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Elimination of the Coffee Export Quota System Revisited: Evaluating International-to-Retail Price Transmission

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We revisit the impact of the International Coffee Agreement (ICA) on international-to-retail price transmission. We account for two distinct dimensions (e.g. symmetry vs. asymmetry and linearity vs. nonlinearity) of price transmission from international to retail coffee prices in France, Germany and the United States. We show that ignoring these two features of the price transmission process may lead to misleading impact assessments of the ICA elimination in 1990. Our results confirm the presence of threshold effects in price transmission in both periods (ICA and post-ICA) in the three countries. Our estimates show that, in the long-run, the speed of adjustment toward equilibrium becomes faster during the post-ICA period in France and Germany. Our results suggest that, for France and Germany, changes in international prices did not influence retail prices in the short-run during the ICA period; in contrast, retail prices responded to changes in international prices in the post-ICA period. We find differences between the two European countries and the United States. Our results indicate that changes in international prices influenced U.S. retail prices in both periods. Nonlinear impulse response analysis indicates that ICA elimination increased the speed of adjustment toward the long-run equilibrium, given a shock in international coffee prices. Overall, our results show that ignoring nonlinearities and asymmetries in price transmission may lead to incorrect impact assessment of policies affecting global agricultural supply chains.

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

JEL Codes: C32, Q17.

Introduction

Price transmission asymmetries (PTAs) in supply chains for agricultural commodities traded internationally have received considerable attention from researchers and decision makers. Observed usually at downstream stage of supply chain, PTAs are closely associated with market structure, market power, consumer behavior and policies, among others. An extensive literature has evolved over time to examine asymmetries existing in price transmission. Proposed by Wolffram (1971) and Houck (1977), early price transmission studies measure asymmetries using the lag of positive and negative first-differences in the exogenous price series. Subsequently, von Cramon-Taubadel (1998) points out that prices at different segments of the supply chain are often co-integrated and that ignoring this feature may lead to spurious regression estimates. Consequently, he suggests the use of error correction models (ECMs) allowing for short-run asymmetric price adjustment to overcome the limitations of Wolffram and Houck approaches (von Cramon-Taubadel and Loy 1996; von Cramon-Taubadel 1998). In the standard ECM, the dependent variable responds identically to deviations from the long-run equilibrium regardless of the magnitude and, moreover, the adjustment occurs in every period (Balke and Fomby 1997). However the presence of transaction costs between spatially separated markets, or other factors generating price friction, may result in the existence of thresholds in the adjustment toward the long-run equilibrium in response to an exogenous shock (Meyer 2004).

We focus on these two distinct and important dimensions of price transmission. One is that price adjustments may respond differently to positive and negative exogenous shocks in the short run (i.e. price transmission asymmetries). The other is that there may be thresholds beyond which long-run adjustments occur. To do this, we employ an error correction model with threshold effects developed 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 based on different regimes separated by threshold values estimated parametrically. In addition, we extend the threshold error correction model by incorporating short-run asymmetric responses to exogenous shocks.

We apply these principles to assess the international-to-retail price transmission implications brought by the elimination of the ICA in the early 1990s. In Figure 1 we show monthly international price and retail coffee prices in France, Germany and the United States during the period 1980 to 2009. The figure suggests that the export quota system eliminations may have affected the three countries in different ways in terms of the response to changes in international prices. The retail prices in the three countries seem to have similar relationship with international prices during the ICA period (Jan/1980-Dec/1989). In contrast, in the post-ICA 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. These differences in retail price responses to international price changes may be in part related to the specific characteristics of the coffee supply chains in each country (Table 1). Coffee processing in the United States is more concentrated than France and Germany. On the other hand, a unique characteristic of the German market is the high share of hard-discount 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). Retail pricing in France, for its part, is more regulated than in Germany and the United States.[1]

A number of researchers have examined the impact of the ICA elimination at various levels of the coffee supply chain. Akiyama and Varangis (1990) employ simulation method for global coffee model and demonstrate that the export quota system contributed to stabilize international coffee prices. Krivonos (2004) conducts a co-integration analysis and finds that the rate of price transmission between farm and international prices increases in the post-ICA period. The author finds that the share of retail value going to coffee growers increased after the ICA elimination; and that in export country domestic prices adjusted faster toward the long-run equilibrium in response to shocks in international prices during this period. Shepherd (2004) examines the impact of ICA's elimination on price transmission using a vector autoregression (VAR) model. The authors argue that elimination of the export quota system did not lead to improved price transmission because of market power exerted by coffee processors. Gemech and Struthers (2007), on their part, find evidence of significant increases in coffee price volatility after the elimination of ICA. Mehta and Chavas (2008) study the impact of ICA on the relationship between farm prices in exporting countries, international prices, and retail prices in importing countries. They find that, in the short-run, retail prices respond asymmetrically to changes in the post-ICA period. In contrast, they find no evidence of asymmetric transmission between wholesale and farm prices. More recently, Gómez, Lee and Körner (2010) examine price transmission from international to retail coffee prices in France, Germany and the United States in post-ICA period during the period 1990-

[1] After the Galland Law is passed in 1996, the 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).

2006 employing error correction model. They find no evidence of long-run price transmission asymmetries. However they provide evidence of short-run asymmetries with substantial differences among countries.

