^{p} \theta_{ij} \Delta P_{it-i} + \varsigma_{it} \tag{5}$$

In this representation, the speed of adjustment differs according to whether the deviation from long run equilibrium is above or below a critical threshold. The speed of adjustment is given by ρ_{i1} if μ_{t-d} is above the threshold level and by ρ_{i2} if μ_{t-d} is below the threshold value.

A TAR model with three regimes

A particular attention is paid to the case where deviations, in the positive and negative directions, from the long run equilibrium relationship lead to a price response only if they exceed a specific threshold level. In that case, the cointegrating relationship is inactive inside a given range and then becomes active once the system gets too far from the equilibrium band. This situation is captured by a TAR(3; p, d) model:

$$\mu_t = \begin{cases} \gamma_0^l + \sum_{k=1}^{p} \gamma_k^l \mu_{t-k} + v_t^l & \text{if } \mu_{t-d} < \mathrm{c}_1 \\ \gamma_0^m + \sum_{k=1}^{p} \gamma_k^m \mu_{t-k} + v_t^m & \text{if } c_1 \le \mu_{t-d} \le c_2 \\ \gamma_0^u + \sum_{k=1}^{p} \gamma_k^u \mu_{t-k} + v_t^u & \text{if } \mu_{t-d} > c_2 \end{cases} \tag{6}$$

v_t are zero-mean random disturbances with constant standard deviation.

An equivalent vector error correction representation is given by:

$$\Delta P_{it} = \begin{cases} \pi_0^l + \phi^l \mu_{t-1} + \sum_{k=1}^{p} \gamma_{1,k} \Delta P_{it-k} + \varsigma_{1,t} & \text{if } \mu_{t-d} < \mathrm{c}_1 \\ \pi_0^m + \phi^m \mu_{t-1} + \sum_{k=1}^{p} \gamma_{2,k} \Delta P_{it-k} + \varsigma_{2,t} & \text{if } \mathrm{c}_1 \le \mu_{\mathrm{t-d}} \le c_2 \\ \pi_0^u + \phi^u \mu_{t-1} + \sum_{k=1}^{p} \gamma_{3,k} \Delta P_{it-k} + \varsigma_{3,t} & \text{if } \mu_{\mathrm{t-d}} > c_2 \end{cases} \tag{7}$$

with $P_{it} = Pc_t, Pb_t$

The two thresholds c_1 and c_2 define three price regimes. The parameters ϕ_1, ϕ_2, and ϕ_3, measure the adjustment speeds of one price to deviations from the equilibrium relationship.

A case often encountered is when $\phi_2 = 0$: small deviations from equilibrium are not corrected (the price series are not cointegrated). Deviation from the equilibrium must reach a critical level before triggering a price response.

If $|c_1| \neq |c_2|$ the interval $[c_1, c_2]$ is not symmetric around the origin and deviations in the positive and negative directions must reach different magnitudes before triggering a price response. This case is more likely when adjustment costs are asymmetric.

3. Empirical application – testing procedure

We follow the threshold cointegration method introduced by Balke and Fomby (1997) and developed by Enders and Siklos (2001). It is a two-step approach that extends the Engle and Granger (1987) testing strategy by permitting asymmetry in the adjustment towards equilibrium. The first step involves the estimation of the long run equilibrium relationship between the price of chocolate and the price of cocoa; cointegration tests are then applied to the equilibrium error. The second step involves testing for nonlinear threshold behaviour. We consider two non linear autoregressive processes with respectively two and three regimes. In each case, tests for the difference in parameters across alternative regimes are conducted and confidence intervals for the threshold values are calculated. We then estimate the asymmetric error correction model corresponding to the best fit model.

3.1. Testing for no cointegration against linear cointegration

We first test for the order of integration of the price series using Dickey and Fuller (1981), Phillips and Perron (1988), and Kwiatkowski, Phillips, Schmidt and Shin (1992) tests (table 1).

