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Impacts of the End of the Coffee Export Quota System on International-to-Retail Price Transmission

Jun Lee and Miguel I. Gómez

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Abstract

We examine the impact of the end of the coffee export quota system (EQS) on international-to-retail price transmission in France, Germany and the United States, taking into account the existence of long-run threshold effects and short-run price transmission asymmetries (PTAs). We find evidence of threshold effects in both periods (EQS and post-EQS) in the three countries and the presence of short-run PTAs during the post-EQS period in the three countries, but not during the EQS period. Our results indicate that the threshold values become smaller and the long-run speed of adjustment decreases during the post-EQS period in the three countries. In the short-run, retail prices in the three countries are more responsive to positive than to negative changes in international prices during the post-EQS period, providing evidence of short-run PTAs. However, changes in international prices are passed on to retail prices to a greater extent in the United States than in the two European countries. Nonlinear impulse response analyses indicate that a shock in international prices tends to persist more during the EQS period than in the post-EQS period. This suggests that price transmission increased in the post-EQS period, regardless of the direction of the change in international prices. Our results suggest that accounting for thresholds and asymmetries improves the accuracy of impact assessments of policy changes on price transmission processes.

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

JEL Classifications: *C32, Q17.*

1. Introduction

The International Coffee Agreement is an international treaty that sets out the objectives and

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the basic framework for the International Coffee Organization member countries, which includes coffee exporters and importers. An important policy instrument of this agreement was the operation of an Export Quota System (EQS) from 1962 to 1989, when the system was ended due to failure to agree on quota distribution and lack of support from importing countries. The EQS contributed to keep international coffee prices high. However, the end of the EQS reduced the influence of exporting countries, leading to decreases in international coffee prices. The average international price of a $1.30 per pound during the EQS period fell rapidly to a $0.88 per pound after the system termination (Figure 1).

[Figure 1 here]

In this study, we examine the implications of the end of the EQS on price transmission between international and retail coffee prices in the three largest coffee importers (France, Germany and the United States). Figure 1 suggests that the end of the EQS in 1990 may have affected these countries in different ways in terms of the response of retail prices to changes in international prices. For instance, retail prices in the three countries seem to have a similar relationship with international prices under the EQS period. In contrast, in the post-EQS period, after the sharp decrease in international prices in the early 1990s, retail prices in France and the United States decreased accordingly, while retail prices in Germany experienced high volatility and stayed relatively high.

There is an extensive literature on price transmission in supply chains for agricultural commodities. Early studies, conducted by Wolffram (1971) and Houck (1977), examined price transmission asymmetries (PTAs) measured by asymmetric supply responses to positive and negative changes in input prices. Observed usually at downstream stages of supply chain, PTAs are closely associated with numerous factors, including market structure, market power, consumer behavior and public policy. Subsequently, von Cramon-Taubadel (1998) pointed out that prices at different segments of the supply chain are often co-integrated and econometric models that ignore this property may generate spurious parameter estimates. The suggested remedy was error correction models (ECMs), which allow for short-run asymmetric price adjustments, to overcome the limitations of Wolffram's and Houck’s approaches (von Cramon-Taubadel and Loy, 1996; von Cramon-Taubadel, 1998). The standard ECM specification assumes that the dependent variable responds identically to deviations from the long-run price equilibrium regardless of the magnitude of the shock and that adjustments occur in every period (Balke and Fomby, 1997). However the presence of any transaction costs between spatially separated markets or other factors generating price friction may result in the existence of thresholds which trigger adjustments toward the long-run equilibrium in response to exogenous shocks (Meyer, 2004).

For the purpose here, we focus on two separate dimensions of price transmission in order to understand their relevance for policy impact evaluation. The first dimension posits that price adjustments respond differently to positive and negative exogenous shocks in the short run (i.e. PTAs). The second dimension is that there may be thresholds beyond which nonlinear price adjustments occur in the long-run (i.e. nonlinear price adjustments). Our approach involves an error correction model with threshold effects, patterned after work by Tong (1983) and extended by Balke and Fomby (1997). The threshold approach allows us to model nonlinear price adjustments toward the long-run equilibrium through threshold values estimated using nonparametric methods. In addition, we extend the threshold error correction model by incorporating short-run PTAs.

The overall objective of this study is to examine the implications of the end of the coffee EQS on international-to-retail coffee price transmission in France, Germany and the United States by taking the existence of long-run threshold effects and short-run asymmetries into account. We show that accounting for these two features of the price transmission process may lead to more accurate evaluations of the EQS termination. We contribute to the literature by improving our understanding of the impact of policy interventions on price transmission in global supply chains of agricultural commodities.

A number of researchers have examined the impact of the EQS termination at various segments of the coffee supply chain. Akiyama and Varangis (1990) employed simulation methods for the global coffee supply chain to demonstrate that the EQS contributed to the stability of international coffee prices. Krivonos (2004) conducted a co-integration analysis and found that the speed of price transmission between producer and international prices increased in the post-EQS period. She found that the share of retail value going to coffee growers increased after the EQS termination; and the analysis also showed that export country retail prices adjusted faster in response to shocks in international prices during the post-EQS period. Shepherd (2004) examined the impact of the end of the EQS on price transmission from producer to international prices; and from international to retail prices employing a vector autoregression (VAR) model. He argued that the EQS termination did not lead to improved price transmission due to market power exerted by coffee processors. Moreover, the study identified asymmetries in price transmission at all levels of the supply chain, particularly during the post-EQS period. Gemech and Struthers (2007) examined impacts of market reforms in Ethiopia on the volatility of coffee producer price using the Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model. They found that coffee price volatility increased after market reforms, triggered by the end of the EQS. Mehta and Chavas (2008) studied impacts of the EQS termination on the relationship between producer prices in exporting countries, international prices, and retail prices in importing countries. They found that, in the short-run, the extent to which retail prices respond to international prices was greater for increases in international prices than for decreases in international prices during the post-EQS period. In contrast, they found no evidence of price transmission asymmetries between producer and international prices. More recently, Gómez, Lee and Körner (2010) examined price transmission from international to retail coffee prices by employing an asymmetric error correction model (AECM) and found evidence of short-run asymmetries with substantial differences among importing countries.

This study extends and refines this literature by accounting for long-run threshold effects and short-run asymmetries in error correction models. This approach adds precision to impact evaluations for private and public decision-makers who are concerned with policies that influence the international trade of agricultural commodities.

This paper is organized as follows. We first review the literature that incorporates the threshold approach into the study of price transmission in agricultural commodity markets. Next, we develop an asymmetric threshold error correction model (ATECM) to examine impacts of the coffee EQS termination. In turn, we describe our data and present the empirical results. Finally we summarize our findings and discuss the benefits and limitations of ATECM representation to assess policies affecting price transmission in global commodity supply chains.

