WP 2009-29
September 2009

Working Paper

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

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April, 2009

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Abstract

We use monthly data spanning the period 1990-2006 to construct error correction representation models to examine price transmission asymmetries between international coffee prices and retail coffee prices in the United States, France and Germany. We find no evidence of long-run price transmission asymmetries. However, we provide evidence of short-run asymmetries with substantial differences among countries. For example, in Germany, decreases in international prices are transmitted faster to retail prices than increases are. Conversely, in the United States increases in international prices are transmitted faster to retail prices than decreases are. In France we find only modest evidence of price transmission asymmetries. We discuss our findings in the context of the differences in supply structures among the three countries.

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

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Introduction

There is evidence in the applied economics literature of price transmission asymmetry (PTA) in supply chains for agricultural commodities. Such asymmetries have been generally explained in terms of market power as well as high cost of inventory adjustment (Meyer and von Cramon-Taubadel 2004; Peltzman 2000; Ward 1982). Various empirical studies focusing on food products find that increases in factor prices are often transmitted more quickly to end consumers than decreases in factor prices (Lass 2005; Meyer and von Cramon-Taubadel 2004; Serra and Goodwin 2003). This observed behavior is particularly relevant to the study of marketing margins in the food industry given the rapid concentration in food processing and retailing worldwide, in particular during the 1990s and early 2000s (McLaughlin 2006). Identifying the occurrence of PTAs is relevant to market practitioners in the design of international supply chain strategy. In addition, the study of PTAs is relevant to policy makers concerned about possible anti-competitive practices in global food supply chains.

PTAs may occur in downstream segments of international supply chains for roasted coffee. Figure 1 shows monthly international commodity and retail coffee prices in the three largest coffee importing countries (France, Germany and the United States) during the period 1990-2006. The Figure suggests that coffee retail prices in these countries tend to respond differently to changes in international coffee prices. For instance, the 1994 international price increase resulted in a contemporaneous increase in US retail prices. In contrast, retail prices in France and Germany increased at a slower pace that in the United States. Moreover, during the period 1999-2002 of declining international prices, retail prices in Germany decreased faster than retail prices in France and the United States.

[Figure 1 here]

We test PTAs between international and retail coffee prices in France, Germany and the United States using monthly data on for the period January/1990 to December/2006. We employ an Error Correction Model representation to measure the significance and the magnitude of these asymmetries. We find significant differences in short-run PTAs among the three countries. In Germany, decreases in international prices are transmitted faster to retail prices than increases are. In the United States, in contrast, increases in international prices are transmitted faster to retail prices than decreases are; and we find modest evidence of PTAs in France. Following Meyer and von Cramon-Taubadel (2004), we interpret our results in light of differences in coffee supply chains across the three importing countries. We contribute to the literature by considering PTAs in downstream coffee markets (between international and retail prices in importing countries) focusing on the post-International Coffee Agreement period. Testing for PTAs is important because they may affect all members of the supply chain including coffee growers in developing countries that became more integrated in the market after the elimination of the export quota system in the early 1990s.

Price Transmission Asymmetries and the Coffee Market

Interest in the study of price transmission mechanisms goes back to Keynesian economics postulates explaining the process of wage and prices adjustment over time. A number of empirical studies identified the presence of PTAs in aggregate price adjustments and led economists to develop theories explaining them (Mankiw and Romer 1991; Peltzman 2000). On the one hand, PTAs are viewed as the result of microeconomic price setting frictions such as costs associated with price adjustments as well as the staggered timing of price changes and inventory management (Levy et al. 1997). On the other hand, at a more aggregate level, PTAs are regarded as the consequence of imperfect competition, including demand externalities and coordination failures (Borenstein et al. 1997; Neumark and Sharpe 1992). These principles have been widely employed to construct testable models of PTAs in vertical and spatial price transmission for markets of agricultural commodities and food products (Ward 1982; Kinnucan and Forker 1987; Bailey and Brorsen 1989; Azzam 1999; Xia 2009).

Econometric methods employed in the study of PTAs have changed over time. Earlier empirical procedures developed by Wolffram (1971) and later improved by Houck (1977) focused on differences in responses of aggregate supply functions to positive and negative changes in prices. Many assessments of PTAs in the food system adopted these methodologies to the study of price transmission with mixed results (Kinnucan and Forker 1987; Boyd and Brorsen 1988; Appel 1992; Hansmire and Willett 1992; Zhang et al. 1995). Nevertheless, von Cramon-Taubadel (1998) points out that these studies may be biased because they disregard the time series properties of the data. Specifically, ignoring that prices at different levels of the supply chain are often co-integrated may lead to spurious regression results.

More recently, 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 behavior. The authors 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 to study PTAs in several contexts, including gasoline prices (Borenstein, Cameron and Gilbert 1997; Balke, Brown and Yücel 1998), interest rates (Frost and Bowden 1999), and consumer products (Peltzman 2000).

Von Cramon-Taubadel and Loy (1996) pioneered the application of EC models to examine PTAs in markets for agricultural commodities and challenge methods utilized to discuss price asymmetry in the international wheat markets. The advantages of EC models to investigate PTAs when price series are co-integrated are formalized later in von Cramon-Taubadel and Loy (1999). Subsequent studies employ EC models to examine PTAs primarily in markets for meats (Ben-Kaabia, Gil and Ameur 2005; Sanjuan and Gil 200l; Miller and Hayenga 2001; Goodwin and Holt 1999; von Cramon-Taubadel 1998) and dairy products (Lass, 2005; Serra and Goodwin 2003; Romain, Doyon and Frigon 2002). These studies provide evidence of short-run price asymmetries along supply chains for agricultural commodities.