Table 1: Unit root tests. monthly data, sample 1960.01 – 2003.02

	ADF	Prob	PP	Prob	KPSS
Price of the cocoa bean (Pb)	-1.88	0.34	-1.84	0.36	0.44***
Price of the chocolate bar (Pc)	-2.20	0.49	-2.16	0.51	0.33***

ADF: Augmented Dickey Fuller test. Ho: unit root
PP: Phillips- Perron. Ho: unit root
KPSS: Ho: I(0) ; critical values (Kwiatkowski-Phillips-Schmidt-Shin, 1992, Table 1)

The ADF and PP tests fail to reject the null hypothesis of unit root for both the cocoa and the chocolate bar price series while the first-differenced price series appear to be stationary[7] (table 1).

Following the two-step method of Engle and Granger (1987) we estimate the long run relationship between the chocolate and the cocoa price using OLS; the estimated equation is given in table 2. The Johansen procedure is conducted to test for linear cointegration between the two price series (table 3).

[7] Results not reported here.

Table 2. OLS Estimate of the cointegration relationship $Pc_t = \alpha_0 + \alpha_1 Pb_t + \alpha_2 T_t + \mu_t$

	α_0	α_1	α_2	Adj R²	SBC	AIC
Pc	-0.04	0.198	0.002	0.95	-2.93	-2.95
	(-6.67)	(3.45)	(85.5)			

T: trend variable
AIC: Akaike Information criterion - SBC: Schwarz criterion
t-Statistics are in parentheses.
Sample: 1960.01 – 2003-02.

Table 3: Cointegration testing results

Trace test: r=0	Max eigenvalue r=0	Max eigenvalue and trace test r=1
35.61	26.18	9.42
(0.002)	(0.004)	(0.16)

H_0: no cointegration, intercept and trend in the cointegrating equation, no trend in VAR

The Johansen tests clearly reject the null hypothesis of no cointegration between the price series of cocoa and of chocolate.

Table 4. OLS Estimate of the autoregressive process

$$\Delta\hat{\mu}_t = \gamma_0 + \gamma_1 \hat{\mu}_{t-1} + \sum_{i=1}^{p-1} \gamma_{i+1} \Delta\hat{\mu}_{t-i} + e_t$$

	γ_1	γ_2	γ_3	γ_4	Adj R²	SSR	AIC	White	Q(33)
$\Delta\mu_t$	-0.005	0.39	0.10	0.12	0.25	0.003436	-9.06	8.25	0.103
	(-2.13)	(8.88)	(2.04)	(2.65)				(0.00)	

t-Statistics are in parentheses
SSR: sum of squared residuals
Q(p) is the Ljung-Box statistic.
White's heteroskedasticity test and Pvalue in parenthesis

Table 4 gives the autoregressive process driving the residuals under the hypothesis of linearity that underlies the standard cointegration tests. The lag order of the linear AR(p) is selected according to the Akaike and the Schwartz information criteria. The former selects 4 lags while the latter selects 3 lags. An evaluation of autocorrelation patterns for the residuals led us to adopt a specification with four lags. We also test for the homoskedasticity of the residuals using the test proposed by White, Engle and Pagan and reject the null of homoskedasticity.

3.2. Testing for non linearity

We look for non linearity in the equilibrium error μ_t using two related tests: the non parametric test of Tsay (1989) and the CUSUM-type test of Petruccelli and Davies (1986). These two tests are based on predictive residuals (e_{ts}) obtained from an arranged autoregressive process where data are sorted according to the value of the potential threshold variable (μ_{t-d}). The sorting process allows to separate the different regimes. The first s cases[8] are located below the first threshold and correspond to the first regime. On this sub-sample the predictive residuals are then orthogonal to the regressors (μ_{ts}). The following cases fall into the second regime and the orthogonality between the predictive residuals and the regressors does not hold any more.

The Tsay's test is thus an orthogonality test which consists in regressing the predictive residuals on the explanatory variables. The test statistic is the *F*-stat associated to the regression:

$$e_{t_s} = \omega_0 + \sum_{v=1}^{p} \varpi_v \mu_{ts-1} + u_{t_s} \tag{8}$$

The test is run for alternative delays. Tsay (1989) proposes to select as an estimate of the delay parameter the delay giving the largest *F*-statistic.

In general the test results depend heavily on the selected AR order and on the starting point for the arranged autoregression. Following Tsay the start-up for the test is taken to be 10 % of the sample. The selected order of the AR process is four (see above).

