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Do retail coffee prices increase faster than they fall? Asymmetric price transmission in France, Germany and the United States

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Abstract.
This investigation examines price transmission asymmetries (PTA) between international and retail coffee prices in the US, France and Germany. Differences in price transmission mechanisms provide evidence for disparities in market structure and market performance across countries. Although all processors of roasted coffee purchase green coffee at the same price in the international markets, one finds significant differences in retail prices among these countries. The study develops an Error Correction (EC) representation model to assess PTA of non-stationary models. Finally, it claims that identifying differences in price transmission asymmetry is an approach to compare market structure across countries.

Keywords
Asymmetric Price Transmission; Roasted Coffee Market; Germany; United States; France; Error Correction Model.

JEL classification.
C13, C22, D43, L13, L11, L66.

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1. Introduction

There is ample evidence in the applied economics literature of price transmission asymmetry in markets for agricultural commodities. The causes of these asymmetries have been extensively studied in the past three decades and at the same time new econometric methods have been developed to measure their effects. Chief among drivers of transmission asymmetry are the exercise of market power by firms as well as the high costs of inventory adjustment. In particular, various studies report that increases in factor prices are often more quickly transmitted to the consumer than decreases in factor prices. This observed behaviour is particularly relevant to the study of marketing margins in the food and fibre industry because this industry segment has experienced substantial increases in the level of concentration in recent years worldwide. Identifying the occurrence of transmission asymmetries is relevant from a policy point of view, because they suggest market failure and, in extreme cases of exertion of market power, might grant government intervention.

The objective of this study is to test the existence of price transmission asymmetries between international and retail coffee prices in the US, France and Germany, the three largest coffee importing countries. Differences in price transmission mechanisms, in turn, provide insights in the study of disparities in market structure and market performance across countries. The motivation of this study is that although processors of all three countries purchase green coffee at similar prices in the international market, one finds significant differences in retail prices among these countries. These differences in retail prices are intriguing, because green coffee represents nearly 70 percent of processing cost to the industry (source). For instance, in the US the average retail price over the period 1990-2000 was $3.55, while the French and German counterparts were $2.58, and $4.25, respectively. This brief comparison suggests that the national coffee markets are far from similar, substantiating a formal test of asymmetries in price transmission in the US, France and Germany. This empirical study claims that identifying differences in price transmission asymmetry is a valid approach to compare market structure across countries (*von Cramon-Taubadel* 1997).

The study is organized as follows: Section Two is a review of relevant literature on price transmission asymmetry (thereafter referred to as PTA) focusing on the theoretical explanations of its occurrence as well as on empirical approaches. Section Three presents an asymmetric error correction representation to empirically test the existence of short and long run asymmetries of price transmission. Section Four presents the data and the study findings and Section Five is a discussion of how the results can shed light on differences in market performance across countries.

2. Review of the literature

Interest in the study of price transmission mechanisms goes back to Keynesian economics postulates explaining the process of wage and prices adjustment over time. The macroeconomic literature offers a large amount empirical research on price adjustment over time and asymmetries of price transmission (cf., *Mankiw and Romer* 1991). These studies contributed largely to the development of a theoretical framework to examine price transmission asymmetries. On the one hand PTA is viewed as the result of frictions in price setting at the microeconomic level such as the cost of price adjustment and the staggered timing of price changes. On the other hand, at a more aggregate level, PTA is regarded as the consequence of imperfect competition, including demand externalities and coordination failures.

These principles have been widely employed in the applied economics literature and in particular to construct testable models of vertical and spatial price transmission in the food sector. This is because various studies demonstrated asymmetries and lagged response in vertical price transmission. Many studies, for example, examine vertical and spatial PTA in terms of frictions in price setting and exertion of market power (*Ward* 1982; *Kinnucan and Forker* 1987; *Bailey and Brorsen* 1989; *Azzam* 1999). There are various aspects when considering the vertical dimension of price transmission. For example, price transmission can be asymmetric in the sense that the price reaction at one level of the marketing chain to changes of prices in another level of the marketing chain depend on the sign (positive or negative) of the initial change. Furthermore, it is possible to distinguish between short and long run PTA (*von Cramon-Taubadel and Loy* 1999). Short run PTA refers to the speed of reaction to price changes while long-run PTA points to the magnitude of the change. Examining PTA also requires addressing the problem of simultaneity in price determination (cf., *Ward* 1982; *Hahn* 1992; *von Cramon-Taubadel* 1997).These considerations suggest that examining PTA calls for careful econometric treatment.

Econometric methods employed to the study of PTA have changed over time. Earlier empirical procedures were developed by *Wolffram* (1971) and later improved by *Houck* (1977). Many assessments of PTA in the food system adopted these methodologies with mixed results (cf., *Kinnucan and Forker* 1987; *Boyd and Brorsen* 1988; *Appel* 1992; *Hansmire and Willett* 1992; *Zhang et al.* 1995). Nevertheless, *von Cramon-Taubadel* (1997) points out that these studies often disregard the time series properties of the data. Thus the problem of spurious regression arises because these time series data on prices are commonly integrated. Moreover, *von Cramon-Taubadel* (1997) challenges the *Wolffram* specification demonstrating that it is no appropriate when prices in the marketing chain are co-integrated.

More recently, the attention turned to empirical procedures based on the model developed by *Engle and Granger* (1987) and extended by *Granger and Lee* (1989) to test for PTA behaviour. They develop a formal model showing that when two price series are co-integrated, there exists an error correction (EC) representation that describes their short and long run relationship as well as the inherent price transmission mechanism. Indeed, the second half of the 1990s saw an increasing interest in EC models as means to assess PTA in a wide variety of commodities and provide a way to discuss market performance and behaviour. Borenstein, *Cameron and Gilbert* (1997), for instance, find that retail gasoline prices respond more slowly to decreases than to increases in crude oil prices. The authors also suggest that short-run market power of retailers produce PTA from retail prices to wholesale price changes.