2. Modeling threshold co-integration in price transmission

A number of studies have utilized the threshold approach to examine price transmission in supply chains of agricultural commodities. Goodwin and Holt (1999) employed a threshold error correction model (TECM) to evaluate linkages between producer, wholesale, and retail prices in U.S. beef markets. Subsequently, Goodwin and Piggott (2001) developed a TECM to examine market integration among four corn and soybean markets in North Carolina accounting for transaction costs. More recently, Abdulai (2002) employed the threshold co-integration model developed by Enders and Granger (1998) to analyze price transmission between producer and retail prices in the Swiss pork supply chain. He compared a standard ECM with a TECM using the Akaike and Schwarz information criteria and showed that the threshold representation is superior to its standard counterpart. Meyer (2004) considered transaction costs potentially occurring in separated markets during the price transmission process and employed a vector error correction model following the procedures of Balke and Fomby (1997). He examined market integration between pig markets in Germany and the Netherlands and found the existence of significant transaction costs. Overall, these studies generally confirm the existence and relevance of thresholds between spatially separated markets and indicate that TECM representations generally show a faster adjustment towards the long-run equilibrium than their standard counterparts.

In this study, we follow and extend the threshold co-integration approach developed by Enders and Granger (1998) to incorporate two relevant properties in price transmission: the existence of thresholds in the co-integrating vector and the possible asymmetries in short-run price responses. As Balke and Fomby (1997) pointed out, the conventional co-integration tests of Johansen (1992a, 1992b, 1995) may be misspecified if the adjustment to the long-run equilibrium is not linear. To overcome this problem, Enders and Granger (1998) suggested an alternative to the standard augmented Dickey-Fuller (ADF) regression model.

Consider RP_t, retail coffee prices and IP_t, international coffee prices at time period *t*. Both price variables are assumed to be integrated of order one, *I(1)*. The co-integration relationship between two price series is given by:

$$RP_t - \sigma_0 - \sigma_1 IP_t = \varepsilon_t, \tag{1}$$

where the error term in equation (1), ε_t, indicates the deviation from the long-run equilibrium between the price series, RP_t and IP_t in period *t*. The threshold autoregressive (TAR) representation proposed by Enders and Granger (1998) is specified as follows:

$$\Delta\varepsilon_t = I_t\left[\rho_0^{(1)} + \rho_1^{(1)}\varepsilon_{t-1}\right] + (1 - I_t)\left[\rho_0^{(2)} + \rho_1^{(2)}\varepsilon_{t-1}\right] + \sum_{i=1}^{p-1}\gamma_i\Delta\varepsilon_{t-1} + \upsilon_t. \tag{2}$$

The Heaviside indicator function in equation (2), I_t, is defined as

$$I_t = \begin{cases} 1 \; if \; |\varepsilon_{t-d}| > \theta \\ 0 \; if \; |\varepsilon_{t-d}| \leq \theta \end{cases}, \tag{3}$$

where θ represents a threshold value beyond which an exogenous shock triggers adjustments toward the long-run equilibrium between prices, and *d* refers to a delay parameter. The Akaike Information Criteria (AIC) and the Schwartz Bayesian Criteria (SBC) are typically employed to determine the appropriate lag structure of equation (2). Price adjustments may occur only when the benefits offset the cost of changing prices due to the presence of transaction costs or other sources of price frictions (Balke and Fomby, 1997). That is, the error correction mechanism operates only when deviations from the long-run equilibrium exceed a critical range [-θ and θ].[2] Here, the inside regime between -θ and θ can be defined as a "neutral band" within which no adjustments take place given an exogenous shock (Goodwin & Piggott, 2001; Meyer, 2004; Meyer & von Cramon-Taubadel, 2004).

Tsay (1998) suggested a nonparametric approach to identify possible thresholds in the error correction term. He employed a recursive least-square method for an arranged autoregressive representation and constructed *F*-tests to examine whether the standardized predicted residuals from a recursive least-squares estimation follow a linear *AR(p)* process (Tsay, 1998). A threshold exists if the null hypothesis that *AR(p)* follows a linear process is rejected. The delay parameter *d* with the largest *F*-statistic value indicates the optimal lag structure for the Heaviside indicator function I_t (Goodwin and Holt, 1999; Goodwin and Piggott, 2001). If nonlinearities in the error correction term are observed, we then proceed to estimate the threshold value θ using Chan's (1993) grid search method in which threshold values are obtained through a search over all possible threshold values minimizing the Squared Sum of Errors (SSE). Specifically, the threshold variable $|\varepsilon_{t-d}|$ is first sorted from the lowest to the highest value. Second, the *TAR* model in equation (2) is estimated using the ordered values of $|\varepsilon_{t-d}|$ as thresholds. Finally the threshold value θ is chosen so that the SSE is minimized. Hansen (1997) argued that the conventional test is not appropriate given that the null hypothesis of linearity in the *AR* process does not follow a standard distribution. Consequently, he proposed a Chow-type test for threshold values using simulation methods, and provided asymptotic *p*-values based on bootstrap methods (Hansen, 1997; Goodwin and Holt, 1999).

Once the presence of threshold effects is confirmed, the error correction model conditional on threshold values can be estimated. Given that international and retail prices may be determined simultaneously, we employ Seemingly Unrelated Regression (SUR) estimation (Zellner, 1962) taking into account the threshold error correction representation to measure

[2] As argued in Hansen and Seo (2002) and Meyer (2004), the threshold co-integration model with three regimes, separated by different threshold values θ_1 and θ_2, is often criticized because such values cannot be tested in the context of a multivariate error correction model. Hansen and Seo (2002) suggested an absolute value of deviations (i.e. error correction terms) as the possible thresholds, following Balke and Fomby (1997), to test the significance of estimated threshold values.

long-run price adjustments and short-run price dynamics. A simultaneous representation of the system of equations therefore yields

$$\Delta RP_t = \alpha_0^{OUT} I_t \widehat{ECT}_{t-1} + \alpha_0^{IN} I_t \widehat{ECT}_{t-1} + \sum_{i=1}^{p} \alpha_{1,i} \Delta RP_{t-i} + \sum_{i=0}^{p} \alpha_{2,i} \Delta IP_{t-i} + \sum_{i=0}^{p} \alpha_{3,i} \Delta z_{1,t-i} + u_{1,t} \quad (4)$$

$$\Delta IP_t = \beta_0^{OUT} I_t \widehat{ECT}_{t-1} + \beta_0^{IN} I_t \widehat{ECT}_{t-1} + \sum_{i=1}^{p} \beta_{1,i} \Delta IP_{t-i} + \sum_{i=0}^{p} \beta_{2,i} \Delta RP_{t-i} + \sum_{i=0}^{p} \beta_{3,i} \Delta z_{2,t-i} + u_{2,t} \quad (5)$$

where $ECT_{t-1} = \varepsilon_{t-1} = RP_{t-1} - \sigma_0 - \sigma_1 IP_{t-1}$ and the Heaviside indicator function I_t is determined by $I_t = \begin{cases} 0 & if\ |ECT_{t-d}| \leq \theta \\ 1 & if\ |ECT_{t-d}| > \theta \end{cases}$. $\Delta z_{k,t-i}\ \forall\ k = 1,2$ are the identifying variables for the short-run dynamics in the retail and international price equations, respectively.