The portmanteau test of Petrucelli and Davies (1986) is based on the cumulative sum of the standardized predictive residuals from the arranged autoregression. The standardized residuals are normalised by the square root of the residual variance computed recursively. The standardized and normalized residuals are standard Gaussian so that the P-value of the test can be evaluated by using an invariant principle for random walk.

Table 5. Nonlinearity tests - ascending order – b=52

d	1	2	3	4	5	6	7	8	9	10	11	12
Tsay	3.26	3.64	2.94	2.89	2.64	2.76	3.09	3.52	3.86	3.81	3.65	3.92
	(0.012)	(0.006)	(0.02)	(0.02)	(0.03)	(0.03)	(0.02)	(0.008)	(0.004)	(0.005)	(0.006)	(0.004)
PD	1.54	1.91	2.19	2.12	1.11	1.05	0.66	0.72	0.80	0.92	0.99	0.74
	(0.25)	(0.11)	(0.057)	(0.067)	(0.54)	(0.58)	(0.93)	(0.88)	(0.81)	(0.70)	(0.64)	(0.86)

Tsay: F stat and (P-value)
PD: Petruccelli and Davies : t stat and P-value.

Table 5 summarises the results of the portmanteau and the *F* tests. The *F* tests with *P*-values below 5 % strongly reject (for all d) the null hypothesis of linearity and thus imply the presence of one or more thresholds. These results contrast with those of the Petruccelli-Davies test that reject linearity only for d=3 and d=4. This may reflects the superiority of the Tsay's test in detecting non linearity (Tsay, 1989).

Given that the hypothesis of linearity is reasonably rejected, we turn to the identification of the thresholds and the estimation of the threshold autoregressive model.

[8] A « case » of data denotes each combination of μ_t and μ_{t-i}. The individual cases of data are ordered according to the threshold variable (μ_{t-d}).

3.3. Testing for TAR adjustment with one threshold

We estimate the threshold value (τ) along with the delay parameter (d) and the autoregressive adjustment process. We follow the methodology of Chan (1993) which consists to minimize the sum of squared residuals by searching over a set of potential threshold values for each possible delay parameters. The lagged residuals of the long run relationship are sorted in ascending order and the largest and smallest 7.5 % of the residuals are discarded. The remaining 85 % of the residual values are considered as potential thresholds. The threshold value and the delay parameter yielding the lowest sum of squared residuals are considered as the appropriate estimate of the threshold value.

The best fit TAR model is presented in table 6. The estimated delay parameter and threshold value corresponding to the lowest sum of squared error are respectively equal to 11 and –0.08531. To check for robustness we conduct the Enders-Siklos (2001) test of no cointegration against the alternative of cointegration with threshold adjustment. The test equation is given by:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1-I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p-1}\gamma_i\Delta\mu_{t-i} + \varepsilon_t \qquad (9)$$

τ : is the threshold value. ε_t is an iid process with zero mean and constant variance.
I_t is an indicator function that depends on the level of μ_{t-d}:
$I_t = 1$ if $\mu_{t-d} \geq \tau$ and $I_t = 0$ otherwise

The test statistics are the t_{MAX} statistic which is given by the larger t-statistic of ρ_1 and ρ_2, and the *F*-statistic, called Φ stat. Results are given in table 6.

Table 6. OLS Estimate of the TAR(2; 4, 11) model

	τ	ρ_0	ρ_1	ρ_2	SSR	Φstat Ho: $\rho_1 = \rho_2 = 0$	SupWald Ho: $\rho_1 = \rho_2$
$\Delta\mu_t$	–0.085831	-0.00	-0.001	-0.033	0.003286	15.12***	9.66
	[-0.0939; -0.0739]	(-2.13)	(-0.27)	(-3.25)			(0.048)[a]

t-stat in parenthesis.
99 % asymptotic confidence interval for τ in brackets.
White Heteroskedasticity-Consistent Standard Errors & Covariance.
Critical value for t_{MAX} (Enders and Siklos, 2001, table 2):
90 % : -1.69 ; 95 %: -1.89; 99 %: -2.29
Critical value for Φ (Enders and Siklos, 2001, table 1)
90 % : 5.21 ; 95 %: 6.33 ; 99 %: 9.09
a: Bootstrapped P-value for 1000 replications.

The t_{MAX} test does not reject the null hypothesis of no cointegration while the F-test does reject the absence of cointegration at the 1% level. The latter test having higher power (Enders and Siklos, 2001) we conclude that the residual are I(0). We then can test for symmetric adjustment.