Researchers have studied price transmission in the international coffee supply chains, primarily in the context of international trade policies. 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 the ICA because they saw it as an efficient way to provide assistance to developing countries, particularly during the cold war (Bohman, Jarvis, and Barichello 1996). In 1990, however, the ICA was eliminated and exporters relied on competition to maintain or gain market share in international markets.

This dramatic policy change generated a stream of studies regarding the impact of the International Coffee Agreement on coffee markets and the implications for the members of the international coffee supply chain (Bohman, Jarvis, and Barichello 1996; Buccola and McCandlish 1999; Boratav 2001) and on price transmission at various levels (Krivonos 2004; Mehta and Chavas; 2008; Fafchamps and Vargas 2008). Krivonos (2004) conducts a co-integration analysis showing that the rate of price transmission between farm and international prices increased during the post-ICA period. However, the study finds evidence of price transmission asymmetries that favor coffee exporters. Fafchamps and Vargas (2008) employ data from growers, traders and exporters in Ghana to examine price transmission from international to prices received by coffee growers. They find that traders enter the market to benefit from higher international prices without transmitting these higher prices to coffee growers. Most recently, Mehta and Chavas (2008) study the impact of the ICA on the relationship between farm prices in exporting countries, international prices, and retail prices in importing countries. Their results suggest that coffee roasters and retailers benefited from price asymmetries between international and retail prices during the ICA period.

This study extends research on price transmission in coffee markets by testing PTAs between international and retail prices in France, Germany, and the United States, the three largest coffee importing countries. In addition, we follow Meyer and von Cramon-Taubadel (2004) to discuss our findings in the context of differences in the coffee supply chains across the three countries.

An Empirical Model of Asymmetric Price Transmission

PTAs can occur in the short- and long-runs, depending on 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 long-run asymmetries yield incomplete price transmission. The differences between positive and negative changes accumulate over time leading to a non-stable long-run equilibrium. In contrast, if two time series are integrated and co-integrated, long-run PTA is inconsistent with theory and only short-run asymmetries are possible (von Cramon-Taubadel and Loy 1996). On the other hand, PTAs can occur in the short-run, as the speed of adjustment toward the long-run equilibrium depends on the sign of the price change.

To address long- and short-run asymmetries, consider a distributed lag model with two non-stationary time series (y_t and x_t) and two lags:

(1) $y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 y_{t-2} + \alpha_3 x_t + \alpha_4 x_{t-1} + a_5 x_{t-2} + \varepsilon_t$

Assuming that y_t and x_t are co-integrated and re-rearranging (1), the general model of an EC representation yields

(2) $$\Delta y_t = \alpha_0 + (\alpha_1 + \alpha_2 - 1)\left[y_{t-1} + \frac{\alpha_3 + \alpha_4 + \alpha_5}{\alpha_1 + \alpha_2 - 1} x_{t-1} \right] - \alpha_2 \Delta y_{t-1} + \alpha_3 \Delta x_t - \alpha_5 \Delta x_{t-1} + \varepsilon_t$$,

where the long-run relationship (co-integration equation) between y_t and x_t is $y_t = \beta_0 + \beta_1 x_t + u_t$. The second term in brackets on the right hand side is the error correction term (ECT) representing the deviation from the equilibrium in the previous period:

(3) $$ECT_{t-1} = \nu_{t-1} = y_{t-1} - \rho_0 - \rho_1 x_{t-1}$$

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 PTAs can take place in 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_{t-1} when the latter 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 to explore short-run asymmetries. Equation (2) can be modified into its asymmetric representation as follows:

(4) $$\Delta y_t = \alpha_0 + \bar{\alpha}^{+} ECT_{t-1}^{+} + \bar{\alpha}^{-} ECT_{t-1}^{-} - \alpha_2 \Delta\, y_{t-1} + \alpha_3^{+} \Delta^{+} x_t + \alpha_3^{-} \Delta^{-} x_t - \alpha_5^{+} \Delta^{+} x_{t-1} - \alpha_5^{-} \Delta^{-} x_{t-1} + \varepsilon_t$$

where $\bar{\alpha} = \alpha_1 + \alpha_2 - 1$. Long-run 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 $\bar{\alpha}^{+} = \bar{\alpha}^{-}$ PTA is rejected and prices adjust equally for positive and negative changes from the long-run equilibrium. The same holds for the estimated parameters of the variables expressed in differences.