Various studies conducted by *von Cramon-Taubadel* and *Loy* pioneered on the application EC representation to examine PTA in markets for agricultural commodities. First, *von Cramon-Taubadel and Loy* (1996) challenge methods utilized to discuss price asymmetry in the international wheat market by arguing that if price series are co-integrated then it is not possible that long run PTA occurs. Subsequently, *von Cramon-Taubadel* (1997) points out that many PTA studies utilized econometric methodologies inconsistent with co-integration, presenting an EC representation to demonstrate PTA in the German pork marketing chain. Limitations of traditional models and adequacy of EC representations to investigate PTA when price series are co-integrated is formalized later in *von Cramon-Taubadel and Loy* (1999) with an application to the international wheat market. Recent applications, all focusing on meat markets, also include *Goodwin and Holt* (1999), *Sanjuan and Gil* (2001) and *Hayenga* (2001).

The utilization of EC representations has been extended to other issues relevant to market structure at both microeconomic and macroeconomic levels. These applications have considered various issues including market efficiency in temporal and spatial dimensions (*Sabuhoro and*

Larue 1997; *Yang and Leatham* 1998), pricing strategies under oligopoly (*Vickner and Davies* 2000), extortion of market power in international trade (*Gómez and Castillo* 2001) as well as price spreads of agricultural commodities (*Chang and Griffith* 1998). EC representations have also been employed in macroeconomic contexts to examine the dynamic relationships between wholesale and retail prices (*Hondroyiannis and Papapetrou* 1988; *Kulger* 1989).

A final note on the utilization of asymmetric EC representations is pertinent. Although the link between theory and observed PTA behaviour is not straightforward, findings of the aforementioned studies suggest that EC models can shed light on the discussion of market structure (e.g., exercise of market power or high costs of adjustment). There are alternative ways to make the link between observed behaviour and theory more direct. For example, *von Cramon-Taubadel* argues that one alternative is to conduct PTA assessments in countries with different market structures. Comparing results across countries, in turn, provide information about the incidence of market structure on price transmission mechanisms. Such is the case of this investigation.

It is pertinent to examine empirical findings of the international coffee market because it has experienced substantial changes in recent years. Before 1990, most coffee exporting countries were part of the *International Coffee Agreement* (*ICA*) which fixed a system of export quotas to meet a target price above competitive prices (*Bates* 1997). Importing countries supported *ICA* probably because they saw it as an efficient way to provide assistance to developing countries (*Bohman, Jarvis, and Barichello* 1996). In 1990, however, *ICA* was eliminated and exporters relied on competition to maintain or gain market share in international markets. We hypothesize that *ICA*'s elimination did not lead to competitive markets but instead to a transfer of market power from exporting countries to international wholesalers in the 1990s.

Most studies on the international coffee market focus on the consequences of the *ICA* and commodity trade policies for coffee exporting countries (cf., *Bohman, Jarvis, and Barichello* 1996; *Bates* 1997; *Buccola and McCandlish* 1999; *Akiyama and Varangis* 1999; *Boratav* 2001). Nevertheless, after the elimination of the export quota system in 1990, researchers turned their attention to coffee markets in importing countries and their links to international markets. These studies commonly develop models of imperfect competition to assess market performance in coffee importing countries (cf., *Feuerstein* 2002). In particular, various empirical investigations examine price transmission mechanisms in importing countries. *Bettendorf and Verboven* (2000) address incomplete transmission of coffee bean prices to consumer prices in the Netherlands, demonstrating that market conduct is nearly competitive; *Koerner* (2002), on its part, report incomplete price transmission from factor to retail prices in the German market; and *Gómez and Castillo* (2001) utilize an EC representation to show that the coffee industry exercise market power in the US.

This study extends the research on price transmission mechanisms in coffee markets of importing countries, linking these mechanisms to market structure. In particular, the study applies an asymmetric EC representation to test for short and long run PTA. Findings of the empirical model as to differences in PTA across countries provide a link between imperfect competition and observed market performance.

3. Model of asymmetric price transmission

PTA can occur in the short-run and in the long-run and it is affected by the stochastic process governing prices. Consider, for instance, two price series that are believed to be interdependent. If these time series are integrated, but not co-integrated, then asymmetry in the long-run yields incomplete price transmission. The adjustment differences between positive and negative changes are cumulated over time and, as a consequence, there exists no stable long-run equilibrium. If, on the other hand, the time series are integrated and co-integrated, long-run PTA is inconsistent with theory and only short-run asymmetry can take place (*von Cramon-Taubadel and Loy* 1996). Second, PTA can arise in the short-run, as prices adjustment towards the long-run equilibrium. In this instance short-run asymmetry is compatible with co-integration and deviations from the long-run equilibrium are reduced at different speed of adjustment (*von Cramon-Taubadel and Loy* 1999).[1]

Employing an Error Correction (EC) representation is appropriate to measure PTA of non-stationary models. Starting with a rational lag distribution (*Jorgensen* 1966) the general model of an EC representation with two non-stationary time series (y_t and x_t) and two lags is given in (1).

$$(1)\ \Delta y_t = \alpha_0 + (\alpha_1 + \alpha_2 - 1)\left[y_{t-1} + \frac{\alpha_2 + \alpha_3 + \alpha_5}{\alpha_1 + \alpha_2 - 1} x_{t-1} \right] - \alpha_2 \Delta y_{t-1} + \alpha_3 \Delta x_t - \alpha_5 \Delta x_{t-1} + \varepsilon_t$$

The second term in brackets on the right hand side is the so-called error correction term (ECT), representing the deviation from the equilibrium in the previous period. Depending on the extent of the deviation, the ECT corrects the dependent variable in the following period toward the long-run equilibrium (*Banerjee et al.* 1993). Thus PTA can take place in both the deviation from equilibrium as well as in the 'short-run dynamics' (first and second differences on the right hand side). Following *Wolffram* (1971) and *Houck* (1977), these deviations can be segmented into positive and negative deviations from the long-run equilibrium, namely ECT_{t-1}^{+} and ECT_{t-1}^{-} respectively. For example, ECT_{t-1}^{+} equals ECT When the later is positive and zero otherwise. Therefore, adding up the segmented vectors ECT_{t-1}^{+} and ECT_{t-1}^{-} yields the original vector ECT_{t-1}. The same can be done for the variables expressed as first-differences. Equation (1) can be modified into its asymmetric representation:

