Price Transmission Asymmetries and Nonlinearities

in the International Coffee Supply Chain

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Abstract: We examine two distinct and important 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 resulting from the elimination of International Coffee Agreement (ICA) in 1990. Our results confirm the presence of threshold effects in both periods (ICA and post ICA) in all 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 these two countries, 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. Our results suggest differences between the two European countries and the United States. Specifically, our results indicate that changes in international prices influence U.S. retail prices in both periods. Nonlinear impulse response analysis indicates that ICA elimination in 1990 increased the speed of adjustment toward the long-run equilibrium, given a shock in international coffee prices. Our results show that ignoring nonlinearities and asymmetries in price transmission may lead to incorrect assessment of the consequences accruing to the elimination of the ICA.

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

JEL Codes: C32, Q17.

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 prices. 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 asymmetric price adjustment to overcome the limitations of Wolframm 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 type of price frictions, may result in nonlinear adjustments toward the long-run equilibrium (Meyer 2004).

We focus on these two distinct and important dimensions of price transmissions. One is that price adjustments may respond differently to positive and negative exogenous (i.e. price transmission asymmetries). The other is that there may be thresholds beyond which long-run adjustments occur (i.e. nonlinearities in price transmission). To do this, we employ an error correction model with threshold proposed by Tong (1983) and later extended by Balke and Fomby (1997). This threshold approach allows us to model nonlinear price adjustments toward the long-run equilibrium based on different regimes which are separated by estimated threshold values. In addition, we extend the threshold error correction model to incorporate asymmetric short-run responses to an exogenous shock.

In this paper we examine price transmission from international to retail coffee bean prices in three largest coffee importing countries (France, Germany and the United States) and examine revisit the implications of the International Coffee Agreement (ICA) elimination in 1990. The primary change brought by this change was the end of the export quota system limiting coffee exports to major importing countries.

In Figure 1 we show monthly international price and retail coffee prices in each country during the period 1980 to 2009. The figure suggests that the termination of export quota system may have affected the three countries in different ways in terms of the response to changes in international prices. The retail prices in 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 of international prices in the early 1990s, while retail prices of France and the United States decrease following the trend of international prices, retail prices in Germany show experienced high volatility and stayed high relative to the other two importing countries.

[Figure 1 here]

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 taking place in the German retail sector in the late 1990s and early 2000s (Körner 2002; McLaughlin 2006). Retail pricing in France is more regulated than in Germany and the United States.[1]

[Table 1 here]

A number of researchers have examined the impact of the ICA elimination and have investigated the impacts at various levels of the coffee supply chains. 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 domestic prices adjusted faster toward the long-run equilibrium in response to shocks in international prices

[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).

during this period. Shepherd (2004) examines the impact of ICA's elimination on price transmission using a vector autoregression (VAR) model. The results indicate that elimination of the export quota system did not lead to improved price transmission because of market power of 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 1990-2006 employing error correction model. They find no evidence of long-run price transmission asymmetries. However they provide the evidence of short-run asymmetries with substantial differences among countries.

In this study, we revisit the implication of the ICA elimination on price transmission between international prices and retail prices in France, Germany and the United States taking into account possible nonlinearities and asymmetries. We show that ignoring these two features of the price transmission process may lead to incorrect impact assessments of the ICA elimination. 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 a ATECM representation model 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 the standard ECMs.

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-integration relationship between two price series is given as:

$$RP_t - \sigma_0 - \sigma_1 IP_t = \varepsilon_t , \qquad (1)$$

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:

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

The Heaviside indicator function, I_t, is

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

where θ represents a threshold value by which movements toward the long-run equilibrium relationship are asymmetric depending on the regime; 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 adjusting overwhelms the cost generated by adapting new price due to the presence of transaction cost or other sources of price frictions (Balke and Fomby 1997). That is, the error correction mechanism operates only when deviations from long-run equilibrium exceed a critical range [θ and -θ][2]. The inside regime, between θ and -θ ,can be defined as a “band of non-adjustment” or “neutral band” in which no adjustments take place, due to small deviations that do not trigger price responses because of the costs of response may be higher than the benefits (Goodwin & Piggot 2001; Meyer 2004; Meyer & von Cramon-Taubadel 2004).