Testing linearity versus a threshold alternative involves a non-standard inference problem as the threshold parameters are not identified under the null hypothesis. In such case conventional test statistics do not have standard distribution.

To test the null hypothesis of linearity versus the two-regime TAR model we follow the methodology proposed by Hansen (1996) and calculate the following standard *F*-statistic:

$$F_n = n\left(\frac{\tilde{\sigma}_n^2 - \hat{\sigma}_n^2}{\hat{\sigma}_n^2}\right) = \sup_{\tau\in\Gamma} F_n(\tau) \text{ with } F_n(\tau) = n\left(\frac{\tilde{\sigma}_n - \hat{\sigma}_n^2(\tau)}{\hat{\sigma}_n^2(\tau)}\right)$$

$\tilde{\sigma}^2$ is the residual variance under the null of linearity.

$\hat{\sigma}^2(\tau)$ is the residual variance under the alternative hypothesis of threshold autoregression calculated over all the possible threshold values.

We follow Hansen (1996) bootstrap procedure to approximate the asymptotic distribution of *F*.[9] Because there is evidence of heteroskedasticity in ε_t (White heteroskedasticity test) we replace the *F*-statistic $F(\tau)$ with a heteroskedascity-consistent Wald statistic and modify accordingly the bootstrap procedure (Hansen, 1997).

We also use the Hansen (1996) methodology to construct confidence-intervals for the threshold estimates.

$$LR_n(\tau) = n\left(\frac{\tilde{\sigma}_n^2(\tau) - \hat{\sigma}_n^2(\hat{\tau})}{\hat{\sigma}_n^2(\hat{\tau})}\right)$$

$\tilde{\sigma}^2(\tau)$ is the residual variance given the true value of the threshold.

$\hat{\sigma}^2(\tau)$ is the residual variance given the estimated value of the threshold.

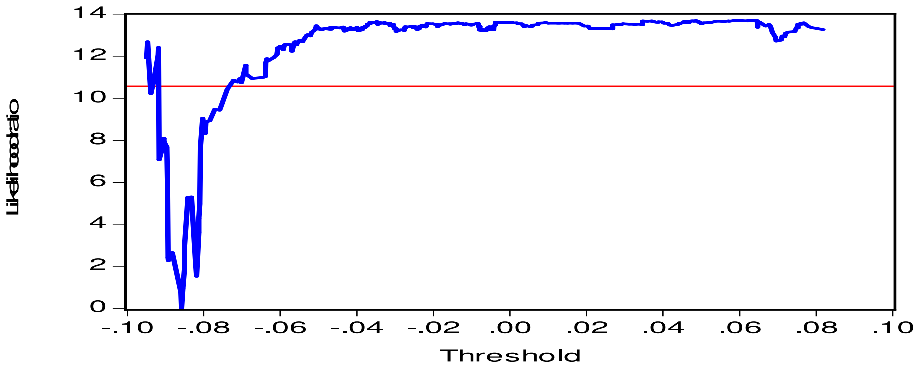
$LR_n(\tau)$ is the likelihood ratio statistic to test the hypothesis $\tau = \tau_0$. The β-percent confidence interval for τ, is given by: $\hat{\Gamma} = \{\tau : LR_n(\tau) \le c\ (\beta)\}$ where c (β) is the β-level critical value from the asymptotic distribution of $LR_n(\tau)$. Following Hansen (1997) we use the convexified region $\hat{\Gamma}^c = [\hat{\tau}_1, \hat{\tau}_2]$ where $\hat{\tau}_1$ and $\hat{\tau}_2$ are the minimum and the maximum elements of Γ.

In the heteroskedastic case, the modified likelihood ratio is: $LR_n^*(\tau) = n\left(\dfrac{\tilde{\sigma}^2(\tau) - \hat{\sigma}^2}{\eta^2}\right)$

and the modified likelihood ratio confidence region is $\hat{\Gamma} = \{\tau : LR_n^*(\tau) \le c\ (\beta)\}$. The nuisance parameter η is estimated using a polynomial regression (Hansen, 1997).

9 1000 simulations are run replacing the dependant variable by standard normal random draws. For each bootstrapped series the grid search procedure is used to estimate the threshold value and the $F_n(\tau)$ statistic is computed.