In this study, we revisit the implications of the International Coffee Agreement (ICA) elimination on price transmission between international prices and retail prices in France, Germany and the United States taking into account the existence of thresholds and short-run asymmetries. The primary change brought by the ICA elimination was the end of the export quota system to major importing countries. We show that ignoring these two features of the price transmission process may lead to incorrect impact assessments of the ICA elimination. We contribute to the literature by improving our understanding of the impact of policy interventions on the price transmission of agricultural commodity global supply chains. This paper is organized as follows. We first review the literature on thresholds in price transmission. Next, we develop the asymmetric threshold error correction model (ATECM) representation employed to examine the implications of ICA elimination. In turn, we describe our data and present and compare the empirical results for three countries. Finally we summarize our findings and discuss the benefits and limitations of using an ATECM representation to evaluate the implications of ICA elimination.

Modeling threshold co-integration in price transmission processes

A number of studies have utilized the threshold approach to examine price transmission in supply chains of agricultural commodities. Goodwin and Holt (1999) employ 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) use a TECM to examine co-integration of prices among four corn and soybean markets in North Carolina accounting for transaction costs. More recently, Abdulai (2002) employs the threshold co-integration model of Enders and Granger (1998) to analyze price transmission between producer and retail prices in the Swiss swine-pork supply chain. He compares a standard ECM with a TECM and uses the Akaike and Schwarz information criteria to shows that the threshold representation is superior. Meyer (2004) considers transaction costs occurring potentially during the process of price transmission and employs a vector error correction model with absolute threshold value following the procedures of Balke and Fomby (1997). These studies generally confirm the existence of nonlinear price transmission (i.e. thresholds) between spatially separated markets. These studies also show that TECM representations generally indicate a faster adjustment towards the long-run equilibrium than their standard counterpart.

Here, 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) point out, the co-integration tests of Johansen and Engel-Granger may be misspecified if the adjustment to the long-run equilibrium is nonlinear. To overcome this problem, Enders and Granger (1998) suggest an alternative to the standard augmented Dickey-Fuller (ADF) regression. Consider RP_t the retail coffee price and IP_t the international coffee price at time period *t*. Both price variables are assumed to be integrated of order one, *I(1)*. Then co-integrating relationship between such two price series is given by:

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

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

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

The Heaviside Indicator function in equation (2), I_t, is

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

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

Tsay (1998) suggests a nonparametric approach to identify possible thresholds in the error correction term. He employs recursive least square method for an arranged autoregressive representation and constructs F-tests to examines whether the standardized predicted residuals from recursive least squares estimation follow a linear *AR(p)* process (Tsay 1998). A threshold exists if the null hypothesis that *AR(p)* follows linear process is rejected. The delay parameter d with the largest *F*-statistic value indicates the optimal lag structure for the Heavyside Indicator function *I* (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 estimated through a search over all possible threshold values minimizing the Squared Sum of Errors (SSE). Specifically, in this approach 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 Squared Sum of Errors (SSEs) are calculated from the TAR parameter estimates for each data point, choose the threshold value θ that minimizes the SSE. Hansen (1997) argues that the conventional test is not appropriate given that null hypothesis of linearity in the *AR* process does not follow a standard distribution. Consequently, he proposes a Chow test for threshold values using simulation methods and provides 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

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[2] As argued in Hansen and Seo (2002) and Meyer (2004), the threshold co-integration model with three regimes, separated by 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) suggest using an absolute value of error correction term following Balke and Fomby (1997) to test for the significance of estimated threshold values.

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 long-run price adjustments and short-run price dynamics. A simultaneous representation of equation therefore yields

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

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

where $ECT_{t-1} = \varepsilon_{t-1} = RP_{t-1} - \sigma_0 - \sigma_1 IP_{t-1}$ from equation (1) and the Heaviside indicator function I_t is determined by $I_t = \begin{cases} 0 & if\ |ECT_{t-d}| \le \theta \\ 1 & if\ |ECT_{t-d}| > \theta \end{cases}$. $\Delta z_{k,t-i}\ \forall\ k = 1,2$ are the identifying variables for the short-run dynamics in retail and international price equations, respectively. To investigate the possible short-run asymmetries in price transmission, the ΔRP_{t-1}, ΔIP_{t-1} and $\Delta z_{1(2),t-i}$ in equations (4) and (5) can be separated according to positive and negative changes (von Cramon-Taubadel and Loy 1996). As a result, equations (4) and (5) can be modified to yield the following asymmetric threshold error correction model (ATECM) representation:

$$\begin{aligned}\Delta RP_t = {} & \alpha_0^{(1)} I_t \widehat{ECT}_{t-1} + \alpha_0^{(2)}(1-I_t)\widehat{ECT}_{t-1} + \sum_{i=1}^{p}\alpha_{1,i}^{+}\Delta^{+}RP_{t-i} + \sum_{i=1}^{p}\alpha_{1,i}^{-}\Delta^{-}RP_{t-i} + \sum_{i=0}^{p}\alpha_{2,i}^{+}\Delta^{+}IP_{t-i} \\ & + \sum_{i=0}^{p}\alpha_{2,i}^{-}\Delta^{-}IP_{t-i} + \sum_{i=0}^{p}\alpha_{3,i}^{+}\Delta^{+}z_{1,t-i} + \sum_{i=0}^{p}\alpha_{3,i}^{-}\Delta^{-}z_{1,t-i} + u_{1,t}\end{aligned} \quad (6)$$