Hitherto the discussion assumes an unidirectional relationship between y_t and x_t. However, it is possible that these two variables are determined simultaneously. Consequently, we conduct weak exogeneity tests to examine whether the co-integrating equation influences both variables. Identification of the short-run dynamics in our model needs at least one restriction on each equation. A simultaneous representation of equations yields

(5a) $\Delta y_t = \alpha_0 + \bar{\alpha}^{+} ECT_{t-1} + \bar{\alpha}^{-} ECT_{t-1} - \alpha_2 \Delta y_{t-1} + \alpha_3^{+} \Delta^{+} x_t + \alpha_3^{-} \Delta^{-} x_t - \alpha_5^{+} \Delta^{+} x_{t-1} - \alpha_5^{-} \Delta^{-} x_{t-1} + \alpha_6 \Delta z_t - \alpha_7 \Delta z_{t-1} + \varepsilon_{1t}$

(5b) $\Delta x_t = \beta_0 + \bar{\beta}^{+} ECT_{t-1} + \bar{\beta}^{-} ECT_{t-1} - \beta_2 \Delta x_{t-1} + \beta_3^{+} \Delta^{+} y_t + \beta_3^{-} \Delta^{-} y_t - \beta_5^{+} \Delta^{+} y_{t-1} - \beta_5^{-} \Delta^{-} y_{t-1} + \beta_6 \Delta z'_t - \beta_7 \Delta z'_{t-1} + \varepsilon_{2t}$

where Δz_t and $\Delta z'_t$ are the identifying variables for the short-run parameters. We employ the system of equations (5a-b) to examine long- and short-run asymmetries between international and retail price transmission asymmetries in France, Germany and the United States.

Data

We employ monthly data on international coffee prices and retail coffee prices in France, Germany and the United States during the period January/1990 to December/2006. 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 in US dollars per pound and international prices are a composite from different coffee varieties, expressed in US-Dollars.[1] We use monthly exchange rates of the Franc and the German Mark to the US dollar from the Federal Reserve Bank (2010) as well as the as the Import Price Index in the United States from the Bureau of Labor Statistics (2010) to identify the retail price equations. We apply the conversion factor between the Franc, the German Mark and the Euro after adoption of the common currency in January/2002.[2] We use the monthly average precipitation in Fortaleza, Brazil to identify the short run dynamics of the international price equation because weather patterns affect international prices (National Centre for Atmospheric Research 2010). We provide descriptive statistics of these data in Table 1.

[Table 1 here]

Tests of Integration, Co-integration and Weak Exogeneity

Integration - Most tests of integration assume non-stationarity under the null hypothesis and often fail its rejection. The Augmented Dickey-Fuller (ADF) and the Phillips-Perron tests are examples of this approach. However, simulations have shown that in small samples both tests show lower diagnostic power than the DF-GLS-test (Elliott, Rothenberg, and Stock 1996; Elliott 1999).Therefore, we test for stationarity under the null and under the alternative hypothesis. The most commonly used test under the null of stationarity is the Lagrange-Multiplier-test of Kwiatowski et al. (1992), known as the KPSS-test.

We construct ADF and DF-GLS tests with non-stationarity under the null hypothesis and KPSS tests with stationarity under the null hypothesis. Test results in Table 2 are robust to the alternative specifications as well as to deterministic processes (i.e. deterministic trends and constants). Our results suggest that all retail price series as well as the international price series contain unit roots with or without constant and trend. However, the null hypotheses for the price series in first differences are rejected (not rejected in the case of the KPSS test) indicating that all time series are *I(1)* without deterministic trends.

[Table 2 here]

Co-integration - Johansen (1992a, 1992b, 1995) as well as Johansen and Juselius (1992) proposed tests to determine whether two *I(1)* time series are co-integrated. The procedures identify the number of equations that determine the co-integration relationship between the two series. The tests are based on the matrix of canonical correlations. One method is the *trace* test (Johansen 1988), which is a likelihood ratio test defined by $trace = -T \sum_{i=r+1}^{n} \log\left(1-\hat{\lambda}_i\right)$, where T is the number of observations, r is the number of co-integration relations and $\hat{\lambda}_i$ is the eigenvalue. The principle is to determine how many eigenvalues equal one and the test is carried out until the null hypothesis cannot be rejected. The second approach, the λ_{max} test, addresses the significance of the estimated eigenvalues , where $\lambda_{\max} = -T \log\left(1-\hat{\lambda}_i\right)$. Critical values for this test are reported in Osterwald-Lenum (1992).

Tests of co-integration are sensitive to the structure of the data generating process - the underlying deterministic process such as constant and trend. Johansen and Juselius (1990) and Osterwald-Lenum (1992) consider three possible cases: (i) intercept 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 (i.e., the co-integrating relationship includes time as trend-stationary variable). Johansen (1992b) suggests testing the joint hypothesis of both rank order and deterministic components. Consequently, our strategy is to move from the most restrictive model (i) to the least restrictive model (iii). At each stage the test statistics are compared to their 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 international price. These results are reported in Table 3, where r is the number of co-integrating vectors.

[Table 3 here]

According to the tests, all countries have one co-integrating vector. The tests also indicate that the model should include an intercept in the error correction term in France and Germany. In contrast, the tests indicate that in the United States the error correction term should include an intercept and a linear trend. The fact that retail prices in the three countries are co-integrated with international prices rules out the existence of long-run PTAs. As a result, asymmetric transmission can only take place in the short-run, as prices adjust towards the long-run equilibrium.