Figure 2. Confidence interval for the estimated threshold (1%)



The value of the sup-Wald statistic is 9.66 (table 6) and the resulting *p*-value is 0.048. Therefore, the null hypothesis of linearity is rejected. The two regimes TAR is thus preferred to a linear autoregressive process. The estimated threshold is $\tau = -0.0858$ with a 99% asymptotic confidence interval [-0.0939; -0.0739]. A plot of the adjusted likelihood ratio ($LR_n^*(\tau)$) is displayed in figure 2. This figure shows that the threshold estimate is quite precise.

3.4. Testing TAR adjustment with 2 thresholds

A three regimes TAR model is fitted to μ_t by minimizing the sum of squared residuals with respect to the threshold and delay parameters, maintaining the lag length at four. Following Goodwin and Piggott (1999), a two dimensional grid search is used, to estimate the two thresholds c_1 and c_2. As a practical matter, we search for the first threshold between 5% and 95 % of the largest (in absolute value) negative residuals and we search for the second threshold between 5% and 95 % of the largest positive residuals.

This estimator selects a delay parameter of 11 and produced the TAR model given in table 7. Three tests are conducted. First, we check the stationnarity of the threshold autoregressive process with a standard Wald test on White heteroskedasticity consistent parameters estimates. The bootstrapped *p*-value is obtained imposing a linear autoregression with a unit root under the null. The calculated test statistic equals 7.30 and the bootstrapped *P*-value is 0.03 (table 7). We thus reject the non-stationnarity hypothesis at conventional levels.

Second, a test to select the appropriate number of regimes for the TAR model is used. To test the null hypothesis of a two regimes TAR model versus the alternative of three regimes, we use the following likelihood ratio statistic: $F_{23} = n\left(\frac{\hat{\sigma}_2^2 - \hat{\sigma}_3^2}{\hat{\sigma}_3^2}\right)$

$\hat{\sigma}_2^2$ is the residual variance from the estimated two-regime TAR model.

$\hat{\sigma}_3^2$ is the residual variance from the estimated three-regime TAR model.

The value of F_{23} is 6.25 and the bootstrapped *P*-value is 0.078. Therefore we do reject the null of a two-regime TAR model against the alternative of a three-regime TAR model.

In a third stage we test the significance of the differences in parameters across regimes using the supWald test defined above. The *P*-value of the test statistic is 0.055 for 1000 replications suggesting that the threshold effects are statistically significant (Table 7).

The estimated threshold values are -0.086 and 0.069. The first regime includes 33 observations, the second 400 and the third 75 observations. The two thresholds are not exactly symmetric around zero suggesting that negative deviations from the long run equilibrium must reach a slightly higher level (in absolute value) than positive deviations before triggering a response in the chocolate price. This suggests asymmetry in adjustment costs.

As long as μ_{t-11} is inside the band - defined by the two thresholds - μ_t acts as a unit root process and consequently has no tendency to drift back towards some equilibrium. When μ_{t-11} is above c_2, μ_t becomes an I(0) process which tends to revert back to the upper border of the band. In the same way, when μ_{t-11} is below c_1, μ_t is I(0) and tends to revert even quicker to the lower border of the band.