[1] Long-run asymmetry implies short-run asymmetry, but not necessarily vice versa (*von Cramon-Taubadel and Loy* 1999).

$$\text{(2)}\quad \begin{aligned}\Delta y_t &= \alpha_0 + \overline{\alpha}^+ ECT^+_{t-1} + \overline{\alpha}^- ECT^-_{t-1} \\ &+ \alpha_2^+ \Delta^+ y_{t-1} + \alpha_2^- \Delta^- y_{t-1} + \alpha_3^+ \Delta^+ x_t + \alpha_3^- \Delta^- x_t - \alpha_5^+ \Delta^+ x_{t-1} - \alpha_5^- \Delta^- x_{t-1} + \varepsilon_t\end{aligned}.$$

where $\overline{\alpha} = \alpha_1 + \alpha_2 - 1$. Asymmetry tests can be utilized to determine whether or not the coefficients of the segmented variables ECT^+_{t-1} and ECT^-_{t-1} are equal. If $\overline{\alpha}^+ = \overline{\alpha}^-$, then PTA is rejected and prices adjust equally to positive and negative changes away fro the long-run equilibrium. The same holds for the estimated parameters of the differentiated variables.

Hitherto a single equation model has been considered implying that there is a unidirectional relationship between x_t and y_y. This is a restrictive assumption because interactive influences between these variables are possible and a simultaneous equation system is appropriate. Within this system, tests of exogeneity can be constructed to examine whether the co-integrating equation influences both or only one variable.[2] Such system must be identified. The estimated coefficients can be assigned unambiguously to the parameters of interest. Identification of the short-run dynamics in our model needs at least one restriction on each equation. This is given if the contemporise term is significant in only one of the two equations (*von Cramon-Taubadel and Loy* 1999). That is, one differentiated variable influences the endogenous in the first equation and not in the second (and vice versa). We need at least one restriction in the short-run parameters on each equation.

Transforming equation (2) into a system generates (3a) and (3b) with Δz_t and $\Delta z'_t$ as identifying variables in the short-run parameters:

$$\text{(3a)}\quad \begin{aligned}\Delta y_t &= \alpha_0 + \overline{\alpha}^+ ECT_{t-1} + \overline{\alpha}^- ECT_{t-1} \\ &- \alpha_2^+ \Delta^+ y_{t-1} - \alpha_2^- \Delta^- y_{t-1} + \alpha_3^+ \Delta^+ x_t + \alpha_3^- \Delta^- x_t - \alpha_5^+ \Delta^+ x_{t-1} - \alpha_5^- \Delta^- x_{t-1} \\ &- \alpha_6^+ \Delta^+ z_t + \alpha_6^- \Delta^- z_t - \alpha_7^+ \Delta^+ z_{t-1} - \alpha_7^- \Delta^- z_{t-1} + \varepsilon_{1t}\end{aligned}$$

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2 *Bessler* (1984a,b) suggested that the decomposition of the forecast error variance gives evidence of exogeneity. Within the ECM-framework tests of exogeneity are carried out straightforwardly, testing whether the co-integration relationship determines the endogenous variable in all equations (*Granger* 1986).

$$
(3b)\qquad \begin{aligned}\Delta x_t &= \beta_0 + \overline{\beta}^{+} ECT_{t-1} + \overline{\beta}^{-} ECT_{t-1} \\ &- \beta_2^{+}\Delta^{+} x_{t-1} - \beta_2^{-}\Delta^{-} x_{t-1} + \beta_3^{+}\Delta^{+} y_t + \beta_3^{-}\Delta^{-} y_t - \beta_5^{+}\Delta^{+} y_{t-1} - \beta_5^{-}\Delta^{-} y_{t-1}\ . \\ &+ \beta_6^{+}\Delta^{+} z'_t + \beta_6^{-}\Delta^{-} z'_t - \beta_7^{+}\Delta^{+} z'_{t-1} - \beta_7^{-}\Delta^{-} z'_{t-1} + \varepsilon_{2t}\end{aligned}
$$

The system (3a/b) is the appropriate model for identifying short-run asymmetry as long as there long-run PTA does to take place. In the next section, we employ this model to examine whether green coffee prices are transmitted asymmetrically. Additionally, we will analyse the causal relationship between green bean prices and retail prices of roasted coffee.

4. Data description and empirical results.

4.1 Data description

The empirical analysis utilizes monthly data from France, Germany and the United States for the period January-1990 to December-2000. We compile national retail prices of roasted coffee and international prices of green coffee from the International Coffee Organization (ICO). Retail prices of roasted coffee are denoted in US-Dollars per pound while international prices are a composite from different coffee varieties, expressed in US-Dollars.[3] Additionally, we use data on the exchange rates of Franc and German Mark to US-Dollar as well as the consumer price index in the United States (*IMF* statistics). Furthermore, the monthly average precipitation in Brazil is used as explanatory variable in the short-run dynamics.[4] Descriptive statistics of these data are given in Table 1.

[3] The indicator price is the arithmetical mean of the weighted average of daily prices for selected coffees of the *Other Mild Arabicas* and *Robusta* groups, calculated in accordance with procedures established under the *International Coffee Agreement*. The weighting reflects the participation of the groups in world trade. The prices are compiled daily from quotations for prompt shipment obtained from various major coffee markets (New York, Bremen/Hamburg and Le Havre/Marseilles) and are weighted to reflect the participation of the various coffees in world trade (*ICO*, 2002).

[4] Monthly average precipitation in mm at Fortaleza/Brazil (3.4S/38.3W) (*WMSSC* 2002).