Weak Exogeneity and Long-run Price Transmission Asymmetry - First, we estimate the equation (5a) and (5b) using Zellner's (1962) Seemingly Unrelated Regressions (SUR) because the error terms in the system are likely to be correlated. For the France and Germany equations (5a) we create a dummy variable that equals 1 during the Euro period and zero otherwise. We employ this specification to test for long-run asymmetry in the error correction term and for weak exogeneity in the price series (Table 4). We first examine whether the magnitude of the estimated coefficient of the positive deviation-vector (ECT_{t-1}^{+}) equals its negative counterpart (ECT_{t-1}^{-}). Table 4 suggests that the null hypothesis (symmetry) cannot be rejected in any country. This means that asymmetry can take place only in the short-run dynamics of the price relationship (i.e. asymmetry in the first-differences variables).

[Table 4 here]

We present the weak exogeneity tests corresponding to the bivariate ECM in equations (5a) and (5b) in Table 4. Test results indicate that the international price is weak exogenous in the bivariate model for France and the United States, but not for Germany. In France and the United States, weak exogeneity of the international price implies that deviations from the equilibrium cause price adjustments in retail prices only. In contrast, test results for Germany suggest feedback between retail and international prices.

There are several strategies to estimate the ECM. Engle and Granger (1987) suggest a two-stage method based on the asymptotic independence between the co-integrating relationship and the short-run dynamics. This method is appropriate if the long-run relationship shows asymmetries in the error correction term and is generally applied to large samples. An alternative, particularly in small samples, is to use a one-stage model in which the components of the error correction term are employed directly in the estimating equation. Based tests presented in Table 4, we modify equations (5a) and (5b) and estimate the following model for each country:

(6a) $\Delta RP_t^j = \alpha_0 + \bar{\alpha} RP_{t-1}^j + \bar{\bar{\alpha}} IP_{t-1} - \alpha_2 \Delta RP_{t-1}^j + \alpha_3^+ \Delta^+ IP_t + \alpha_3^- \Delta^- IP_t - \alpha_5^+ \Delta^+ IP_{t-1} - \alpha_5^- \Delta^- IP_{t-1} + \alpha_6 \Delta z_t - \alpha_7 \Delta z_{t-1} + \varepsilon_{1t}$

(6b) $\Delta IP_t = \beta_0 - \beta_2 \Delta IP_{t-1} + \beta_3^+ \Delta^+ RP_t^j + \beta_3^- \Delta^- RP_t^j - \beta_5^+ \Delta^+ RP_{t-1}^j - \beta_5^- \Delta^- RP_{t-1}^j + \beta_6 \Delta z_t' - \beta_7 \Delta z_{t-1}' + \varepsilon_{2t}$,

where $\bar{\alpha} = \alpha_1 + \alpha_2 - 1$, $\bar{\beta} = \beta_1 + \beta_2 - 1$; $\bar{\bar{\alpha}} = \alpha_3 + \alpha_4 + \alpha_5$; and $\bar{\bar{\beta}} = \beta_3 + \beta_4 + \beta_5$. Equation (6a) includes a trend in the error correction term in the US model; and equation (6b) includes the error correction term as explanatory variable in the German model. Statistical inference requires identification of the short run dynamics. For the retail equation, we employ the exchange rate between the domestic currency and the US dollar, EX_t^f, EX_t^g in France and Germany, respectively; and for the United States we employ the monthly import price index for food and beverage products (IPI_t^{us}). The identifying restriction on the international price equation (6b) is the monthly average precipitation in Fortaleza, Brazil ($RAIN_t$).

Results

Table 5 presents Seemingly Unrelated Regression (SUR) parameter estimates of the system (6a) and (6b) for each country. The retail price equations explain about 77, 60 and 58 percent of the variation in retail prices in France, Germany and the United States, respectively. Similarly, the international price equations explain 18, 19 and 15 percent of the variability in international coffee prices. The relatively lower explanatory power of the international price models may be due to the fact that factors other than trade (e.g., future prices in the stock market) generate speculative investments which we cannot model within this framework. Durbin-Watson statistics indicate no autocorrelation in the error terms. Our discussion below focuses primarily on the retail price equations, given that our objective is to examine asymmetries in price transmission from international to retail prices.

[Table 5 here]

Long-run equilibrium between international and retail prices – The estimated coefficient of IP_{t-1} describes the long run relationship between international and retail prices and the estimated coefficient of RP_{t-1} indicates the speed of adjustment towards the long-run equilibrium following a change in international prices. The parameters estimates of IP_{t-1} are positive in all three countries, as predicted by theory, although the United States coefficient is statistically insignificant. In Germany (France), a $1 increase in international coffee price leads to a $0.14 ($0.08) in retail price; but this adjustment takes place at a rate on 0.039 (0.043) per month. In the United States, international prices may have only short-term effects on retail prices and these effects do not persist in the future; and the trend coefficient suggests that the price spread between international and retail price increased at a modest significant rate of $0.0002 per pound per month during the period of analysis. These results suggest differences between the three countries: the long-run relationship between international and retail prices is stronger in Germany than in France, yet the speed of adjustment is similar in these two countries. In the United Sates, our results do not provide evidence of a long-run equilibrium between international and retail prices.