$$\begin{aligned}\Delta IP_t = {} & \beta_0^{(1)} I_t \widehat{ECT}_{t-1} + \beta_0^{(2)}(1-I_t)\widehat{ECT}_{t-1} + \sum_{i=1}^{p}\beta_{1,i}^{+}\Delta^{+}IP_{t-i} + \sum_{i=1}^{p}\beta_{1,i}^{-}\Delta^{-}IP_{t-i} + \sum_{i=0}^{p}\beta_{2,i}^{+}\Delta^{+}RP_{t-i} \\ & + \sum_{i=0}^{p}\beta_{2,i}^{-}\Delta^{-}RP_{t-i} + \sum_{i=0}^{p}\beta_{3,i}^{+}\Delta^{+}z_{2,t-i} + \sum_{i=0}^{p}\beta_{3,i}^{-}\Delta^{-}z_{2,t-i} + u_{2,t}\end{aligned} \quad (7)$$

where $\Delta^{+}RP_{t-i} = \Delta RP_{t-i}$ if $\Delta RP_{t-i} > 0$ and $\Delta^{-}RP_{t-i} = \Delta RP_{t-i}$ if $\Delta RP_{t-i} < 0$. $\Delta^{+}IP_{t-i}$, $\Delta^{-}IP_{t-i}$, $\Delta^{+}z_{1(2),t-i}$ and $\Delta^{-}z_{1(2),t-i}$ are similarly obtained.

We follow a systematic approach to determine the appropriate specification to assess impacts of the ICA elimination. 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 co-integration test. Second, we examine possible nonlinearities in the co-integrating vector following Tsay (1998). If nonlinearities exist, we then find the threshold value θ using the grid search method of Chan (1993); and we test the significance of threshold effect following Hansen (1997). Third, for each 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 ICA period, from January 1980 through December 1989; and the post-ICA period, from January 1990 to December 2009. We employ SUR methods to obtain parameter estimates. Next, we employ the AIC and the BIC criteria to assess whether a symmetric or an asymmetric representation is more appropriate to examine price transmission during and after the ICA export quota system.

Data

We employ monthly data on international composite 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 are from the International Coffee Organization (ICO). Retail prices of roasted coffee and international composite prices are in 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 the retail price equations in 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 National Centre for Atmospheric Research (2010) for identification purposes, because weather patterns in this country influence international prices. We present descriptive statistics of these data in Table 2.

Results

Test of Integration and Co-integration

We first test the time-series properties of the price data. 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 (Table 3). The *ADF-t* and *DF-GLS* tests for all variables (international price and retail prices in the three countries) suggest rejection of the null hypothesis of nonstationarity. Furthermore, the KPSS tests cannot reject the null hypothesis of stationarity, indicating that all price series in first differences follow *I(0)* processes.

We follow Johansen's (1992a, 1992b, 1995) approach to test whether the international and retail price series are co-integrated. This procedure identifies the number of equations that determine the co-integration relationship between the international and retail prices in each importing country. For each country, we therefore construct λ_{max} and *trace* tests between the retail price and the international price. We present the results from these tests in Table 4, where *r* represents the co-integration rank (i.e., the number of co-integration vectors). According to these tests, the international price and the retail price in each country have at least one co-integrating vector. This implies the existence of a long-run relationship between international prices and retail prices in each country.

Parameter Estimates

Table 5 presents the estimated parameters from the TAR model in Equation (2). We employed 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* statistics (Goodwin and Holt 1999; Goodwin and Piggott 2001). As a result of Tsay test, we find strong evidence of nonlinearity in series of co-integrating vector (ε_{t-1}) in both periods (during and post-ICA) and three countries. The test statistics imply that the null hypothesis of a linear *AR* process in the co-integrating vector is rejected at the 5 percent significance level in the three countries. Table 5 shows that the percent share of observations in the 'inside' regime (i.e., deviations from the long-run equilibrium in the interval $[-\theta, \theta]$) decreased during the post-ICA period

[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.

in Germany and in the United States. However, somewhat surprisingly, the percent of observations in the ‘inside’ regime increases in France in the post-ICA period. Following Balke and Fomby (1997) and Goodwin and Piggot (2001), the interval $[-\theta, \theta]$ can be interpreted as the range where no adjustment takes place due to transaction costs arising from adjusting retail prices in response to changes in international prices. Therefore a shrinking threshold interval means that price adjustments are more common during the post-ICA period than the ICA period. In this sense, Germany experiences the steepest decline in the range of threshold value from 55% to 24% between periods, implying substantial changes in the price adjustment mechanism in Germany’s coffee supply chain after the elimination of the export quota system. In contrast, the coffee market in the United States seems to be the least affected from the elimination of ICA among the three countries examined in this study. The Hansen tests also reject the null hypothesis of no threshold effects for both periods and all three countries at the 5 percent level of significance. These results provide additional evidence of threshold effect in the co-integrating vector of each country. Additionally, the F statistics to test the null hypothesis of symmetry (last row in Table 5) confirm the existence of the long-run asymmetries across regimes supporting the hypothesis of presence of nonlinearities in the error correction term.