Table 1: Descriptive statistics of the sample

	Name	Mean	Standard deviation	Minimum	Maximum
Composite price vector (in US$)	CP	0.924	0.361	0.459	2.024
Retail price in France (in US$)	RP^{f}	2.818	0.553	1.992	4.179
Retail price in Germany (in US$)	RP^{g}	4.575	0.630	3.150	6.180
Retail Price in the US (in US$)	RP^{us}	3.313	0.623	2.352	4.669
Exchange Rate Franc/US-Dollar	EX^{f}	5.701	0.607	4.831	7.694
Exchange Rate German Mark/US-Dollar	EX^{g}	1.682	0.191	1.381	2.294
Consumer Price Index in the US	CPI^{us}	1.519	0.127	1.274	1.741
Monthly average precipitation (in mm)	$RAIN$	105.462	144.819	0	668.000

4.2 Tests of integration

Most tests of integration assume non-stationarity under the null hypothesis and often tend to not reject it. To address this issue, stationarity should be tested under the null as well as under the alternative hypothesis. Common tests are The *Augmented Dickey-Fuller (ADF)*[5] and the *Phillips-Perron*[6] are examples of tests assuming non-stationarity under the Null hypothesis. Simulations have shown that, especially in small samples, both tests show lower diagnostic power than the *DF-GLS*-test[7] (*Elliott, Rothenberg, and Stock* 1996; *Elliott* 1999). The most commonly used test under the null of stationarity is the Lagrange-Multiplier-test of *Kwiatowksi et al.* (1992)[8] - the so-called *KPSS*-test.

We use the *ADF-t*-test as well as the *DF-GLS*-procedure to test non-stationarity under the Null and the *KPSS*-test to test stationarity under the null. Our results are robust to the alternative tests as well as to the deterministic processes (with or without constant trend respectively). Test results and critical values are given in Table 2.

5 See *Dickey and Fuller* (1979) and *Dickey and Fuller* (1981) as well as *Hall* (1986).

6 See *Phillips* (1987), *Perron* (1988) and *Phillips and Perron* (1988).

7 *Dickey Fuller*-Generalized Least Squares-test (*Elliott, Rothenberg, and Stock* 1996; *Elliott* 1999).

8 *Kwiatowski, Phillips, Schmidt, and Shin* (1992).

Table 2: Tests of integration

	Hypothesis	Critical values	Retail price in France	Retail price in Germany	Retail price in the US	Composite Price
ADF-t	$H0 :\sim I(1)$	-2.88	-1,.67	-1.53	-2.31	-1.44
	$H0 :\sim I(1)$ and no constant	4.63	1.44	1.28	2.68	1.06
DF-GLS	$H0 :\sim I(1)$	-1.95	-0.52	-0.59	-0.23	-0.92
	$H0 :\sim I(1)$ and no constant	-1.95	-1.73	-1.35	-1.86	-2.02
	$H0 :\sim I(1)$ and no linear trend	-2.89	-1.80	-1.49	-2.36	-2.17
KPSS	$H0 :\sim I(0)$ and no constant	0.46	0.55	0.45	0.53	0.47
	$H0 :\sim I(0)$ and no linear trend	0.15	0.16	0.18	0.15	0.19
Perron	$H0 :\sim I(1)$ and no break	-5.55	-6.77	-4.77	-6.72	-6.34

All retail price series are integrated of order one, ~ I(1), none of them has a deterministic trend and their first differences are stationary. The same is true for the composite price of green coffee. Additionally, we examined whether these time series include a time break. This break could change the intercept or the slope of the curve. *Perron* (1997) suggested a test of unit root with endogenous time break. The date of possible change in the intercept or the slope is not fixed a priori. While the composite, the French and the US retail prices show significant time breaks, the German retail price does not.[9] The significant time breaks give further information which should be used in the regressions.

4.3 Test of co-integration

Johansen (1991,1992a, 1995) as well as *Johansen and Juselius* (1992) proposed a test to determine whether two I(1) time series are co-integrated. The procedure allows identifying the number of equations that determine the co-integration equation. It tests the rank based of the matrix on canonical correlations. The test statistic of the trace test (Johansen 1988) is a likelihood

[9] Retail price in France in 1994:06, in the US in 1994:05 and the composite price in 1994:03. the possible time break in Germany was in 1994:05, but does not influence significantly.

ratio test defined by $trace = -T \sum_{i=r+1}^{n} \log\left(1-\hat{\lambda}_i\right)$ with T as total number of observations, r as number of co-integration relations and $\hat{\lambda}_i$ as eigenvalue. The principle is to determine how many eigenvalues equal one. The test is carried out until the null hypothesis cannot be rejected. Another common procedure is given by testing the significance of the estimated eigenvalues themselves. It is called the λ_{max} test with $\lambda_{max} = -T \log\left(1-\hat{\lambda}_i\right)$ as test statistic. Critical values are reported by *Osterwald-Lenum* (1992).

Tests of co-integration are sensitive to the structure of the data generating process, e.g. the underlying deterministic process such as constant and trend. *Johansen and Juselius* (1990) as well as *Osterwald-Lenum* (1992) consider three cases: (i) intercept is restricted to the co-integration space, (ii) intercept in the short-run model (which corresponds to a model with drift) and (iii) linear trend in the co-integration vector (co-integrating relationship includes time as trend-stationary variable).[10] *Johansen* (1992b) suggests testing the joint hypothesis of both rank order and deterministic components. These three tests are conducted in turn. The strategy here is to move from the most restrictive model (i) to the less restrictive one (iii). At each stage the test statistics are compared to its critical values. These tests are conducted as long as the null hypothesis is rejected. For each country we conducted λ_{max} as well as trace tests for each national retail price with respect to the composite price.[11] The relevant test results are reported in Table 3 (with r as the number of co-integrating vectors).

Table 3: Test of co-integration (*Johansen*-test), 2 lags

H0: *r*	**(i) intercept in long-run model**	**(iii) linear trend in long-run model**

[10] Of course, no deterministic component either in the long-run nor in the short-run model is possible. But this is unlikely to occur in practise, because the intercept is needed to account for the units of measurement of the variables.

[11] All test results are presented in the Appendix (Table A).

	H0: *r*	Critical values	France	Germany	critical values	USA
λ_{max}	0	11.44	18.00	11.99	14.90	27.26
	1	3.84	***3.41***	***2.74***	8.18	***6.19***
trace	0	12.53	21.41	14.73	17.95	33.46
	1	3.84	***3.41***	***2.74***	8.18	***6.19***

According to the tests all countries have one co-integrating vector. While in France and Germany the EC model contains an intercept in the long-run model, in the US the EC model includes an intercept as well as a drift term. The fact that retail prices in the three countries are co-integrated with international prices rules out the existence of long-run PTA. As a result, asymmetric transmission can only take place in the short-run, as prices adjustment towards the long-run equilibrium.