Short-run asymmetries between international and retail prices – In Table 6, we present tests results for short-run asymmetries regarding the impact of contemporaneous and lagged changes in international prices (ΔIP_t and ΔIP_{t-1}) on changes in retail prices (ΔRP_t). Test results suggest differences in short-run dynamics across countries. In Germany, there is evidence that negative changes in international prices have a larger effect on retail prices than positive changes:

a $1 decrease (increase) in international price is associated with a $0.68 ($0.23) contemporaneous decrease (increase) in retail prices. Asymmetry tests in Table 6 suggest that negative changes have significantly larger impacts than their positive counterparts. Lagged changes in international prices in the previous month, either positive or negative, do not affect current changes in retail prices. Our German results are in sharp contrast with parameter estimates for the United States, in which positive changes in international prices appear to have a greater effect on retail prices than do negative changes. Specifically, for the United States, our results suggest that while a $1 increase in international price leads to a $0.45 contemporaneous increase in retail prices, negative changes in international prices do not affect retail prices. Moreover, a $1 increase in lagged international prices is associated with a $1.12 increase in retail prices; and, contrary to expectations, a $1 decrease leads to a $0.78 increase in retail prices. These results provide evidence that in the United States changes in retail prices are much more sensitive to positive than to negative changes in international prices (Table 6).

[Table 6 here]

Results in Table 5 and Table 6 suggest further differences in the French coffee supply chain in comparison to Germany and the United States. In France, our results indicate asymmetries on the lagged changes in international prices (ΔIP_{t-1}): a $1 increase in lagged changes international prices leads to a $0.23 increase in retail prices while a $1 decrease does not result in lower retail prices. In fact, of the coefficient of negative changes is unexpected (0.17) because it suggests that negative changes in international prices lead to positive changes in retail prices. Nevertheless, the segmented coefficients of contemporaneous changes in international prices (ΔIP_t) correct this apparent inconsistency: a $1 negative contemporaneous change in international coffee prices results in a $0.23 decline in coffee retail prices, whereas positive contemporaneous changes in international prices do not influence changes in retail prices.

The variables employed for identification of short-run dynamics are significant in France and Germany but not in the United States. As expected, changes in the exchange rate are negative and significant given that retail prices are converted into US dollars. There are modest differences in Germany during the common currency period, as reflected by the interaction coefficient $D\Delta z_t$. In the United States, the price index of imported food and beverages is used for identification and its coefficient is positive but statistically insignificant.

Short-run dynamics of the international price equation – The parameter estimates suggest that international prices are influenced by increases of retail prices in all three countries. If retail prices were to increase $1 in each importing country then the international price would increase $0.48, $0.40 and $0.13 in France, Germany and the United States, respectively. In contrast our results suggest that negative changes in retail prices do not have an effect on international prices in the three countries. Although Table 4 suggests feedback effects from retail to international prices in Germany, the estimated coefficient of lagged retail price, which represents the long-term effect that retail prices have on international prices, is statistically insignificant. Consequently, our results suggest that such effects take place only in the short-run. Lagged changes in precipitation levels in Fortaleza-Brazil, the variable employed for identification, are positive and significant in the three models. This suggests that short run weather patterns, as well as changes in harvest expectations, are important determinants of international prices.

Summary of findings – Our findings reject the hypothesis of long-run asymmetries in price transmission between international and retail coffee prices. In contrast, we find asymmetric price behaviour in the short run with marked differences across the three countries. In Germany, reductions in international prices produce faster adjustments of retail prices than do increases in international prices. In contrast, in the United States, positive changes in international prices produce immediate increases in retail prices and negative changes do not affect retail prices in the short-run. In France, our results suggest modest evidence of price transmission asymmetries: contemporaneous and lagged changes in international prices exhibit asymmetries of comparable magnitudes in opposite directions.

Short Run Price Asymmetries and Market Structure

The observed differences in short-run price transmission behavior can be discussed in the context of differences in coffee supply chains among the three importing countries. In Table 7 we present selected characteristics of the coffee supply chain in each country relevant to our period of analysis. The United States coffee market is the largest, even though the US per capita consumption is substantially smaller than in France and Germany. The coffee processing sector is slightly more concentrated in the United States than in France and Germany. The share of private label coffee brands in France and Germany (18.8 and 22.0 percent, respectively) is substantially larger than in the United States (7.8 percent). The degree of concentration of food retailing in the European countries is substantially higher than in the United States; and the primary difference between the food retailing sectors in France and Germany is the high market share of hard discounters (e.g. Aldi, Lidl) in the latter (7.8 and 34.0 percent, in France and Germany, respectively). In the US, on the other hand, the share of hard discounters was less than 2 percent during the period of analysis. Hard discounters offer limited assortment of products (typically five to six thousand stock keeping units, which is small relative to the forty-five thousand stock keeping units offered by traditional supermarkets) in large quantities, which allow them to operate extremely low-cost supply chains.

[Table 7 here]

We argue that country differences in Table 7 can be discussed in the context of PTAs identified in the econometric model. In Germany, for example, the large market share of hard-discount retailers, as well as the large market share of private label coffee brands, may explain that reductions in international coffee prices are transmitted faster than are price increases. Hard discounters often employ aggressive competitive strategies based on low prices relative to competitors. Large market share of private label brands increases the ability of food retailers to control their pricing strategies. Indeed, a number of academic and industry studies document price wars in the German retail sector in general and in the coffee product category in particular, mostly during the late 1990s and early 2000s (e.g. Koerner 2002; McLaughlin 2006). The Aldi coffee brand is the market leader in Germany, the company owns coffee roasting plants and buys green coffee directly from international commodity exchanges. Therefore, Aldi has the ability to control the supply chain and to pass lower international prices on to the end consumer.