Given the existence of thresholds (i.e. nonlinearities) in the co-integrating vector of each country, we estimate the system of equations (6)-(7) using SUR and we test for possible short-run asymmetries in contemporary and lagged explanatory variables explaining the short-run dynamics between international and retail prices. In Table 6 we show the χ^2 statistics corresponding to the null hypothesis of symmetry in each period (ICA and post-ICA). Our results show that, during the ICA period, there is no evidence of asymmetries in France and very modest evidence of asymmetries in Germany and in the United States. Goodness-of-fit (AIC and SBC) measures presented in Table 7 provide additional support to this finding. During the ICA period, the AIC and SBC values for the symmetric model specifications are lower than for their asymmetric model counterparts, in all three countries. The implication is that a symmetric threshold error correction model specification is more appropriate to explain price transmission during the ICA period. In contrast, during the post-ICA period, we find strong evidence of short-run price asymmetries in France and Germany, and modest evidence of price transmission asymmetries in the United States. Goodness-of-fit measures also suggest that an asymmetric formulation in the post-ICA period is more appropriate than a symmetric model.

Tables 8, 9 and 10 show the parameter estimates corresponding to a symmetric TECM model during the ICA period and an asymmetric TECM model for post-ICA period for France, Germany and the United States, respectively. The estimated coefficients of $ECT_{t-1}^{(1)}$ and $ECT_{t-1}^{(2)}$, in Tables 8, 9 and 10, describe the speed of adjustment towards the long-run equilibrium in each regime after a change in international coffee prices. Regime (1) represents deviations beyond the threshold, outside range between threshold values $[-\theta, \theta]$; and regime (2) 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 are statistically significant. Our results indicate that, for both countries, the speed of adjustment decreases in post-ICA period. In France (Germany), deviations from the long-run equilibrium adjust at the rate of 0.048 (0.062) in the ICA period. However, in the post-ICA period, these speeds decreased to a rate of 0.043 (0.046). 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 elimination of the export quota system than France.

Additionally, the parameter estimates suggest that there are no significant adjustments in the interior regime, or Regime (1), for both countries, consistent with the existence a “band of no adjustment.” In contrast to France and Germany, in United States the speed of adjustment accelerates after the ICA elimination. The estimated coefficient of the error correction term during the ICA period is not significant. However, the speed of adjustment is faster in regime (2) than in regime (1) which is contrary to model expectations. These results show fundamental differences in long-run behavior of price adjustments between the two European countries and the United States.

We examine the short-run dynamics through analyzing contemporary and lagged parameters of both international price and the identification variables (exchange rate in France and Germany and consumer price indexes for food and beverages in the United States). We present these results in Tables 8, 9 and 10 and they indicate substantial differences across periods and countries. In France, for example, Table 8 suggest that while changes in contemporary and lagged international price do not affect the retail price in the ICA period, a contemporary negative change in international price of $1 is associated with a $0.24 retail price decrease in the post-ICA period. Furthermore, a $1 increase of lagged international price leads to $0.23 increase in retail price. But contrary to expectations, a $1 decline of lagged international price leads to $0.20 increase in retail price. Exchange rates have a significant effect on retail price in France in both periods. However, in the post-ICA period, retail price response to the positive changes in exchange rates is stronger while the response to the negative changes in exchange rate is weaker.

We present the parameter estimates for Germany in Table 9. The results show that changes in international prices do no influence retail prices in the ICA period. In contrast in the post-ICA period, both positive and negative changes in international prices affect retail prices. Our results indicate that a $1 increase in the contemporary international price causes a $0.37 increase in retail price. On the other hand, a $1 decrease of contemporary international price leads to $0.57 decrease in retail price. The short-run behavior in Germany is fundamentally different than in France. In Germany, the transmission of international prices to retail prices is larger for negative than for positive changes in international prices.[4] Similar to France, the response for exchange rate fluctuation becomes faster for positive changes and slower for negative changes in the post-ICA regime. Specifically, a one unit increase in the value of the German Mark with respect to the US dollar is associated with a $1.42 increase in retail price. However, the effect of exchange rate is much stronger in Germany than in France, reflecting the differences of unit currency values with respect to the US dollar. Overall, our estimates suggest that in in the ICA period, changes in international prices had modest effects on retail prices in both France and Germany. However, retail prices in these countries became more responsive to changes in international prices, with strong evidence of asymmetric price transmission.

Contrary to France and Germany, international price variations in both periods have a significant effect on the retail price in the United States. However, the extent of such effect differs depending on the regime. In the ICA period, a $1 increase (decrease) in contemporary international prices leads to $0.58 decrease (increase) in retail price, which is unexpected.

[4] These results are consistent with our former results in Gómez, Lee and Körner (2010) which represents the characteristics of Germany’s coffee supply chain where the market share of hard discounter such as Aldi is larger than that of France. Since hard discounters often choose low prices relative to competitors as a strategy to gain the market share.