4.4 Tests of equal coefficients and test of exogeneity.

First, we estimate the equation (3a) and (3b) using *Zellner*'s (1962) Seemingly Unrelated Regressions (SUR). The error terms in the system of equations are assumed to be not independent. The disturbances of the different linear equations are correlated. Due to the non-diagonality of the error covariance matrix SUR takes into account the correlations across equations and therefore improves inference. Thus, in the presence of cross-equation correlation the SUR estimator is more efficient than individual regression estimates (Zellner 1962). After the initial regression we conduct two different tests: one of equal coefficients concerning the segmented error correction-vectors ECT_{t-1}^{+} and ECT_{t-1}^{-} and the other of weak exogeneity of the co-integrating equation. Results from these tests indicate that modifications to the system (3a) and (3b) are required.

The first test is conducted to examine whether the estimated coefficient of the positive deviation-vector (ECT_{t-1}^{+}) is equal to it negative counterpart (ECT_{t-1}^{-}). Rejecting the Null gives evidence that adjustments towards equilibrium are asymmetric and dependent of whether deviations from the equilibrium are positive or negative (Table 4).

Table 4: Tests of equal coefficients of long-run asymmetry and weak exogeneity

	$\chi^2(1)$ Critical value at 5%	France	Germany	United States
Test of equal coefficients (H_0: $\overline{\alpha}^{+} = \overline{\alpha}^{-}$)	3.84	2.61	3.29	3.26
Tests of exogeneity (H_0: co-integrating vector has no influence on endogenous)				
Retail price as endogenous variable (3a)	3.84	27.94***	26.30***	17.15***
Composite price as endogenous variable (3b)	3.84	2.04	0.25	4.06

*** Significant at the 1% level; ** significant at the given 5% level.

The null hypothesis cannot be rejected for all countries. This means that asymmetry could take place in the short-run dynamics only, (i.e. asymmetry in the first-differences variables). Additionally, Table 4 suggests that the segmented EC term must not be employed in the model. Instead, the un-segmented co-integration relationship should be utilized.

Table 4 also presents tests of weak exogeneity in a bivariate ECM - equations (3a) and (3b). The results indicate that the composite price is weak exogenous in the bivariate model for France and Germany, but not in the US between the American retail price and the composite price subsist feedback connections. Weak exogeneity of one long-run parameter implies that deviations from the equilibrium causes price adjustments in just one market. If the long-run parameters in both equations are not weak exogenous then there exist feedback relationships between the retail and the composite price. One expects at least one price to be weak exogenous because this price gives the impulse to correct deviations from the long-run equilibrium.

The ECM can be estimated in different ways. *Engle* and *Granger* (1987) suggest a two-stage method based on the asymptotical independence of the co-integrating relationship and the short-run dynamics.[12] This method is appropriate if the long-run relationship shows signs of segmentation into positive and negative deviations and especially in large samples. Another way is the one-stage OLS-estimation of the EC model. It produces *t*-values for the short-run parameters and for each parameter in the co-integrating vector.[13] Simulations have shown that one-stage OLS-estimation is preferred in the case of small samples. We use the latter procedure because tests show that the long-run relationship should not be segmented and the sample is relatively small. It follows that we use the lagged variables instead of the EC term. Making use of results in Table 4 we can modify equations (3a) and (3b) into the relevant estimation model yielding:[14]

$$\begin{aligned}(4a)\qquad \Delta RP_t^i &= \alpha_0 + \overline{\alpha} RP_{t-1}^i + \overline{\overline{\alpha}} CP_{t-1} \\ &- \alpha_2^+ \Delta^+ RP_{t-1}^i - \alpha_2^- \Delta^- RP_{t-1}^i + \alpha_3^+ \Delta^+ CP_t + \alpha_3^- \Delta^- CP_t - \alpha_5^+ \Delta^+ CP_{t-1} - \alpha_5^- \Delta^- CP_{t-1} \\ &- \alpha_6^+ \Delta^+ z_t + \alpha_6^- \Delta^- z_t - \alpha_7^+ \Delta^+ z_{t-1} - \alpha_7^- \Delta^- z_{t-1} + \varepsilon_{1t}\end{aligned}$$

$$\begin{aligned}(4b)\qquad \Delta CP_t &= \beta_0 + \overline{\beta} CP_{t-1} + \overline{\overline{\beta}} RP_{t-1}^i \\ &- \beta_2^+ \Delta^+ CP_{t-1} - \beta_2^- \Delta^- CP_{t-1} + \beta_3^+ \Delta^+ RP_t^i + \beta_3^- \Delta^- RP_t^i - \beta_5^+ \Delta^+ RP_{t-1}^i - \beta_5^- \Delta^- RP_{t-1}^i\ . \\ &+ \beta_6^+ \Delta^+ z_t' + \beta_6^- \Delta^- z_t' - \beta_7^+ \Delta^+ z_{t-1}' - \beta_7^- \Delta^- z_{t-1}' + \varepsilon_{2t}\end{aligned}$$

4.5 Estimation results and tests of the asymmetry hypothesis.

Before discussing the empirical findings a brief discussion on the identifying variables is appropriate. Tests in the previous section indicate that identification of the short-run dynamics is crucial. Because asymmetry can only take place in the first-difference variables, statistical inference requires identification of the system's dynamics. Regarding the retail equation (4a) and in European countries, the exchange rate between the domestic currencies and the US-Dollar

[12] The moments of the integrated variables converge faster than the moments of the stationary variables.

[13] But the estimation of the ECM does not compute *t*-values for the long-run relationship itself. Because we are preliminary interested in the PTA the *t*-values of the first-differences are important.

[14] With $\overline{\alpha} = \alpha_1 + \alpha_2 - 1$, $\overline{\beta} = \beta_1 + \beta_2 - 1$ as well as $\overline{\overline{\alpha}} = \frac{\alpha_2 + \alpha_3 + \alpha_5}{\alpha_1 + \alpha_2 - 1}$ and $\overline{\overline{\beta}} = \frac{\beta_2 + \beta_3 + \beta_5}{\beta_1 + \beta_2 - 1}$.