In France, both the market concentration at the processing and retail levels, as well as the share of private label brands in the coffee category are comparable to Germany. However, the market share of hard-discount retailers in France is substantially smaller than in Germany. Furthermore, a unique feature of the French market is the role of public policies in regulating the pricing behavior along the food supply chain. A report by Dobson Consulting (1999), for example, states that the French coffee market was heading to a price war in the early 1990s, similar to its German counterpart. Nevertheless, price promotions were restricted substantially after the Government passed the Galland Law in 1996. This law is intended to avoid conflicts and imbalances in the relationship between large retailers and their suppliers as well as with small retailers. The law prevents processors and retailers from selling at a loss and retailers cannot reduce prices to take advantage of volume discounts and other promotions offered by coffee processors.[3] This regulation, together with the smaller participation of hard discounters in the French market and the similar market concentration between processors and retailers, may explain the modest evidence of price transmission asymmetries in France.

In the United States, the coffee supply chain exhibits considerable differences with respect to its European counterparts. Consider the following unique characteristics of the supply chain in this country: 1) higher concentration in the coffee processing sector; 2) moderate concentration in food retailing; 3) small share of private label brands in the coffee product category; and 4) less than two percent market share of hard-discount retailers. In addition, US government regulation regarding price promotion is less strict than in France. Therefore, coffee processors in the United States may have more ability to coordinate the supply chain than do their European counterparts. Our econometric estimates show that negative changes in international prices are not passed on to consumers as fast as are positive changes, suggesting a certain degree of oligopoly power of coffee processors. This conjecture, however, should be interpreted with caution because a formal analysis of market power is beyond the scope of the study.

Conclusion

Price transmission asymmetries can provide valuable information for private and public decision makers about supply chain behavior. We develop error correction models to statistically test for long- and short-run PTAs in France, Germany and the United States, during the post International Coffee Agreement period (1990-2006). The analysis focuses on the impact of changes in international coffee prices on retail prices and also on the links between PTA econometric estimates and coffee supply chain structures.

Our analysis provides evidence of asymmetric price transmission behavior only in the short-run with important differences between Germany, France and the United States. In Germany, negative changes in international prices have higher impacts on retail prices than do positive changes. Large share of hard-discount retailers may drive this asymmetric behavior. Price transmission behavior is opposite in the United States: positive changes in international prices produce immediate positive changes in retail prices while negative changes do not affect retail prices. The characteristics of the coffee supply chain may allow coffee processors to obtain economic rents in the short-run. Finally, we find modest evidence that asymmetric price transmission behavior may be due to public policies aimed at regulating relationships among supply chain members.

While our study provides insights regarding PTAs and market structures in coffee importing countries, several areas call for further research. Future research should take into consideration differences in consumer preferences across countries, primarily between *robusta* and *arabica* variety types. Such level of disaggregation would provide more precise estimates of price transmission asymmetries given the high level of product differentiation in the coffee product category in high income countries. Future research should also explore alternative methods such as threshold vector error correction models to assess price transmission asymmetries. Finally, more research on formal models to assess market structure and conduct is required to assess the welfare implications of the elimination of the International Coffee Agreement.

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Table 1: Descriptive statistics of the estimating sample

	Mean	Ste. Dev	Max	Min
International price	0.829	0.340	2.024	0.412
Retail price in France	2.703	0.523	4.179	1.904
Retail price in Germany	4.115	0.897	6.179	2.473
Retail price in the US	3.217	0.528	4.669	2.352
Exchange Rate (Franc/US Dollar)	5.799	0.731	7.694	4.831
Exchange Rate (Mark/US Dollar)	1.718	0.225	2.294	1.381
Import Price Index, Foods, Feeds, and Beverages[a]	1.026	0.079	1.226	0.885
Precipitation (100mm)	1.292	1.508	6.680	0

[a] Index 2000 = 1

Table 2: Tests of integration in levels and in first differences

Variables in Levels			Critical Value[a]	Retail Price France	Retail Price Germany	Retail Price US	International Price
ADF-t	H_0 : ~$I(1)$		-2.88	-1.83	-1.47	-2.59	-2.49
	H_0 : ~$I(1)$	no constant	-1.95	0.001	-0.25	-0.317	-0.71
DF-GLS	H_0 : ~$I(1)$		-2.93	-1.82	-1.46	-2.43	-2.36
	H_0 : ~$I(1)$	no linear trend	-2.03	-1.83	-1.33	-2.40	-2.25
KPSS	H_0 : ~$I(0)$	no constant	1.66	13.73	14.49	13.54	12.83
	H_0 : ~$I(0)$	no linear trend	0.463	0.469	1.54	0.5	0.56
Variables in First Differences			**Critical Value**	**Δ Retail Price France**	**Δ Retail Price in Germany**	**Δ Retail Price in US**	**Δ Internat. Price**
ADF-t	H_0 : ~$I(1)$		-2.88	-8.33	-9.67	-9.42	-12.11
	H_0 : ~$I(1)$	no constant	-1.95	-8.35	-9.70	-9.45	-12.13
DF-GLS	H_0 : ~$I(1)$		-2.93	-5.35	-7.11	-6.28	-6.64
	H_0 : ~$I(1)$	no linear trend	-2.03	-4.15	-6.92	-5.87	-6.59
KPSS	H_0 : ~$I(0)$	no constant	1.66	0.09	0.15	0.06	0.06
	H_0 : ~$I(0)$	no linear trend	0.463	0.12	0.14	0.05	0.07

[a] At the 10% level of significance.