To examine possible short-run asymmetries in price transmission, the ΔRP_{t-i}, ΔIP_{t-i} and $\Delta z_{1(2),t-i}$ in equations (4) and (5) can be split into positive and negative changes (von Cramon-Taubadel and Loy, 1996). As a result, equations (4) and (5) can be modified to incorporate the following asymmetric threshold error correction model (ATECM) representation:

$$\Delta RP_t = \alpha_0^{OUT} I_t \widehat{ECT}_{t-1} + \alpha_0^{IN} I_t \widehat{ECT}_{t-1} + \sum_{i=1}^{p} \alpha_{1,i}^{+} \Delta^{+} RP_{t-i} + \sum_{i=1}^{p} \alpha_{1,i}^{-} \Delta^{-} RP_{t-i} + \sum_{i=0}^{p} \alpha_{2,i}^{+} \Delta^{+} IP_{t-i} +$$

$$\sum_{i=0}^{p} \alpha_{2,i}^{-} \Delta^{-} IP_{t-i} + \sum_{i=0}^{p} \alpha_{3,i}^{+} \Delta^{+} z_{1,t-i} + \sum_{i=0}^{p} \alpha_{3,i}^{-} \Delta^{-} z_{1,t-i} + u_{1,t} \quad (6)$$

$$\Delta IP_t = \beta_0^{OUT} I_t \widehat{ECT}_{t-1} + \beta_0^{IN} I_t \widehat{ECT}_{t-1} + \sum_{i=1}^{p} \beta_{1,i}^{+} \Delta^{+} IP_{t-i} + \sum_{i=1}^{p} \beta_{1,i}^{-} \Delta^{-} IP_{t-i} + \sum_{i=0}^{p} \beta_{2,i}^{+} \Delta^{+} RP_{t-i} +$$

$$\sum_{i=0}^{p} \beta_{2,i}^{-} \Delta^{-} RP_{t-i} + \sum_{i=0}^{p} \beta_{3,i}^{+} \Delta^{+} z_{2,t-i} + \sum_{i=0}^{p} \beta_{3,i}^{-} \Delta^{-} z_{2,t-i} + u_{2,t}\ , \quad (7)$$

where $\Delta^{+} y_{t-i} = \Delta y_{t-i}$ if $\Delta y_{t-i} > 0$, zero otherwise; and $\Delta^{-} y_{t-i} = \Delta y_{t-i}$ if $\Delta y_{t-i} < 0$, zero otherwise, for $y = (RP, IP, z_1, and\ z_2)$.

We follow a systematic approach to determine the appropriate model specification to assess impacts of the end of the EQS. We first investigate the time series properties of international and retail coffee prices including nonstationarity and co-integration using various unit-root tests and the Johansen's co-integration test. Second, we examine possible nonlinearities in the co-integrating vector following Tsay (1998). If nonlinearities exist, we then estimate the threshold value θ using the grid search method of Chan (1993); and we test for the statistical significance of the threshold estimates following Hansen (1997). Third, for each importing country, we estimate the system of equations (4)-(5) for a symmetric TECM and the system of equations (6)-(7) for an asymmetric TECM for two periods: The EQS period, from January 1980 through December 1989; and the post-EQS period, from January 1990 to December 2009. We employ SUR methods to obtain parameter estimates. Next, we test the short-run parameter asymmetries using the *F*-tests under the null hypothesis of symmetries and employ the model selection criteria to assess whether a symmetric or an asymmetric representation is more appropriate to examine price transmission during and after the EQS termination.

3. Data

We employ monthly data of international coffee prices (the weighted average price of different coffee varieties) and retail prices of roasted coffee in France, Germany and the United States during the period January/1980 to December/2009. These data were obtained from the International Coffee Organization (ICO). Retail prices of roasted coffee and international composite prices are denoted by US dollars per pound. We compile monthly exchange rates of the French Franc and the German Mark[3] to the US dollar from the Federal Reserve Bank Statistics (2010) and we use them as identification variables in the retail price equations of France and Germany, respectively. In the United States equation, we employ the Consumer Price Index for food and beverages from the Bureau of Labor Statistics (2010). In the international price equation, we use monthly average precipitation in Fortaleza, Brazil from the National Centre for Atmospheric Research (2010) for identification purposes, because weather patterns in this country influence international prices.

In Table 1, we present descriptive statistics of these data. The results show that the average retail price in Germany ($4.13 per pound) is the highest among the three countries. Table 1 also suggests that international prices exhibit less variability than retail prices in the three countries. Specifically, the standard deviation of international prices is 0.37, whereas those for retail prices in France, Germany, and the United States are 0.67, 0.81, and 0.51, respectively. Among the three countries, Germany shows the largest variation in retail prices.

[Table 1 here]

4. Results

[3] Conversion factor between the Franc and the Mark and Euro has been employed since January 2002. For German Marks, 1 Euro = 1.95583 DM; for French Francs, 1 Euro = 6.55957FF.

4.1 Test of Integration and Co-integration

We first test the time-series properties of the price data during the EQS and the post-EQS periods. We conduct augmented Dickey-Fuller (*ADF*) and *DF-GLS* (Elliott, Rothenberg and Stock, 1996; Elliott, 1999) tests under the null hypothesis of nonstationarity; and we also use the *KPSS* (Kwiatkowski et al., 1992) test under the null hypothesis of stationarity. We present test results for first differences of each price series in Table 2.[4] The *ADF-t* and *DF-GLS* tests for all first difference variables (international price and retail prices in the three countries) indicate rejection of the null hypothesis of nonstationarity. Furthermore, the KPSS tests cannot reject the null hypothesis of stationarity, showing that all price series in first differences follow *I(0)* processes.

[Table 2 here]

We follow Johansen's (1992a, 1992b, 1995) approach to test whether international and retail price series are co-integrated. This procedure identifies the number of equations that determine the co-integration relationship between prices in each country. For both periods and the three countries, we construct λ_{max} and *trace* tests between international and retail prices. We present test results in Table 3, where *r* represents the co-integration rank (i.e., the number of co-integration vectors). Test results indicate that the relationship between international and retail prices in each country has at least one co-integrating vector. These results imply the existence of a long-run relationship between two price series in the three countries.

[Table 3 here]

4.2 TAR Parameter Estimates

Table 4 presents the parameter estimates of the TAR model in Equation (2). We employ the AIC and SBC criteria to identify the optimal lag structure of each TAR model. The delay parameter, *d*, was selected based on the Tsay test indicating the largest *F* statistic (Goodwin and Holt, 1999; Goodwin and Piggott, 2001). According to the test, we find strong evidence of threshold effects in the co-integrating vector (ε_{t-1}) in both periods (EQS and post-EQS) and in the three countries. The test statistics imply that the null hypothesis of a linear autoregressive (AR) process in the co-integrating vector is rejected at the 5 percent significance level in the three countries and in both periods.

[4] The unit root test results for level variables indicate all data series for both periods follow *I(1)* processes. The specific test results are available on request from authors.