EX_t^f, EX_t^g are used. Because the composite price is denoted in US-Dollars, increases and decreases in the exchange rate affect the national price even if the composite price stays the same (ceteris paribus). Conversely, changes in the exchange rate do not influence the world prices of green coffee. Since identification in the US retail equation needs another variable different than the exchange rate, monthly average consumer price index in the United States (CPI_t^{us}) is employed. On the other hand, the identifying restriction on the composite price equation (4b) is the monthly average precipitation in Brazil $RAIN_t$.

Table 5 presents parameter estimates of the system (4a) and (4b) utilizing SUR as well as the PTA test.

Table 5a: Estimation results, retail price equation following (4a), *t*-values in brackets

Retail price equation (4a)		**France**		**Germany**		**United States**	
Constant		0.0869	(2.9846) ***	-0.1726	(-3.7831)	-0.1676	(-2.7609)***
Trend		--		--	--	-0.0005	(-1.9858) **
RP_{t-1}^i	for $i = f, g, us$	-0.0618	(-4.8275)***	0.0555	(4.4768) ***	0.1085	(3.1198)***
CP_{t-1}		0.0798	(3.0834) ***	-0.0979	(-3.6568)***	-0.0673	(-1.1339)
$\Delta^+ RP_{t-1}^i$	for $i = f, g, us$	0.5753	(6.7644) ***	0.2919	(2.1941) **	0.0077	(0.0415)
$\Delta^- RP_{t-1}^i$	for $i = f, g, us$	-0.0726	(-0.5407)	0.4608	(4.6896) ***	0.1834	(2.7678) ***
$\Delta^+ CP_t$		0.0198	(0.3004)	0.3664	(3.0693) ***	-0.2574	(-1.3293)
$\Delta^- CP_t$		0.2738	(2.3669) **	0.0839	(1.2636)	0.3862	(3.3897) ***
$\Delta^+ CP_{t-1}$		0.1766	(2.2841) **	-0.1484	(-1.3504)	-0.6495	(-3.6312) ***
$\Delta^- CP_{t-1}$		-0.3516	(-3.2099) ***	0.1373	(1.7566)	1.1601	(8.0518) ***
$\Delta^+ EX_t^{f,g}$	resp. $\Delta^+ CPI_t^{us}$	-0.5333	(-8.8540) ***	2.7904	(13.574) ***	-0.6968	(-0.2444)
$\Delta^- EX_t^{f,g}$	resp. $\Delta^- CPI_t^{us}$	-0.3621	(-5.7455)	2.0617	(9.7189) ***	18.9459	(0.6566)
$\Delta^+ EX_{t-1}^{f,g}$	resp. $\Delta^+ CPI_{t-1}^{us}$	-0.0181	(-0.2189) ***	-0.4846	(-1.1945)	-4.6025	(-1.6216)
$\Delta^- EX_{t-1}^{f,g}$	resp. $\Delta^- CPI_{t-1}^{us}$	0.1174	(1.5751)	-0.7195	(-2.2094) **	-68.9420	(-2.3745) **
$\overline{R}^2$ (adjusted)		0.7849		0.8740		0.7310	
Durbin-Watson statistic		2.1438		1.9714		1.8722	

*** Significant at the 1% level; ** significant at the 5% level.

Table 5b: Estimation results, composite price equation following (4b), *t*-values in brackets

Composite price equation (4b)	**France**		**Germany**		**United States**	
Constant	-0.0205	(-1.3632)	0.0269	(1.5141)	-0.1671	(-2.6374) ***

Trend		--	--	--	--	-0.0003	(-1.2765)
RP_{t-1}^{i}	for $i = f, g, us$	--	--	--	--	0.0815	(2.3340) **
CP_{t-1}		--	--	--	--	-0.0416	(-0.7555)
$\Delta^{+}CP_{t-1}$		-0.0329	(-0.2770)	0.2276	(1.3991)	0.1201	(0.6300)
$\Delta^{-}CP_{t-1}$		0.3166	(1.9700)**	0.0068	(0.0558)	-0.3474	(-1.9812) **
$\Delta^{+}RP_{t}^{i}$	for $i = f, g, us$	0.6504	(4.3431) ***	-0.0958	(-0.8661)	-0.2519	(-13732)
$\Delta^{-}RP_{t}^{i}$	for $i = f, g, us$	-0.1241	(-0.8032)	0.4483	(3.9390) ***	0.2177	(2.1096) **
$\Delta^{+}RP_{t-1}^{i}$	for $i = f, g, us$	-0.4638	(-3.2370) ***	0.0529	(0.4730)	-0.2884	(-1.6825)
$\Delta^{-}RP_{t-1}^{i}$	for $i = f, g, us$	0.0064	(0.0416)	-0.2502	(-2.1958)**	0.1973	(2.8725) ***
$\Delta^{+}RAIN_{t}$		0.0002	(1.9506)**	-0.0001	(-1.7097)	-0.0002	(-2.6259) ***
$\Delta^{-}RAIN_{t}$		-0.0002	(-2.6211)***	0.0002	(2.3458)**	0.0003	(2.9826) ***
$\Delta^{+}RAIN_{t-1}$		-0.0002	(2.6764)***	-0.0002	(-2.7738)***	-0.0002	(-2.3713) **
$\Delta^{-}RAIN_{t-1}$		0.0002	(2.3818)**	-0.0002	(-1.9393)**	-0.0002	(-2.5005) **
$\overline{R}^{2}$ (adjusted)		0.3786		0.3751		0.4366	
Durbin-Watson statistic		1.8422		1.8772		1.7171	

*** Significant at the 1% level; ** significant at the 5% level.

Considering the retail equation the model explains about 78, 86 and 72 percent of the variation of changes in retail prices in France, Germany and the US, respectively. The estimated adjusted $\overline{R}^{2}$ of the composite price equation, on the other hand, is 0.38, 0.37 and 0.42, respectively. Such low explanatory power of the composite equation is because other factors different than trade (i.e., future prices in the stock market) generate speculative investments which we cannot model within this framework.[15] The *Durbin-Watson* statistics indicate no autocorrelation in the error terms. The variables of the long-run relationship are highly significant in France and Germany, as it is expected from the co-integration vector and weak exogeneity tests. In contrast, the long-run parameters in the US equation indicate that long-run equilibrium in the coffee market does not hold for this country.