However, a $1 increase (decrease) of lagged international price results in $0.45 increase (decrease) in domestic price. These results indicate that the effects of international price changes take place after a period of one month. The degree of price transmission for lagged international price change becomes stronger after the elimination of the ICA. Based on the presence of asymmetries in post-ICA period, a 1$ positive change in international prices results in a $0.45 and $0.94 changes in retail price in the ICA and post-ICA periods, respectively. The estimates of consumer price index for food and beverages are not statistically significant.

The results for the United States in the post-ICA period are in sharp contrast to the findings in the European countries. In the United States, increases in contemporary international prices impact retail prices, but decreases appear to have no effect on the retail prices. In contrast, in Germany, negative changes (price decreases) in contemporary international prices have larger effects on retail prices than positive changes do. These results are also observed in parameter estimates for lagged variables (ΔIP_{t-1}^{+} and ΔIP_{t-1}^{-}), although the sign of negative change's variable (ΔIP_{t-1}^{-}) is unexpected. The results for France also show differences in short-run price transmission behavior in comparison to the United States. In this country, a negative contemporary change in international price lead to reductions in retail prices, but positive changes do not affect retail prices. In addition, a $1 increase in international prices in the prior month has a larger effect on retail prices in the United States ($0.94) than in France ($0.23).

We provide further evidence of asymmetric price transmission through an impulse-response analysis. Given the nonlinear characteristics of model, we employ Potter's (1995) approach to conduct an impulse-response analysis. The author points out that in the case of the nonlinear model, the effect of exogenous shocks on the time path of responses is affected by the magnitude and sign of the history of shocks (Goodwin and Holt 1999; Abdulai 2002; Enders 2004). This is in contrast to linear impulse responses model, where the response is independent from the time series history and the sign and magnitude of the shock have no effect on the time-path of responses. Potter (1995) suggests a modified representation of the linear impulse response function, replacing the linear predictor with a conditional expectation as follows:

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

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

In marked contrast to the ICA period, the responses of retail prices to international shocks seem to be asymmetric in the post-ICA period. In Germany, for instance, a positive change in international price 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.

In this country, positive shocks are absorbed almost entirely after two months, while negative shocks tend to last about three months. In the case of France, retail prices adjust more rapidly to negative shocks than to positive shocks, although the adjustment takes place after three months in both cases. Overall, our results suggest that we find in contrast to ICA period where shocks more last, the effects of changes in international prices disappear faster in the post-ICA period than in the ICA period, regardless of the direction of the shock.

Finally, in Table 11 we compare parameter estimates for short-run dynamics between the Threshold Error Correction model employed here and a standard Error Correction Model[5]. In the ICA period, during which a symmetric model fits better the data, deviations from equilibrium generally adjust faster in the threshold model than in the standard model in the three countries. In the post-ICA period, in which an asymmetric specification fits the data better, the estimates of threshold model for France show faster adjustments. However, the estimates for Germany and the United States show the mixed results, depending on the variables. In short, our results show that ignoring thresholds and asymmetries can lead to biased assessments regarding the ICA elimination.

Concluding Remarks

In this study we investigated price transmission between international and retail coffee prices in the three largest coffee-importing countries. We examined the impact of the elimination of export quota system in 1990, taking into account the existence of thresholds and asymmetries in the price transmission process. Our findings suggest the existence threshold effects in the long-run adjustment process in both periods (ICA and post-ICA) and in all three countries. Based on this evidence, our approach to model selection suggests that a symmetric model is more appropriate during the ICA period and that an asymmetric model is more appropriate during the post-ICA period. We find that the speed of adjustment towards the long-run equilibrium after an exogenous shock decreased in the post-ICA period in France and Germany. In contrast, this long-run adjustment became faster during the post-ICA period in the United States.

Our results indicate that the threshold range became smaller in the post-ICA period, particularly for Germany. This indicates that in the post-ICA period retail prices were more responsive to changes in international prices than in the ICA period, even if the change in the latter were of small magnitude. In the short-run, our estimates suggest that during the ICA period, changes in international prices had modest influence on retail prices in France and Germany. In contrast, retail prices in the United States responded to changes in international prices in both periods (ICA and post-ICA). Our analysis of the Impulse Response Functions provides additional evidence of symmetric price transmission during the ICA period and asymmetric price transmission in the post-ICA period in all three countries. Our results also indicate faster adjustment to the long run equilibrium after an exogenous shock in international prices during the post-ICA period than in the ICA period. Overall, our results indicate that ignoring thresholds and asymmetries in the price transmission process may lead to inexact assessment of the impacts of policy changes affecting international supply chains for agricultural commodities.

Our study provides valuable insights regarding the application of an ATECM representation

[5] We estimate a standard symmetric (asymmetric) error correction model following Gómez, Lee and Körner (2010). For ICA period, we employ a symmetric model and for post-ICA period, we use asymmetric model.

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 consumer and firm behavior 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. The characteristics of coffee supply chain in three countries

	France	Germany	United States
Share of leading brands (%) [a]	27.0	28.5	34.7
Share of three leading brands (%) [a]	66.8	63.1	70.2
Share of private labeled brands (%) [a]	14.4	31.1	8.1
Share of five leading supermarkets (%)	76.4	61.8	35.5 [b]
Share of hard-discounter retailers (%)	7.8	34.0	<2.0 [c]

a. Mintel's Market Intelligence for France and Germany (2001 and 2003); Grocery Headquarters State of the Industry Almanac (2002 and 2004).

b. Average for years 1998-2003, from the Food Industry Management Program, Cornell University.

c. Estimates from the Food Industry Management Program, Cornell University.