[Table 4 here]

The threshold value θ presented in Table 4 is estimated with Chan's grid search method. The magnitude of the threshold decreases during the post-EQS period relative to its EQS counterpart: from 0.28 to 0.20 in France; 0.48 to 0.18 in Germany; and 0.18 to 0.08 in the United States. These values suggest that the magnitude of deviations beyond which price adjustments take place decreased during the post-EQS period in all countries. We also note that the threshold in the United States is smaller than in both European countries.

Table 4 also shows the percentage of observed deviations that lie in the inside regime (i.e., deviations from the long-run equilibrium within the interval $[-\theta, \theta]$). These percentages decrease during the post-EQS period in Germany and in the United States. However, somewhat surprisingly, the percent of observations in the inside regime increases in France during the post-EQS period. Following Balke and Fomby (1997) and Goodwin and Piggott (2001), the interval $[-\theta, \theta]$ can be interpreted as the range where no adjustment toward the long-run equilibrium occurs. This may be due to transaction costs arising from adjusting retail prices in response to changes in international prices. Therefore, a shrinking threshold interval suggests that price adjustments have taken place more frequently during the post-EQS period compared to the EQS period. The German data yield the steepest decline (from 55% to 24%) in observations in the inside regime, suggesting a greater extent of adjustment toward the long-run equilibrium after the EQS termination. In contrast, the United States coffee market seems to be the least affected by the EQS termination, as the percentage of observations in the inside regime decreases modestly, from 32% in the EQS period to 25% in the post-EQS period.

The Hansen tests presented in Table 4 also reject the null hypothesis of no threshold effects for both periods and all three countries at the 5 percent significance level. These results provide additional evidence of threshold effects in the co-integrating vector of each country. Additionally, the *F* statistics to test the null hypothesis of symmetry (last row in Table 4) confirm the existence of the long-run asymmetries across regimes, further suggesting the presence of nonlinearities in the error correction term.

4.3 Model Selection

Given the existence of thresholds generating nonlinearities in the co-integrating vector of each country, we first estimate the system of equations (6)-(7) using SUR to test possible short-run asymmetries in contemporary and lagged explanatory variables.

In Table 5 we show the χ^2 statistics corresponding to the null hypothesis of symmetry between positive and negative variables in contemporary and lagged international price changes in equation (6)-(7) for both periods (EQS and post-EQS). Our results provide strong evidence of symmetries during the EQS period and asymmetries during the post-EQS period in France. The null hypothesis of symmetries between split variables in contemporary and lagged international prices is strongly accepted in the EQS period, but strongly rejected in the post-EQS period. In contrast, there is modest evidence of asymmetries for Germany and the United States in both periods. The null hypotheses of symmetries are accepted for contemporary variables, but rejected for lagged variables in both periods.

[Table 5 here]

The AIC measures presented in Table 6 provide additional information for model selection. During the EQS period, the AIC values of the symmetric model specifications are smaller than for their asymmetric model counterparts in all three countries. This indicates that a symmetric formulation (equation (4)-(5)) is more appropriate for the EQS period. In contrast, goodness-of-fit measures for the post-EQS period favor an asymmetric representation. Overall, Table 6 supports a symmetric TECM for the EQS period and an asymmetric TECM for the post-EQS period in the three countries.

[Table 6 here]

4.4 Price Transmission in the long-run

Tables 7, 8 and 9 show the parameter estimates of the retail price equations (4) and (6) corresponding to a symmetric TECM model during the EQS period and an asymmetric TECM model in the post-EQS period for France, Germany and the United States, respectively.[5] The estimated coefficients of ECT_{t-1}^{OUT} and ECT_{t-1}^{IN} describe the speed of adjustment towards the long-run equilibrium in each regime. The 'OUT' regime represents deviations outside the threshold values $[-\theta, \theta]$; and the 'IN' regime represents deviations of magnitude smaller than the threshold (i.e. the inside threshold range).

For France and Germany, the estimated coefficients of the outside regime are negative, as predicted by theory, and statistically significant in both periods. Our results indicate that, in both countries, the speed of adjustment in the outside regime decreases slightly during the post-EQS period. In France (Germany), deviations from the long-run equilibrium adjust at a rate of 0.048 (0.062) in the EQS period; whereas this speed decreases to a rate of 0.043

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[5] We employ SUR model for parameter estimates. We did not present the parameter estimates of the international price equation. The estimates results are available on request from authors.

(0.046) during the post-EQS period. The extent to which the speed of adjustment decreases is much larger in Germany than in France, suggesting that Germany went through a more dramatic change in the price transmission process after the end of the EQS than France. Additionally, the parameter estimates suggest that there are no significant adjustments in the inside regime in both countries, which is consistent with the existence of no adjustment range for deviations smaller than the threshold.

In the United States, the speed of adjustment of the outside regime is negative and significant for the post-EQS period, while this parameter estimate for the EQS period is not significant. In contrast to France and Germany, the estimated coefficients of error correction terms for the inside regime are significant for both periods and the speed of adjustment is much faster in the inside regime than in the outside regime. These results are contrary to expectations and may be due to the smaller deviations from the long-run equilibrium in the United States than in France and Germany. As mentioned in TAR estimates results, the threshold values of the United States for the EQS period and the post-EQS period are 0.18 and 0.08, respectively, which are much smaller than those of the two European countries. That means, in the United States, price adjustment toward the long-run equilibrium takes place for even small deviations.

These results show fundamental differences in long-run behavior of price adjustments between the two European countries and the United States. In France and Germany, the speed of adjustment toward the long-run equilibrium occurs only when deviations exceed a critical range; whereas in the United State, there is a meaningful price adjustment of deviations which lie in the inside regime. Moreover, the magnitude to which the speed of adjustment decreases is larger in the United States than in the two European countries. Overall, our results indicate that during the post-EQS period, retail prices became more responsive to changes in international prices (i.e. the threshold values decrease, according to Table 4), but the speed of adjustment toward the equilibrium tends to decrease.

[Table 7 here]

[Table 8 here]

[Table 9 here]

4.5 Price Transmission in the short-run

We examine the short-run dynamics of price transmission through the contemporary and lagged parameters of changes in international prices, retail prices, and the identification variables. The results show substantial differences across periods and countries.

In France, the retail price equation explains 85 and 77 percent of the variation in retail prices during the EQS period and the post-EQS period, respectively. Table 7 indicates that changes in contemporary and lagged international prices do not affect the retail prices in the short run during the EQS period. In contrast, a contemporary negative change in international price of $1 is associated with a $0.24 decrease in retail prices during the post-EQS period. Furthermore, a $1 increase in lagged international prices leads to a $0.23 increase in retail prices in the same period; and a $1 decrease in lagged international prices leads to a $0.20 increase in retail prices. Our results suggest that a $1 increase in international prices in the previous month (*t-1*) leads to a $0.23 ($\alpha_{2,t}^{+} + \alpha_{2,t-1}^{+}$= 0 + 0.23) increase in retail prices in the current month (*t*). In contrast, a $1 decrease in international prices in the previous month (*t-1*) leads to a $0.04 ($\alpha_{2,t}^{-} + \alpha_{2,t-1}^{-}$= 0.24 – 0.20) decrease in retail prices in the current period (*t*). These results suggest that in the post-EQS period, retail prices are more responsive to positive changes than to negative changes in international prices, providing evidence of short-run price transmission asymmetries. Exchange rates have a significant effect on retail prices in France in both periods. However the response of retail prices to the positive changes in exchange rates is stronger and the response to the negative changes is weaker during the post-EQS period.