One expects positive (negative) deviations of the composite price to have positive (negative) effects on retail prices. Due to adjustment lags price changes from $t-2$ to $t-1$ may

15 We assume that the price at time t contains all relevant and available information at time t.

be more important than the price changes from *t-1* to *t*. Parameter estimates show differences across countries. Considering the retail price equation, results indicate that France and the US have similar behaviour. That is, increases in composite price from period $t-2$ to $t-1$ ($\Delta^{+}CP_{t-1}$) have significant impacts on retail prices; this is not the case of contemporary positive deviations ($\Delta^{+}CP_{t}$). On the other hand, negative changes have significant influence in both time lags ($\Delta^{-}CP_{t}$ and $\Delta^{-}CP_{t-1}$). In Germany, ion the other hand, changes in composite price have much smaller effects on retail price when compared to the US and France: only positive contemporaneous changes in composite prices result in retail price changes ($\Delta^{+}CP_{t}$).

While the CPI in the US has no impact on the retail price, the exchange rate in France and Germany does. An increase (decrease) in the exchange rate between the German Mark with respect to the US-Dollar leads to an increase (decrease) in the consumer price (e.g. when green beans become more expensive (cheaper) in Germany). This is true for the first differences, but not for the second differences, indicating rapid adjustments to changes in the exchange rate. In France, instead, results indicate that a one time increase in the exchange rate reduces retail prices, which is an unexpected result. Finally, reductions in the exchange rate do not affect consumers of coffee in France.

Regarding the composite price equation, the lagged first-differences of the composite price do not have a significant impact on current price. Parameter estimates show that composite prices are influenced by increases of retail prices in all three countries and by the precipitation in Brazil. The latter result is not surprising since Brazil is world's largest coffee exporting country. Changes in weather as well as changes in harvest expectations are important determinants of the world market price.[16] Changes in retail prices in the US have significant effect on composite prices, in particular when reductions in retail price occur. Furthermore, Table 5b indicates that decreases in the retail prices result in lower composite prices.[17] The same relationship holds for Germany in the case of first-difference lagged one period, but not in the case of first differences lagged two periods. Thus, reductions in German retail prices lagged one period results in increments of composite prices at time *t*. This is a plausible reaction because lower consumer prices are likely to increase retail demand for coffee. Parameter estimates for France show that an increase in retail price from $t-1$ to t produces increases in composite prices at first. But an increase in the retail price over the period $t-2$ to $t-1$ reduces composite prices.

Before drawing conclusions about PTA it is necessary to test whether or not the segmented vectors are equal, employing χ^2-tests. Under the null hypothesis we have the hypothesis of equal coefficients. If we could reject the null asymmetry is present. The following Table 6 displays the test results.

Table 6a: Tests of asymmetric adjustment – retail price equation (4a).

Null hypothesis: $\alpha_j^+ = \alpha_j^- \forall j$	$\chi^2(1)$ **Critical value at 5%**	**France**	**Germany**	**United States**
$\Delta^+ RP_{t-1}^i$ *and* $\Delta^- RP_{t-1}^i$	3.84	58.0402***	1.6105	0.8888
$\Delta^+ CP_t$ *and* $\Delta^- CP_t$	3.84	14.8233***	5.5988**	11.0475***
$\Delta^+ CP_{t-1}$ *and* $\Delta^- CP_{t-1}$	3.84	46.6615***	6.7611***	102.3573***
$\Delta^+ z_t$ *and* $\Delta^- z_t$	3.84	8.0752***	12.5661***	47.4729***
$\Delta^+ z_{t-1}$ *and* $\Delta^- z_{t-1}$	3.84	2.6852	0.3344	513.8395***

Table 6b: Tests of asymmetric adjustment – composite price equation (4b)

Null hypothesis: $\beta_j^+ = \beta_j^- \forall j$	$\chi^2(1)$ **Critical value at 5%**	**France**	**Germany**	**United States**

[16] Because the coffee tree is very sensitively to timing and quantity of precipitation one cannot say that increase in precipitation is always good or bad. Thus, we do not concentrate on the estimated coefficients. We just need them to identify the equations.

[17] All entries in the vector of the first differences ($\Delta^- RP_t^i$ and $\Delta^- RP_{t-1}^i$) have a negative sign. The positive estimated coefficient indicates that the effect on the endogenous must be negative.

$\Delta^{+}CP_{t-1}$ *and* $\Delta^{-}CP_{t-1}$	3.84	8.6522***	1.8421	6.0130***
$\Delta^{+}RP_{t}^{i}$ *and* $\Delta^{-}RP_{t}^{i}$	3.84	26.7471***	24.2064***	6.5528***
$\Delta^{+}RP_{t-1}^{i}$ *and* $\Delta^{-}RP_{t-1}^{i}$	3.84	10.7713***	7.3385***	8.0295***
$\Delta^{+}RAIN_{t}$ *and* $\Delta^{-}RAIN_{t}$	3.84	23.4762***	18.6301***	36.3089***
$\Delta^{+}RAIN_{t-1}$ *and* $\Delta^{-}RAIN_{t-1}$	3.84	0.0086	0.6357	0.0586

*** denotes a significance level of 1% and ** a significance level of 5%.

The test indicates that asymmetric adjustment in the short-run dynamics occurs in all three countries. The first and second differences of composite price have an asymmetric influence on the retail price in the three countries while the lagged retail prices do not have. Exchange rates have asymmetric impact on the retail prices as well as the consumer price index in the US, but only first differences. Only the first differences of all national retail prices have an asymmetric effect on the composite price. The same holds for the first difference of the average precipitation.

Results from Table 5 and Table 6 combined indicate significant asymmetric effects of the differentiated composite price in France and the US. In Germany, however, the vector of first differences is statistically significant and asymmetric only in the case of positive changes. These results suggest that the price transmission in Germany differs from the one in France and the US. The German adjustment process indicates that the behaviour of composite prices is less important for the retail price determination than in the other two countries.