Table 2. Descriptive statistics of the estimating sample, 1980:1-2009:12

	Mean	**Ste. 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 3. Tests of integration in first differences

Variables in First Differences			**Critical Value**	**Δ Retail Price France**	**Δ Retail Price in Germany**	**Δ Retail Price in U.S.**	**Δ Inter. Price**
ADF-t	$H_0: \sim I(1)$	ICA	-2.88	-7.13	-7.96	-7.30	-7.64
		post-ICA	-2.88	-12.47	-11.89	-10.32	-13.21
	$H_0: \sim I(1)$ *no constant*	ICA	-1.95	-7.15	-7.94	-7.32	-7.62
		post-ICA	-1.95	-12.48	-11.92	-10.34	-13.22
DF-GLS	$H_0: \sim I(1)$	ICA	-1.95	-7.15	-7.95	-7.32	-7.63
		post-ICA	-1.95	-9.61	-10.87	-10.32	-13.21
	$H_0: \sim I(1)$ *no constant*	ICA	-1.95	-6.15	-7.36	-6.54	-7.62
		post-ICA	-1.95	-7.26	-10.61	-9.78	-11.68
	$H_0: \sim I(1)$ *no linear trend*	ICA	-2.89	-7.24	-8.07	-7.30	-7.69
		post-ICA	-2.89	-9.00	-10.83	-10.02	-12.69
KPSS	$H_0: \sim I(0)$ *no constant*	ICA	0.15	0.07	0.14	0.07	0.06
		post-ICA	0.15	0.07	0.09	0.04	0.06
	$H_0: \sim I(0)$ *no linear trend*	ICA	0.46	0.31	0.26	0.10	0.08
		post-ICA	0.46	0.17	0.14	0.04	0.06

Table 4. Test of co-integration (*Johansen* test)

France	H_0:r	ICA period	post-ICA period
λmax	0	10.43*	29.56**
trace	0	11.92*	29.63**
Germany	**H_0:r**	**ICA period**	**post-ICA period**
λmax	0	16.31**	16.99**
trace	0	17.74**	17.03**
United States	**H_0:r**	**ICA period**	**post-ICA period**
λmax	0	26.56**	38.60**
trace	0	28.34**	38.60**

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

Table 5. TAR estimates

		France	Germany	United States
Optimal Lags (p)[a]	ICA	1	2	2
	post-ICA	5	5	7
Delay Parameters (d)[b]	ICA	6	6	6
	post-ICA	2	1	3
Tsay (1997) Test[c]	ICA	4.42** (0.01)	3.91** (0.01)	2.96** (0.04)
	post-ICA	2.56** (0.02)	3.70** (0.00)	2.37** (0.02)
Hansen (1997) Test[d]	ICA	7.77** (0.00)	5.74** (0.00)	9.33** (0.00)
	post-ICA	4.83** (0.00)	3.63** (0.03)	6.80** (0.00)
Threshold (θ)[e]	ICA	0.275 (20.2%)	0.484 (55.3%)	0.177 (31.9%)
	post-ICA	0.195 (36.1%)	0.181 (23.8%)	0.081 (24.5%)
Long-run Asymmetry across Regimes[f]	ICA	2.721*	10.262**	23.641**
($\boldsymbol{\rho_1^{(1)} = \rho_1^{(2)}}$)	post-ICA	17.148**	14.804**	15.809**

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

b. Delay parameters are chosen the delay giving the largest F-statistics in *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. Parenthesis indicates the share of inside range among all data points.

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

Table 6. Tests of short-run asymmetries (Retail price equation)

Null Hypothesis	$\chi^2(1)$ critical value at 5%	Time Period	France	Germany	United States
$\Delta RP_{t-1}^{+} = \Delta RP_{t-1}^{-}$	3.84	ICA	0.600	1.820	0.984
		Post-ICA	17.433***	2.445	1.016
$\Delta IP_{t}^{+} = \Delta IP_{t}^{-}$	3.84	ICA	0.120	0.018	0.527
		Post-ICA	4.670**	0.736	1.603
$\Delta IP_{t-1}^{+} = \Delta IP_{t-1}^{-}$	3.84	ICA	1.576	6.117**	8.169***
		Post-ICA	15.082***	4.234**	42.484***
$\Delta z_{t}^{+} = \Delta z_{t}^{-}$	3.84	ICA	1.020	0.300	2.761
		Post-ICA	3.764	4.679**	0.095
$\Delta z_{t-1}^{+} = \Delta z_{t-1}^{-}$	3.84	ICA	0.056	1.192	0.233
		Post-ICA	7.126**	1.105	0.002

Table 7. Comparison of goodness-of-fit across model specifications[a]

France	ICA period		Post-ICA period	
	TECM	ATECM	TECM	ATECM
AIC	-1290.02	-1275.73	-2635.71	-2667.10
SBC	-1245.69	-1203.69	-2580.15	-2576.82
Germany	TECM	ATECM	TECM	ATECM
AIC	-1168.79	-1163.93	-2333.57	-2337.27
SBC	-1124.46	-1091.89	-2278.02	-2246.99
United States	TECM	ATECM	TECM	ATECM
AIC	-1269.25	-1252.46	-2327.11	-2391.95
SBC	-1224.92	-1180.42	-2271.55	-2301.67

a. AIC = Akaike Information Criteria; SBC = Schwartz Bayesian Criteria; ICA = International Coffee Agreement; TECM = Threshold Error Correction Model; and ATECM is Asymmetric Threshold Error Correction Model.