We present the parameter estimates for Germany in Table 8. The retail price equation accounts for 69 and 55 percent of the variation in retail prices during the EQS period and the post-EQS period, respectively. The results show that, similar to France, changes in international prices do not influence retail prices in the EQS period. In contrast, in the post-EQS period, both positive and negative changes in international prices affect retail prices. Our results indicate that a $1 increase (decrease) in the contemporary international prices causes a $0.37 increase (decrease) in retail prices. However, a $1 increase positive change in the lagged international price does not have a significant impact on retail prices, while a $1 decrease leads to a $0.36 increase in retail prices. Overall, a $1 increase in international prices in the previous period (*t-1*) is associated with a $0.37 ($\alpha_{2,t}^{+} + \alpha_{2,t-1}^{+}$= 0.37 + 0) increase in retail prices in the current period (*t*); whereas, a $1 decrease in international prices in the previous period (*t-1*) is associated with a $0.21 ($\alpha_{2,t}^{-} + \alpha_{2,t-1}^{-}$= 0.57 – 0.36) decrease in retail prices in the current period (*t*). Similar to France, the response to changes in exchange rate is larger for positive changes and smaller for negative changes during the post-EQS period. However, the extent to which exchange rate affects retail prices is much greater in Germany than in France, reflecting the differences of unit currency values with respect to the US dollar.

The short-run price dynamics in Germany appear to be similar to those in France. Namely, retail prices are more responsive to positive than to negative changes in international prices. However, the magnitude of the impact is much larger in Germany. In particular, a $1 change (increase and decrease) in international prices, is associated with modest retail price changes in France ($0.23 increase and $0.04 decrease) relative to price changes in Germany ($0.37 increase and $0.21 decrease). These differences in retail price response to international price changes may relate to market characteristics of each country. A unique characteristic of the German coffee sector is the high market share of hard-discounter retailers (e.g. Aldi, Lidl) which is often associated with the price war taking place in the German retail sector in the late 1990s and early 2000s (Körner, 2002). Hard discounters in Germany often choose to lower prices as a strategy to gain market share when international prices experience negative changes. In contrast, retail prices in France are less responsive than in Germany, perhaps because of stricter regulations on promotional pricing in retail sector.[6]

Table 9 presents the parameter estimates for the United States. The retail price equation explains 41 and 53 percent of the variation in retail prices during the EQS period and the post-EQS period, respectively. Parameter estimates indicate that, in contrast to France and Germany, changes in international prices have a significant effect on retail prices in both periods. However, the extent of such effects differs between periods. Specifically, in the EQS period, a $1 increase in contemporary international prices leads to a $0.58 decrease in retail prices, while a $1 increase in lagged international prices results in a $0.45 increase in retail prices. In sum, a $1 increase in international prices in the previous month (*t-1*) leads to a $0.13 ($\alpha_{2,t} + \alpha_{2,t-1}$= -0.58 + 0.45) decrease in retail prices in the current month (*t*). These results are contrary to expectations and may be due to other factors (e.g. changes in distribution costs in the United States), not accounted for in this study. The degree of price transmission for changes in international prices is stronger after the end of the EQS. Based on the presence of asymmetries during the post-EQS period, a $1 increase in contemporary international prices leads to a $0.26 increase in retail prices, whereas decreases do not influence retail prices. In addition, a $1 increase in lagged international prices is associated with a $0.94 increase in retail prices. However, a $1 decrease in lagged international prices leads to a $0.45 increase in retail prices. Overall, for the post-EQS period, our results suggest that a $1 increase in international prices in the previous period (*t-1*) leads to a $1.20

[6] Since the Galland Law passed in 1996, price promotions in France are restricted to prevent processors and retailers from selling at a loss to take advantage of volume discounts and other promotions offered by coffee processors (Dobson Consulting, 1999; Gómez, Lee and Körner, 2010).

($\alpha^{+}_{2,t} + \alpha^{+}_{2,t-1}$= 0.26 + 0.94) increase in retail prices in the current month (t), while a \$1 decrease in international prices in the previous period (*t-1*) is associated with a \$0.45 ($\alpha^{-}_{2,t} + \alpha^{-}_{2,t-1}$= 0 – 0.45) increase in retail prices in the current period (t). That is, during the post-EQS period, price transmission exceeds hundred percent in the case of positive changes in international prices. Moreover, retail prices increase even in response to negative change in international prices. Changes in consumer price index for food and beverages do not have a significant effect on retail prices in the United States in both periods.

Our results on the short-run dynamics in the post-EQS period for the United States differ from the findings in the two European countries. In the United States, positive changes in international prices are transmitted to retail prices to a greater extent than in its European counterparts; whereas negative changes in international prices lead to a fall in retail prices in France and Germany, but not in the United States. These differences in short-run price transmission between the United States and the two European countries may be related to market characteristics in the United State where the market share of leading brands is higher and share of private labeled brands is lower, indicating higher concentration in the coffee processing sector (Gómez, Lee and Körner, 2010).

4.6 Nonlinear impulse response analysis

We provide further evidence of PTAs through a nonlinear impulse response analysis. Potter (1995) pointed out the limitations of linear impulse response functions, given that the effect of exogenous shocks on the time path of responses can be affected by both the magnitude and direction (Goodwin and Holt 1999; Abdulai 2002). To overcome these limitations, Potter (1995) suggested a modified representation of the linear impulse response function by replacing the linear predictor with a conditional expectation as follows:

$$NIRF_n(\delta;\ X_t, X_{t-1}, \cdots) = E[x_{t+n} | X_t = x_t + \delta, X_{t+1} = x_{t+1}, \cdots] - E[x_{t+n} | X_t = x_t, X_{t+1} = x_{t+1}, \cdots], \quad (8)$$

where X_t is observed data and δ is the postulated impulse. Figure 2 illustrates retail price responses to positive and negative shocks in international prices with a magnitude of one standard deviation, for the three countries. The impulse response paths are calculated for the EQS and the post-EQS period for comparative purposes. During the EQS period, the international price shocks exhibit symmetric effects on retail prices in all three countries, regardless of the direction of the changes in international prices. However, the responses are slightly different across countries. The responses of retail prices in Germany disappear after eight months, while those of the United States vanish after five months for both positive and negative shocks. Furthermore, retail prices adjust fast in the United States (after one month), while these adjustments take place gradually in Germany and France.

In marked contrast, the responses of retail prices to international shocks seem to be asymmetric in the post-EQS period. In Germany, for instance, a positive shock in international prices persists for four months, while a negative shock of the same magnitude disappears after three months. The adjustment process is quite different in the United States. Positive shocks are absorbed almost entirely after two months, while negative shocks tend to last until three months. In the case of France, retail prices adjust more rapidly to negative shocks than to positive shocks, although the adjustment disappears after three months in both cases. Overall, our results suggest that in contrast to EQS period where shocks impacts last longer, the effects of changes in international prices disappear faster in the post-EQS period, regardless of the direction of the shock.