National exchange rates (France and Germany) deviate significantly from zero and show asymmetric adjustments, but signs of the estimated coefficients are unexpected: While a positive difference leads to an increase in the German retail price (positive sign), it leads to a decrease in the French retail price (negative sign). And the opposite is true for the negative difference: If the exchange rate French Franc to US-Dollar falls (green coffee beans become less expensive for French roasters) the retail price rises in France.

5. Conclusions

Findings of the econometric model contribute to explain differences in market structure across countries. In Germany, for example, an increase in the composite price is transmitted to retail prices instantaneously, and this is the only significant effect of composite prices on the retail prices in Germany. Other factors that affect German retail prices are rising and shrinking exchange rates which cause an increasing resp. decreasing retail price without significant effects from earlier periods. The results for France and the US are different and their interpretation is more complex. For these two countries a reduction in the composite price leads to an instantaneous (i.e. in the same period) reduction in the retail price. Changes in composite price, however, produce lagged adjustments in the retail prices. An increase in the factor price from $t-2$ to $t-1$ has as effect a growing consumer price in France. In the US the opposite is true since a decline in composite prices from $t-2$ to $t-1$ results in a reduction to the retail price at time t. These observations suggest that price adjustments in Germany tend to take place more quickly than in France and the US.

Differences in the observed asymmetries and adjustment speeds can be linked to differences in market structure and differences in price elasticities of demand.

Consider the coffee market in Germany first. High degree of competition is one important characteristic of the supply side in this country. Although the main producers are highly concentrated,[18] pressure is put on them to keep prices low in order to compete with the price leader ("*Aldi*"). This implies an upper limit for retail prices that the market leaders charge though the leading firms have large market shares (*Dobson Consulting* 1999). Competition between suppliers could be described as a price war (*Koerner* 2002).[19] On the demand side, per capita consumption level associated with a relatively high level of consumption limits pricing strategy. Thus, coffee suppliers are limited to exert market power both by retailers and by consumers.

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[18] Concentration ratio of the six biggest suppliers is 90%.

[19] Additionally, the margins of German coffee suppliers are minimal or non-existent (*Marketing in Europe* 1998).

In the US and France, in contrast, market structure shows signs of oligopoly power on the supply side. Changes in production costs are not completely passed on to the consumer within the same period. Adjustments take place over at least three periods. In France, market observers have claimed that retailers have market power against the roasters despite the fact that the six biggest producers have a combined market share of 80%. Retailers try to impose a delay on producers before they can pass the increase on in wholesale prices (*Dobson Consulting* 1999). That may be an explanation why retail prices do not always reflect changes in the underlying coffee prices and are transmitted asymmetrically.[20] Concerning demand, the French consumption level is lower than in Germany indicating a lower price elasticity of demand relative to Germany.

There is also evidence of exercise of market power by the coffee industry in the US. *Gómez and Castillo* (2001), for instance, find that changes in the formation of international prices after the elimination of the International Coffee Agreement resulted in a transfer of rents from coffee exporting countries to coffee roaster in the US. *Koerner* (2002) reports that US suppliers exercise market power and that retail prices are above marginal costs. Moreover, another study conducted by World Bank on the increasing differences between world prices and retail prices in commodity markets, found that the price of coffee decreased on world markets by 18% between 1975 and 1993, while coffee prices for consumers in the United States increased by 240% during the last 25 years (*Consumers International* 2002). This suggests is that price falls of commodities traded internationally have not been passed on to consumers.

Our analysis has shown that PTA takes place in France, Germany and the US, but not for all variables and time lags. We relate the differences in price adjustments mechanism to differences in market structure and differences in demand behaviour. While increases in the composite price are transmitted positively to the retail price in Germany, factor price losses are

[20] At the beginning of the period analysed French coffee suppliers has tended to price competition. Since the *Galland Law* (entered into force in 1998) has forbidden sales at a loss, roasters and retailers are not allowed to have negative margins. Possibilities of price war are limited in contrast to Germany (*Dobson Consulting* 1999).

passed on to consumer price in France and the US. Such differences are based on the relatively high degree of competition and the relatively low margins in Germany (compared to France and the US).

Further research should use world market prices of different varieties instead of the composite price. Due to national differences in taste, this disaggregation would give provide a more detailed insight into the highly differentiated coffee market in developed countries. In addition, tests of stability could provide information whether the estimated coefficients are stable over several periods. Such extension could only be done if the time series are longer. Finally, one of the limitations of the study is the data frequency. Concerning world market prices monthly data does not reveal adjustments that are likely to occur within a single month. It would be ideal to work with weekly of daily data, but unfortunately these data at the retail level do not exist. Retail prices, on the other hand tend to be sticky and adjustments do not take place as often as in the world market. Weekly or daily retail price data would not provide enough variation to get useful estimates.

Appendix

Table A: Tests of co-integration (with *r* as number of co-integrating vectors)**.**

Critical Values	**H0: *r***	**(i) intercept in long-run model**	**(ii) intercept in short-run model**	**(iii) linear trend in long-run model**
λ_{max}	**0**	11.44	14.07	14.90
	1	3.84	3.76	8.18
trace	**0**	12.53	15.41	17.95
	1	3.84	3.76	8.18

France	**H0: *r***	**(i) intercept in long-run model**	**(ii) intercept in short-run model**	**(iii) linear trend in long-run model**
λ_{max}	**0**	18.00	17.93	25.50
	1	***3.41***	3.35	3.72
trace	**0**	21.41	21.28	29.22
	1	***3.41***	3.35	3.72

Germany	**H0: *r***	**(i) intercept in long-run model**	**(ii) intercept in short-run model**	**(iii) linear trend in long-run model**
λ_{max}	**0**	11.99	11.83	20.72
	1	***2.74***	2.66	3.28
trace	**0**	14.73	14.49	24.00
	1	***2.74***	2.66	3.28

USA	**H0: *r***	**(i) intercept in long-run model**	**(ii) intercept in short-run model**	**(iii) linear trend in long-run model**
λ_{max}	**0**	17.23	17.23	27.26
	1	6.25	6.16	***6.19***
trace	**0**	23.48	23.39	33.46
	1	6.25	6.16	***6.19***

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