Table 8. Estimation results for France (Retail price equation)

Variables	ICA period	Variables	Post-ICA period
	STECM		ATECM
Constant	-0.001	*Constant*	-0.001
	(0.004)		(0.007)
$ECT_{t-1}^{(1)}$	-0.048***	$ECT_{t-1}^{(1)}$	-0.043***
	(0.010)		(0.009)
$ECT_{t-1}^{(2)}$	-0.026	$ECT_{t-1}^{(2)}$	0.009
	(0.021)		(0.026)
ΔRP_{t-1}	0.500***	ΔRP_{t-1}^{+}	0.576***
	(0.074)		(0.063)
		ΔRP_{t-1}^{-}	0.036
			(0.106)
ΔIP_{t}	-0.130	ΔIP_{t}^{+}	-0.008
	(0.049)		(0.053)
		ΔIP_{t}^{-}	0.239***
			(0.084)
ΔIP_{t-1}	-0.003	ΔIP_{t-1}^{+}	0.230***
	(0.049)		(0.057)
		ΔIP_{t-1}^{-}	-0.196**
			(0.078)
Δz_{t}	-0.461***	Δz_{t}^{+}	-0.559***
	(0.020)		(0.044)
		Δz_{t}^{-}	-0.423***
			(0.039)
Δz_{t-1}	0.179***	Δz_{t-1}^{+}	0.013
	(0.042)		(0.067)
		Δz_{t-1}^{+}	0.246***
			(0.046)
R^2	0.85	R^2	0.77

a. Standard errors in parenthesis, *** significant at 1% level, ** significant at 5% level.

Table 9. Estimation results for Germany (Retail price equation)

Variables	ICA period	Variables	Post-ICA period
	STECM		ATECM
Constant	-0.010 (0.007)	*Constant*	0.006 (0.015)
$ECT_{t-1}^{(1)}$	-0.062*** (0.016)	$ECT_{t-1}^{(1)}$	-0.046*** (0.012)
$ECT_{t-1}^{(2)}$	-0.038 (0.024)	$ECT_{t-1}^{(2)}$	-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)
Δz_{t}	-1.415*** (0.100)	Δz_{t}^{+}	-2.464*** (0.304)
		Δz_{t}^{-}	-1.405*** (0.278)
Δz_{t-1}	0.140 (0.202)	Δz_{t-1}^{+}	-0.463 (0.372)
		Δz_{t-1}^{+}	0.113 (0.309)
R^2	0.69	R^2	0.55

a. Standard errors in parenthesis, *** significant at 1% level, ** significant at 5% level.

Table 10. Estimation results for United States (Retail price equation)

STECM	ICA period STECM	ATECM	Post-ICA period ATECM
Constant	-0.002 (0.012)	*Constant*	-0.057*** (0.014)
$ECT_{t-1}^{(1)}$	-0.044 (0.031)	$ECT_{t-1}^{(1)}$	-0.092*** (0.022)
$ECT_{t-1}^{(2)}$	-0.243*** (0.052)	$ECT_{t-1}^{(2)}$	-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)
Δz_t	0.820 (2.365)	Δz_t^{+}	2.076 (1.699)
		Δz_t^{-}	-0.256 (6.828)
Δz_{t-1}	0.129 (2.321)	Δz_{t-1}^{+}	1.087 (1.695)
		Δz_{t-1}^{+}	0.702 (6.817)
R^2	0.41	R^2	0.53

a. Standard errors in parenthesis, *** significant at 1% level, ** significant at 5% level.

Table 11. Comparison of estimates from TECM with ECM

Variables	France		Germany		United States	
	ECM	TECM	ECM	TECM	ECM	TECM
ICA Period						
ΔRP_{t-1}	0.528*** (0.067)	0.500*** (0.074)	0.182** (0.083)	0.135 (0.120)	0.389*** (0.079)	0.522*** (0.078)
ΔIP_t	-0.151*** (0.047)	-0.130 (0.049)	0.030 (0.079)	0.069 (0.081)	-0.445*** (0.085)	-0.575*** (0.080)
ΔIP_{t-1}	-0.015 (0.048)	-0.003 (0.049)	0.061 (0.079)	0.127 (0.084)	0.446*** (0.096)	0.447*** (0.088)
Post-ICA Period						
ΔRP_{t-1}^{+}	0.489*** (0.066)	0.576*** (0.063)	0.169 (0.096)	0.250** (0.101)	0.195*** (0.060)	0.182*** (0.060)
ΔRP_{t-1}^{-}	0.171 (0.089)	0.036 (0.106)	0.060 (0.087)	0.017 (0.094)	-0.037 (0.135)	0.019 (0.137)
ΔIP_t^{+}	-0.017 (0.054)	-0.008 (0.053)	0.342*** (0.111)	0.372*** (0.110)	0.366*** (0.099)	0.263*** (0.097)
ΔIP_t^{-}	0.293*** (0.091)	0.239*** (0.084)	0.692*** (0.184)	0.571*** (0.172)	-0.075 (0.160)	0.010 (0.149)
ΔIP_{t-1}^{+}	0.168*** (0.063)	0.230*** (0.057)	0.037 (0.126)	0.127 (0.125)	1.019*** (0.116)	0.944*** (0.111)
ΔIP_{t-1}^{-}	-0.129 (0.084)	-0.196** (0.078)	-0.228 (0.177)	-0.358** (0.167)	-0.543*** (0.154)	-0.447*** (0.153)