[Figure 2 here]

Lastly, we compare parameter estimates for the speed of adjustment (ECT_{t-1}) and the short-run dynamics (ΔIP_{t-i}) between a standard ECM and an asymmetric TECM for the post-EQS period.

Our findings indicate that parameter estimates of error correction terms in the standard ECM generally show a slower speed of adjustment towards the long-run equilibrium compared to those of the threshold ECM.[7] Moreover, a comparison for the short-run dynamics suggests that the parameter estimates of the symmetric ECM tend to underestimate the short-term impacts of positive changes in international prices; and overestimate the short-term impacts of negative changes in international prices.[8] Overall, this comparison suggests that ignoring

[7] For France, the parameter estimate for the standard ECM and the ATECM are -0.034 and -0.043, respectively; for Germany, the parameter estimate for the standard ECM and the ATECM are -0.045 and -0.046, respectively; and for the United States, the parameter estimate for the standard ECM and the ATECM are -0.108 and -0.138, respectively. All parameter estimates are significant at 5% significance level.

[8] For France, the short-term impacts of a $1 increase (decrease) in international prices are associated with a $0.09 increase ($0.09 decrease) in the standard ECM and a $ 0.23 increase ($ 0.04 decrease) in the ATECM; for Germany, the short-term impacts of a $1 increase (decrease) in international prices lead to a $0.39 increase ($0.39 decrease) in the standard ECM and a $ 0.37 increase ($ 0.21 decrease) in the ATECM; and for the United States, the short-term impacts of a $1 increase (decrease) in international prices are associated with a $0.45 increase ($0.45 decrease) in the standard ECM and a $ 1.21 increase in the ATECM.

threshold effects and asymmetries may lead to incorrect impacts assessments of the EQS termination.

5. Concluding Remarks

In this study we investigate price transmission between international and retail coffee prices in the three largest coffee-importing countries. We examine the impact of the EQS termination, taking into account the existence of thresholds and asymmetries in the price transmission process. Our findings confirm the existence of long-run threshold effects in both periods (EQS and post-EQS) and in the three countries. Based on this evidence, our approach to model selection suggests that a symmetric TECM is more appropriate to study price transmission in the EQS period. In contrast, an asymmetric TECM is more suited to study price transmission in the post-EQS period.

Our results indicate that retail prices become more responsive to changes in international prices (i.e. the threshold values become smaller) and that the speed of adjustment in the long-run decreases slightly in post-EQS period in the three countries. In the short-run, changes in international prices had a modest effect on retail prices in France and Germany during the EQS period, while they had a significant impact on retail prices in the United States in same period. In contrast, after the end of the EQS, changes in international prices affect retail prices in all three countries asymmetrically, but the magnitude of the effects differs across countries. For instance, after the end of the EQS, the transmission of positive changes in international prices to retail prices is higher for the United States than for France and Germany. Conversely, the transmission of negative changes in international prices to retail prices is higher for the two European countries than for the United States. Moreover, our nonlinear impulse response analyses suggest that shocks in international coffee prices trigger faster adjustment toward the equilibrium in the post-EQS period than in the EQS period. Lastly, our results suggest that accounting for threshold effects and asymmetries in price transmission can lead to more accurate impact assessments of policy changes affecting international supply chains for agricultural commodities.

Our study provides valuable insights regarding the application of an ATECM representation for policy evaluation, but several limitations indicate the need for future research. In particular, price transmission from upstream to downstream markets in food supply chains are closely related to market structure. That is, the extent of price transmission depends on the behavior of consumer and firm as well as on the exertion of market power by supply chain participants. Consequently, future research on price transmission using threshold error correction models should incorporate formal models of market structure and their conduct.

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Table 1

Descriptive statistics of data, 1980:1-2009:12 (N=360)

	Mean	Std. Dev	Max	Min
International price	1.014	0.365	2.042	0.412
Retail price in France	3.061	0.674	4.717	1.904
Retail price in Germany	4.125	0.810	6.179	2.473
Retail price in the US	3.136	0.510	4.669	2.352
Exchange Rate (Franc/US Dollar)	5.982	1.171	4.041	10.093
Exchange Rate (Mark/US Dollar)	1.861	0.420	3.303	1.241
Consumer Price Index, Foods and Beverages [a]	1.480	0.373	0.833	2.192
Precipitation (100mm)	1.348	1.527	8.310	0

a. Index 2000 = 1.

Table 2

Tests of integration in first differences of price series

Variables in First Differences			Δ Retail Price in France	Δ Retail Price in Germany	Δ Retail Price in the U.S.	Δ Inter. Price
ADF-t	$H_0: \sim I(1)$	EQS	-7.13**[a]	-7.96**	-7.30**	-7.64**
		Post- EQS	-12.47**	-11.89**	-10.32**	-13.21**
	$H_0: \sim I(1)$ *no constant*	EQS	-7.15**	-7.94**	-7.32**	-7.62**
		Post- EQS	-12.48**	-11.92**	-10.34**	-13.22**
DF-GLS	$H_0: \sim I(1)$	EQS	-7.15**	-7.95**	-7.32**	-7.63**
		Post- EQS	-9.61**	-10.87**	-10.32**	-13.21**
	$H_0: \sim I(1)$ *no constant*	EQS	-6.15**	-7.36**	-6.54**	-7.62**
		Post- EQS	-7.26**	-10.61**	-9.78**	-11.68**
	$H_0: \sim I(1)$ *no linear trend*	EQS	-7.24**	-8.07**	-7.30**	-7.69**
		Post- EQS	-9.00**	-10.83**	-10.02**	-12.69**
KPSS	$H_0: \sim I(0)$ *no constant*	EQS	0.07	0.14	0.07	0.06
		Post- EQS	0.07	0.09	0.04	0.06
	$H_0: \sim I(0)$ *no linear trend*	EQS	0.31	0.26	0.10	0.08
		Post- EQS	0.17	0.14	0.04	0.06

a. ** and * indicate 5% and 10% significant level, respectively.

Table 3

Test of co-integration (*Johansen* test)

France	H_0:r	EQS period	Post- EQS period
λmax	0	10.43*[a]	29.56**
trace	0	11.92*	29.63**
Germany	**H_0:r**	**EQS period**	**Post- EQS period**
λmax	0	16.31**	16.99**
trace	0	17.74**	17.03**
United States	**H_0:r**	**EQS period**	**Post- EQS period**
λmax	0	26.56**	38.60**
trace	0	28.34**	38.60**

a. ** and * indicate 5% and 10% significant level, respectively.