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

Critical Value	H_0:r	intercept in long-run model	intercept in short-run model	linear trend in long-run model
λmax	0	11.44	14.07	19.67
	1	3.84	3.76	9.24
trace	0	12.53	15.41	19.96
	1	3.84	3.76	9.42

France	H_0:r	intercept in long-run model	intercept in short-run model	linear trend in long-run model
λmax	0	13.680	19.528	19.574
	1	0.004	3.704	3.788
trace	0	13.685	23.232	23.361
	1	0.004	3.704	3.788

Germany	H_0:r	intercept in long-run model	intercept in short-run model	linear trend in long-run model
λmax	0	12.542	15.289	15.319
	1	0.039	2.658	2.695
trace	0	12.581	17.937	18.014
	1	0.039	2.648	2.695

United States	H_0:r	intercept in long-run model	intercept in short-run model	linear trend in long-run model
λmax	0	10.652	25.444	25.446
	1	0.135	8.944	9.031
trace	0	10.787	34.387	34.477
	1	0.135	8.944	9.041

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

	$\chi^2(1)$ Critical value at 5%	**France**	**Germany**	**United States**
Long-run Asymmetry Test ($H_0: \bar{\alpha}^+ = \bar{\alpha}^-$)	3.84	0.00	0.02	0.68
Weak Exogeneity Test (H_0: co-integrating vector has no influence on endogenous variable)				
Retail price as endogenous variable (5a)	3.84	13.71***	9.59***	17.00***
International price as endogenous variable (5b)	3.84	3.08	10.79***	0.22

Table 5: Estimation results, Standard Errors in brackets

Retail price equation (6a)	**France**	**Germany**	**U.S.**
Constant	0.047** (0.018)	0.043 (0.031)	0.181*** (0.050)
Trend	-	-	0.0002* (0.0001)
RP^i_{t-1}	-0.043*** (0.010)	-0.039*** (0.011)	-0.094*** (0.024)
IP_{t-1}	0.078*** (0.021)	0.142*** (0.038)	0.044 (0.041)
ΔRP^i_{t-1}	0.411*** (0.059)	0.174*** (0.066)	0.123** (0.051)
$\Delta^+ IP_t$	0.038 (0.057)	0.226** (0.109)	0.445*** (0.106)
$\Delta^- IP_t$	0.231** (0.099)	0.681*** (0.192)	-0.180 (0.181)
$\Delta^+ IP_{t-1}$	0.174*** (0.066)	0.109 (0.124)	1.120*** (0.126)
$\Delta^- IP_{t-1}$	-0.173* (0.092)	-0.286 (0.180)	-0.708*** (0.174)
Δz_t	-0.433*** (0.025)	-1.996*** (0.164)	0.261 (1.471)
Δz_{t-1}	0.148*** (0.037)	-0.021 (0.218)	0.943 (1.479)
$D \cdot \Delta z_t$	0.003 (0.023)	-0.275* (0.154)	-
$D \cdot \Delta z_{t-1}$	0.009 (0.024)	-0.146 (0.156)	-
R^2	**0.749**	**0.573**	**0.571**
International price equation (6b)			
Constant	-0.005 (0.009)	-0.005 (0.027)	0.008 (0.013)
Trend	-	-	0.00002 (0.0001)
IP_{t-1}	-	-0.074*** (0.026)	-
RP^i_{t-1}	-	0.014 (0.009)	-
ΔIP_{t-1}	0.051** (0.066)	0.173** (0.068)	0.061 (0.082)
$\Delta^+ RP_t$	0.483*** (0.115)	0.403*** (0.081)	0.132* (0.072)
$\Delta^- RP_t$	-0.090 (0.134)	0.054 (0.068)	0.160 (0.136)
$\Delta^+ RP^i_{t-1}$	-0.361*** (0.114)	-0.127 (0.083)	0.061 (0.060)
$\Delta^- RP^i_{t-1}$	-0.022 (0.134)	-0.006 (0.067)	-0.180 (0.135)
$\Delta^+ Rain_t$	0.003 (0.004)	-0.003 (0.004)	-0.003 (0.004)
$\Delta^+ Rain_{t-1}$	0.016*** (0.004)	0.015*** (0.004)	0.016*** (0.004)
R^2	**0.148**	**0.186**	**0.103**

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

Table 6: Tests of asymmetric adjustment – Retail price equation

Null hypothesis: $\alpha_j^+ = \alpha_j^-, \forall j$	$\chi^2(1)$ Critical value, 10%	France	Germany	U.S.
$\Delta^+ IP_t$ and $\Delta^- IP_t$	3.84	2.08 (0.15)[a]	3.14 (0.08)	5.85 (0.02)
$\Delta^+ IP_{t-1}$ and $\Delta^- IP_{t-1}$	3.84	6.63 (0.01)	2.35 (0.12)	54.52 (0.00)