Figure 1. Monthly international coffee prices and retail coffee prices in France, Germany and the United States, 1980-2009

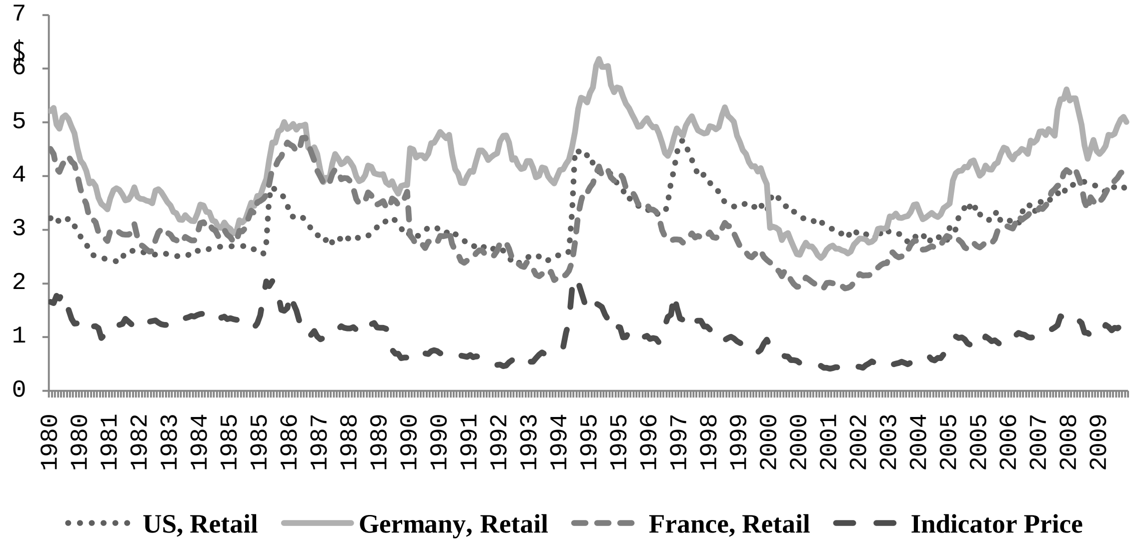
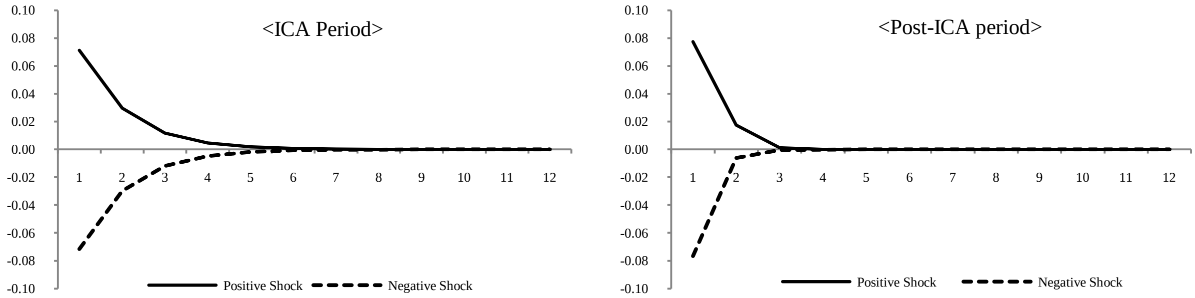
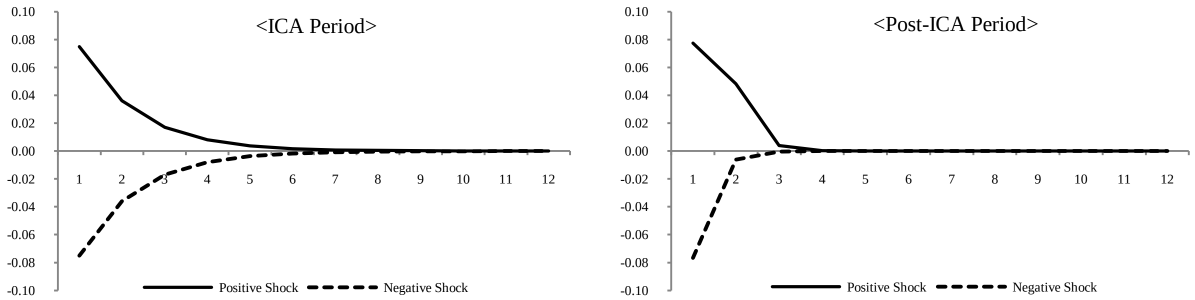


Figure 2. Responses of retail price to the change of international price

A. France



B. Germany



C. United States