Table 4

TAR estimates

		France	Germany	United States
Optimal Lags (p)[a]	EQS	1	2	2
	Post- EQS	5	5	7
Delay Parameters (d)[b]	EQS	6	6	6
	Post- EQS	2	1	3
Tsay (1997) Test[c]	EQS	4.42** (0.01)	3.91** (0.01)	2.96** (0.04)
	Post- EQS	2.56** (0.02)	3.70** (0.00)	2.37** (0.02)
Hansen (1997) Test[d]	EQS	7.77** (0.00)	5.74** (0.00)	9.33** (0.00)
	Post- EQS	4.83** (0.00)	3.63** (0.03)	6.80** (0.00)
Threshold (θ)	EQS	0.28 (20.2%)[e]	0.48 (55.3%)	0.18 (31.9%)
	Post- EQS	0.20 (36.1%)	0.18 (23.8%)	0.08 (24.5%)
Long-run Asymmetry across Regimes[f]	EQS	2.72*	10.26**	23.64**
($\rho_1^{(1)} = \rho_1^{(2)}$)	Post- EQS	17.15**	14.80**	15.81**

a. Optimal lags are determined by *AIC* and *SBC*

b. Delay parameters are chosen by the lags giving the largest *F*-statistics from *Tsay* test.

c. *F* test for no linear process and parenthesis shows asymptotic *p* values for test statistics.

d. *F* test for no threshold effects and parenthesis indicates asymptotic *p* values of bootstrap simulations with 100 replications.

e. Numbers in parenthesis indicate the percentage of observed deviations that lie inside the threshold.

f. ** and * indicate 5% and 10% significant level, respectively.

Table 5

Tests of short-run asymmetries (Retail price equation)

Period	$\chi^2(1)$[a]	Null Hypothesis	France	Germany	United States
EQS period	3.84	$\Delta IP_t^+ = \Delta IP_t^-$	0.120	0.018	0.527
		$\Delta IP_{t-1}^+ = \Delta IP_{t-1}^-$	1.576	6.117**	8.169***
Post- EQS period	3.84	$\Delta IP_t^+ = \Delta IP_t^-$	4.670**	0.736	1.603
		$\Delta IP_{t-1}^+ = \Delta IP_{t-1}^-$	15.082***	4.234**	42.484***

a. Critical value at 5% significance level.

Table 6
Comparison of the AIC measures across model specifications[a]

	EQS period		Post- EQS period	
	Symmetric TECM	*Asymmetric TECM*	*Symmetric TECM*	*Asymmetric TECM*
France	**-1290.02**[b]	-1275.73	-2635.71	**-2667.10**
Germany	**-1168.79**	-1163.93	-2333.57	**-2337.27**
United States	**-1269.25**	-1252.46	-2327.11	**-2391.95**

a. AIC is the Akaike Information Criteria; TECM is Threshold Error Correction Model; and ATECM is Asymmetric Threshold Error Correction Model.

b. Number in bold indicates smaller value of AIC measure.

Table 7

Estimation results for France (Retail price equation)

Variables	EQS period (*Symmetric TECM*)		Post- EQS period (*Asymmetric TECM*)	
	Parameter Estimates	Standard Errors	Parameter Estimates	Standard Errors
Constant	- 0.001	(0.004)	- 0.001	(0.007)
ECT_{t-1}^{OUT}	- 0.048***[a]	(0.010)	- 0.043***	(0.009)
ECT_{t-1}^{IN}	- 0.026	(0.021)	0.009	(0.026)
ΔRP_{t-1}	0.500***	(0.074)	--	
ΔRP_{t-1}^{+}	--		0.576***	(0.063)
ΔRP_{t-1}^{-}	--		0.036	(0.106)
ΔIP_{t}	- 0.130	(0.049)	--	
ΔIP_{t}^{+}	--		- 0.008	(0.053)
ΔIP_{t}^{-}	--		0.239***	(0.084)
ΔIP_{t-1}	- 0.003	(0.049)	--	
ΔIP_{t-1}^{+}	--		0.230***	(0.057)
ΔIP_{t-1}^{-}	--		- 0.196**	(0.078)
$\Delta Exrate_{t}$	- 0.461***	(0.020)	--	
$\Delta Exrate_{t}^{+}$	--		- 0.559***	(0.044)
$\Delta Exrate_{t}^{-}$	--		- 0.423***	(0.039)
$\Delta Exrate_{t-1}$	0.179***	(0.042)	--	
$\Delta Exrate_{t-1}^{+}$	--		0.013	(0.067)
$\Delta Exrate_{t-1}^{-}$	--		0.246***	(0.046)
R^2	0.85		0.77	
N	120		240	

a. *** and ** indicate 1% and 5% significant level, respectively.

Table 8
Estimation results for Germany (Retail price equation)

Variables	EQS period (*Symmetric TECM*)		Post- EQS period (*Asymmetric TECM*)	
	Parameter Estimates	Standard Errors	Parameter Estimates	Standard Errors
Constant	- 0.010	(0.007)	0.006	(0.015)
ECT_{t-1}^{OUT}	- 0.062***[a]	(0.016)	- 0.046***	(0.012)
ECT_{t-1}^{IN}	- 0.038	(0.024)	- 0.045	(0.075)
ΔRP_{t-1}	0.135	(0.120)	--	
ΔRP_{t-1}^{+}	--		0.250**	(0.101)
ΔRP_{t-1}^{-}	--		0.017	(0.094)
ΔIP_{t}	0.069	(0.081)	--	
ΔIP_{t}^{+}	--		0.372***	(0.110)
ΔIP_{t}^{-}	--		0.571***	(0.172)
ΔIP_{t-1}	0.127	(0.084)	--	
ΔIP_{t-1}^{+}	--		0.127	(0.125)
ΔIP_{t-1}^{-}	--		- 0.358**	(0.167)
$\Delta Exrate_{t}$	- 1.415***	(0.100)	--	
$\Delta Exrate_{t}^{+}$	--		- 2.464***	(0.304)
$\Delta Exrate_{t}^{-}$	--		- 1.405***	(0.278)
$\Delta Exrate_{t-1}$	0.140***	(0.202)	--	
$\Delta Exrate_{t-1}^{+}$	--		- 0.463	(0.372)
$\Delta Exrate_{t-1}^{-}$	--		0.113***	(0.309)
R^2	0.69		0.55	
N	120		240	

a. *** and ** indicate 1% and 5% significant level, respectively.

Table 9
Estimation results for the Unites States (Retail price equation)

Variables	EQS period (*Symmetric TECM*)		Post- EQS period (*Asymmetric TECM*)	
	Parameter Estimates	Standard Errors	Parameter Estimates	Standard Errors
Constant	- 0.002	(0.012)	- 0.057	(0.014)
ECT_{t-1}^{OUT}	- 0.044	(0.016)	- 0.092***[a]	(0.022)
ECT_{t-1}^{IN}	- 0.243***	(0.052)	- 0.138**	(0.054)
ΔRP_{t-1}	0.522***	(0.078)	--	
ΔRP_{t-1}^{+}	--		0.182***	(0.060)
ΔRP_{t-1}^{-}	--		0.019	(0.137)
ΔIP_{t}	- 0.575***	(0.080)	--	
ΔIP_{t}^{+}	--		0.263***	(0.097)
ΔIP_{t}^{-}	--		0.010	(0.149)
ΔIP_{t-1}	0.447***	(0.088)	--	
ΔIP_{t-1}^{+}	--		0.944***	(0.111)
ΔIP_{t-1}^{-}	--		- 0.447**	(0.153)
$\Delta CPIFB_{t}$	0.820	(2.365)	--	
$\Delta CPIFB_{t}^{+}$	--		2.076	(1.699)
$\Delta CPIFB_{t}^{-}$	--		- 0.256	(6.828)
$\Delta CPIFB_{t-1}$	0.129	(2.321)	--	
$\Delta CPIFB_{t-1}^{+}$	--		1.087	(1.695)
$\Delta CPIFB_{t-1}^{-}$	--		0.702	(6.817)
R^2	0.41		0.53	
N	120		240	

a. *** and ** indicate 1% and 5% significant level, respectively.