Tests of asymmetric adjustment – International price equation

Null hypothesis: $\beta_j^+ = \beta_j^-, \forall j$	$\chi^2(1)$ Critical value at 5%	France	Germany	U.S.
$\Delta^+ RP_t^i$ and $\Delta^- RP_t^i$	3.84	6.74 (0.01)	8.37 (0.00)	0.02 (0.88)
$\Delta^+ RP_{t-1}^i$ and $\Delta^- RP_{t-1}^i$	3.84	2.81 (0.09)	0.97 (0.32)	2.31 (0.13)

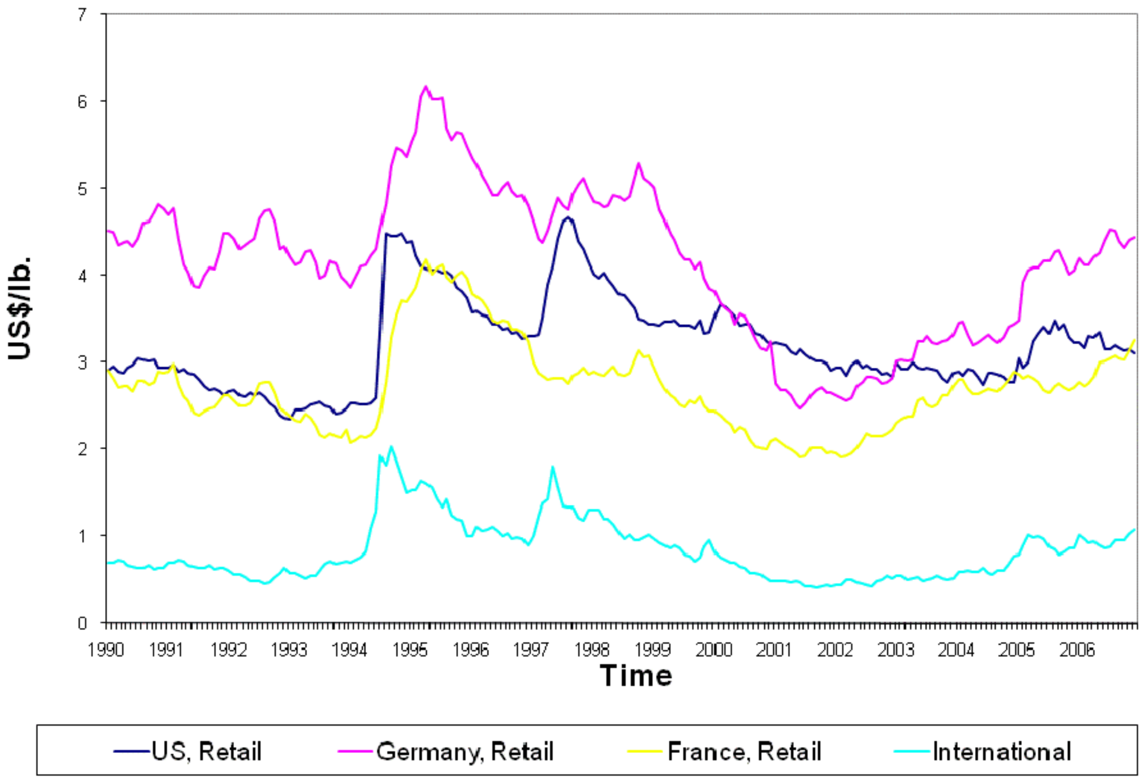
[a] Probability > *Chi* square in parenthesis.

Table 7: Selected characteristics of the coffee supply chains

	France	Germany	United States
Per Capita consumption[a]	5.42	6.23	4.18
Roasted coffee retail sales (Million US Dollars)[a]	1,039	2,297	4,145
Brand Manufacturers[b]			
Share of leading brand (%)	27.0	30.3	35.6
Share of three leading brands (%)	64.0	62.6	68.6
Share of private label brands (%)	18.0	22.0 (Aldi excluded)	7.8
Supermarket Sector			
Share of five leading supermarkets (%)[c]	76.4	61.8	35.5
Share of hard-discount retailers (%)[d]	7.8	34.0	<2.0%

[a] Averages for years 1995, 2001 and 2005, from *Tropical Products: World Markets and Trade,* Foreign Agricultural Service, United States Department of Agriculture.
[b] All figures represent averages for years 2001 and 2003, from Mintel's Market Intelligence. Private label brand share in the United States is from *Private Label* (2007) and corresponds to years 2005 and 2006.
[c] For France and Germany the figures are the average for years 2001 and 2003, from Mintel's Market Intelligence. United States figures are for years 1998-2003 the Food Industry Management Program, Cornell University.
[d] For France and Germany the figures are the average for years 2001 and 2003, from Mintel's Market Intelligence. For the United States the figure corresponds to estimates from the Food Industry Management Program at Cornell University.

Figure 1: Monthly International Coffee Prices and Retail Prices for Coffee in France, Germany and the United States: 1990-2006



Source: International Coffee Organization. International price is the mean of the weighted average of daily prices for selected coffees of the Other Mild, Arabicas and Robusta varieties, calculated by the International Coffee Organization.

Endnotes

[1] 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, 2010).

[2] 1 Euro = 1.95583 German Marks; and 1 Euro = 6.55957 French Francs.

[3] The Franch Government passed an amendment in 2005 to make the Galland Law less restrictive, but the primary principles of the law are still in in place.