Table 7. OLS Estimate of the TAR(3; 4, 11) model

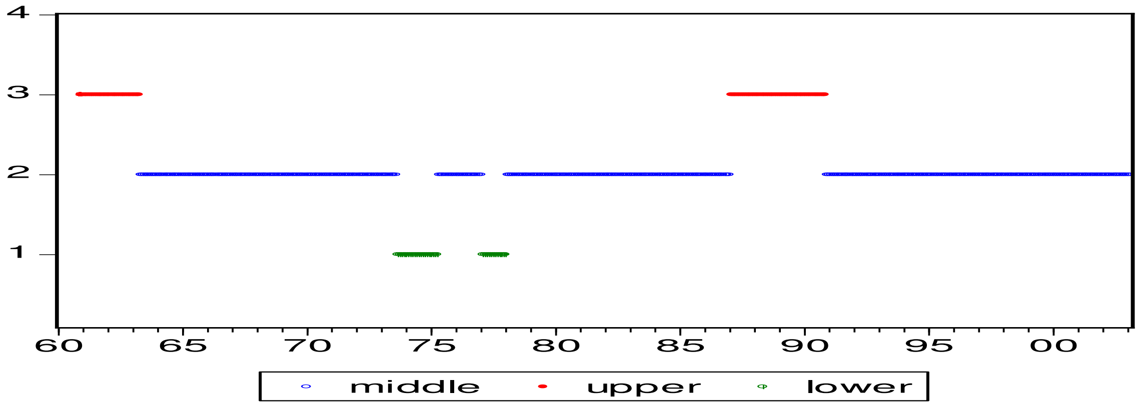
	γ^l	γ^m	γ^u	Regimes (obs)	Estimated thresholds	Φ $\gamma^l = \gamma^m = \gamma^u = 0$	SupWald $\gamma^l = \gamma^m = \gamma^u$	F_{23}
$\Delta\mu_t$	-0.033	0.005	-0.009	Regime I (32)	$-\infty < \mu_{t-11} \leq -0.085831$	7.30	8.43	6.25
	(-3.20)	(1.46)	(-3.10)	Regime II (400)	$-0.085831 < \mu_{t-11} \leq 0.06933$	(0.031)[a]	(0.055)[a]	(0.078)[a]
				Regime III (75)	$0.06933 < \mu_{t-11} < \infty$			

White Heteroskedasticity-Consistent Standard Errors & Covariance
F_{23}: Ho: TAR (2,4,11) versus TAR(3, 4, 11)
a: Bootstrapped-calculated asymptotic P-value using 1000 replications.

Figure 3 depicts the regime shifts during the period under consideration. During the major part of the period, the deviations from the long run equilibrium relationship - linking the price of chocolate to the price of cocoa - fall inside the band regime. Within the band prices are not cointegrated. During a short period of time, 1973-1979, corresponding to the two successive booms in the world cocoa prices, the deviation from the long run equilibrium falls below the first threshold and the chocolate bar price is well above its long run equilibrium value. At the opposite the 1987-1991 period corresponds to a phase of extremely low cocoa prices. The chocolate price is then well above its long run equilibrium value and tends to move back toward the equilibrium but rather slowly.

The timing of the regime shift may reflect a change in market structure and price leadership in the cocoa chain. Araujo and Brun (2007) have shown that until mid 80s Côte d'Ivoire can be considered as a price leader. The major cocoa producing country lost the price leadership at the end of the eighties to the benefit of the processing and distributing industry.

Figure 3. Timing of regime switching



3.4. Short run dynamics

The error correction model for the chocolate bar shows that the chocolate price displays asymmetric error correction toward long run equilibrium (table 8). The price of chocolate adjusts faster to negative shocks than to positive shocks. In the middle regime the coefficient for the error correction term is not significantly different from 0. Deviations of the chocolate price from its long run equilibrium thus have to reach a critical level before adjustment operates. For small deviations the chocolate and the cocoa prices move independently one from each other.

Table 8. OLS Estimate of the error correction model

	π_0	ϕ^l	ϕ^m	ϕ^u	Wald $\phi^l = \phi^m = \phi^u=0$	Wald $\phi^l = \phi^m = \phi^u$	Wald $\phi^l = \phi^u$
ΔPc_t	0.00	-0.027	0.001	-0.008	7.87	4.86	3.66
	(4.09)	(-2.81)	(0.29)	(-2.94)	(0.00)	(0.01)	(0.06)

White Heteroskedasticity-Consistent Standard Errors & Covariance

The Wald tests reject the different null hypothesis of linear and symmetric adjustment.

Concluding remarks

The tests clearly reject linear cointegration between the price of cocoa beans and the price of chocolate bar. The results tend to favour a three regimes error correction model. Most of the time, the price of chocolate does not adjust to changes in the price of cocoa beans. This may partly be explained by excessive adjustment costs and stocks management.

Two periods of adjustment are identified corresponding on the one hand, to an historical cocoa boom on the world market, and on the other hand, to a sharp decline in the cocoa price and, at the same time, to an increased concentration in the chocolate industry. In the first case, the chocolate price responded rather quickly to the disequilibrium. In the second case, the chocolate price reverted slowly to the equilibrium band. This asymmetry in the pass through of large price shocks to consumers may reflect a non-competitive market structure.

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