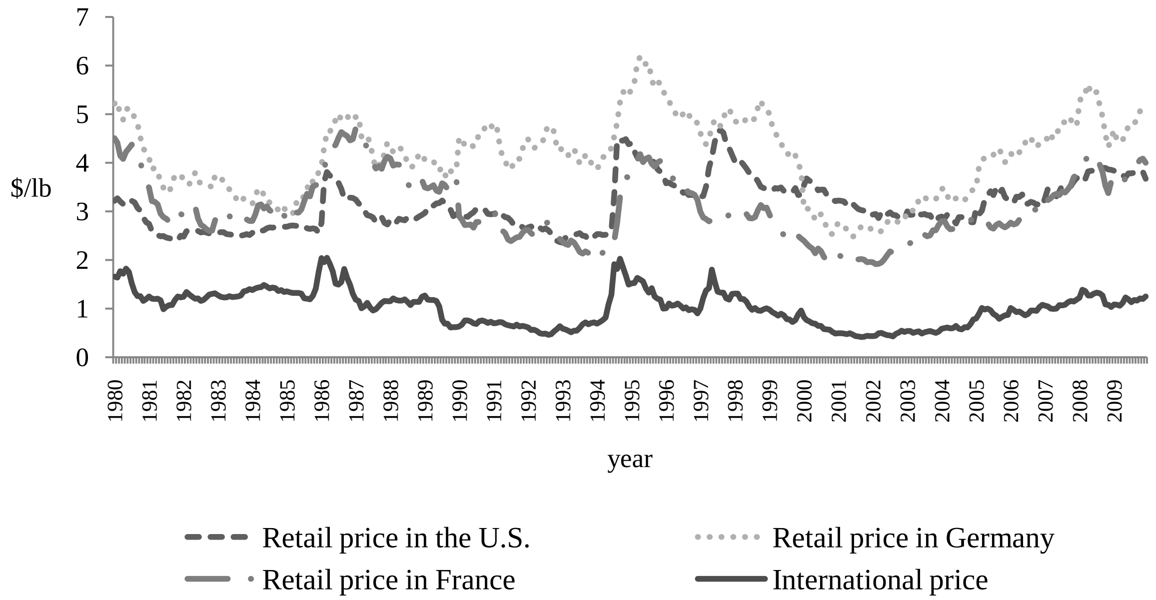

Figure 1. Monthly international coffee prices and retail coffee prices in France, Germany and the United States

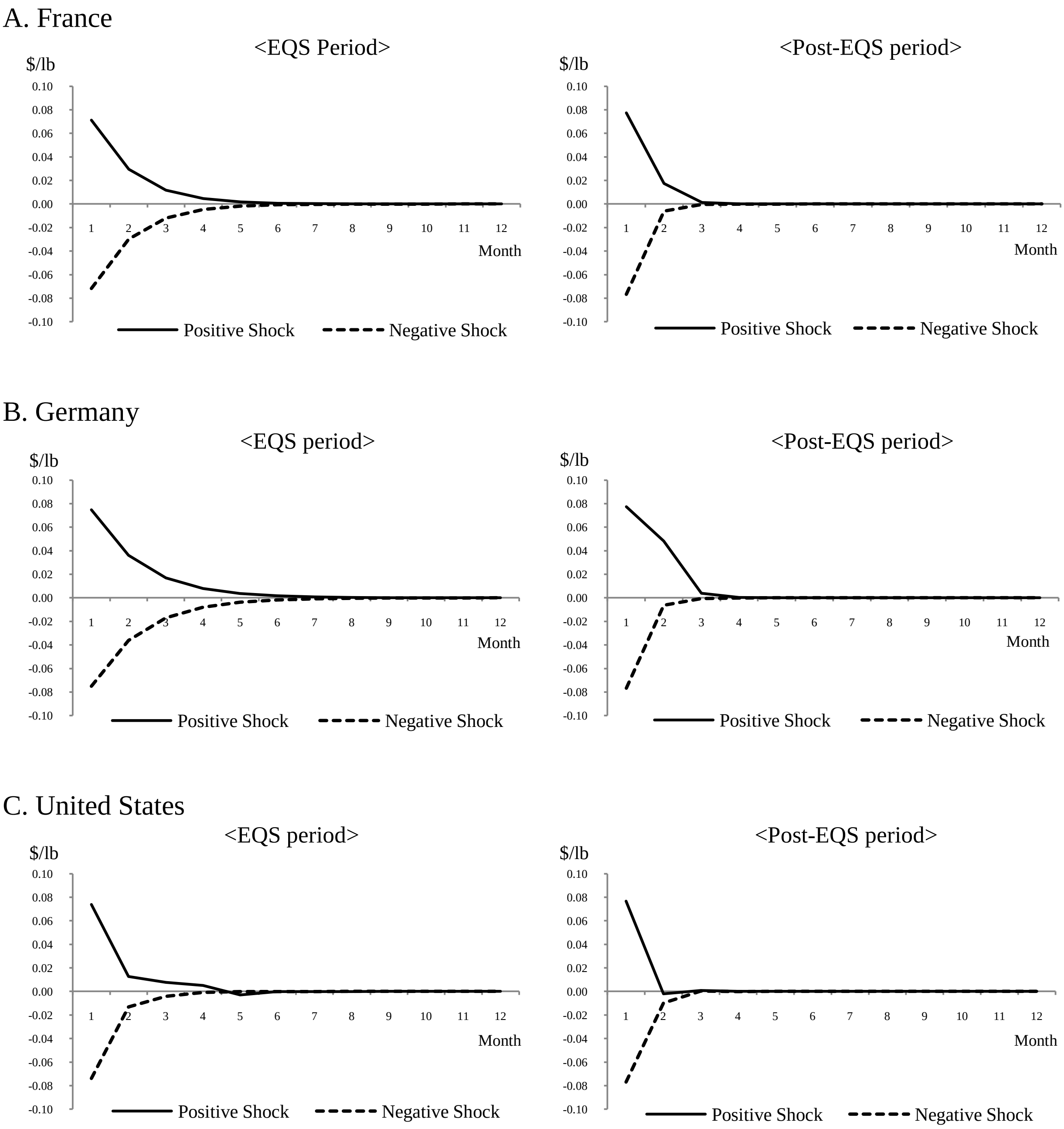

Figure 2. Responses of retail prices to the changes in international prices

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