[2] As argued in Hansen and Seo (2002) and Meyer (2004), threshold co-integration model with three regimes divided by two different threshold values, θ_1 and θ_2 is often criticized by the fact that no exists any significance test for two threshold values in multivariate error correction model. Hansen and Seo (2002) suggest absolute value of error correction term in line with Balke and Fomby (1997) for significance test of threshold effect.

Tsay (1998) suggests a nonparametric approach to identify possible nonlinearities 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 delay (Goodwin and Holt 1999; Goodwin and Piggott 2001). If nonlinearities in the error correction term are observed, we 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 sum of squared 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 Square 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 threshold values can be estimated. Since it is likely that international and retail prices are 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 $I_t = \begin{cases} 0 & if\ |ECT_{t-d}| \leq \theta \\ 1 & if\ |ECT_{t-d}| > \theta \end{cases}$. $\Delta z_{k,t-i}\ \forall\ k = 1,2$ are the identifying variables for the short-run dynamics in retail and international price equations, respectively. To investigate the possible short-run asymmetries in price transmission, theΔRP_{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 defined as in the systems of equations (4)-(5).

In this study, we follow a systematic approach to determine the appropriate specification to assess impacts of the elimination of the ICA. 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 quote 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 is from the International Coffee Organization (ICO). Retail prices of roasted coffee and international composite prices are presented as the 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) as the identification variables the retail price equations in France and Germany, respectively. In the U.S. 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) because weather patterns in this country influence international prices. Descriptive statistics of these data are presented in Table 2.

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[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.

[Table 2 here]

Results

Test of Integration and Co-integration

We first test the time-series properties of the price data. We first 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.

[Table 3 here]

We follow Johansen's (1992a, 1992b, 1995) approach to test whether our international and retail price series are co-integrated. These procedures identify 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.

[Table 4 here]

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 test by Tsay (i.e., choosing *d* that maximizes the *F* statistic (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 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 $[-\theta, \theta]$) is decreases during the post-ICA period 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.

[Table 5 here]

Given that 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 for both periods (ICA and post-ICA). Our results show that there is no evidence of asymmetries in France and very modest evidence of asymmetries in Germany and in the United States, during the ICA period. Measures of model goodness-of-fit (AIC and SBC) presented in Table 7 provide additional support to this finding. During the ICA period, the values of AIC and SBC for the symmetric model specifications are lower than their asymmetric model counterparts, in all three countries. The implication is that a symmetric threshold error correction model specification is more 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.

[Table 6 here]

[Table 7 here]

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. 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 range). For France and Germany, the estimated parameters for outside regime are negative in both countries, as predicted by theory, and are statistically significant. 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 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 France, which implies Germany went through more dramatic change in price transmission after the elimination of the export quota system. Additionally, the parameter estimates suggest that there are no significant adjustments in interior regime for both countries, consistent with the existence a “band of no adjustment.” In contrast to France and Germany, for the United States the speed of adjustment accelerates after the collapse of the ICA. The parameter estimated 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 expectation. These results show different long-run behaviors of price adjustments between France, Germany and the United States.

We examine the short-run dynamics through analyzing contemporary and lagged parameters of both international price and identification variables such as exchange rate and consumer price indexes. Our results presented in Table 8, 9 and 10 show that for France, while the change of contemporary and lagged international price does not affect the retail price in ICA period, for contemporary negative international shock of $1, retail price decreases by $0.24 in post ICA period. Furthermore, a $1 increase of lagged international price leads to $0.23 increase in retail price. But contrary to expectation, a $1 decline leads to $0.20 increase in retail price. Exchange rate appears to have a significant effect on retail price in France. However, the responses to the change of exchange rate are different depending on ICA regime. While response to the positive change of exchange rate becomes stronger, response to the negative change becomes weaker in post ICA period.

Our German results show while the change of international price has no influence on retail price in ICA period, the variation of international price in post ICA period affect retail price both positive and negative change. Based on the presence of asymmetries in price transmission in post ICA period, a $1 increase of contemporary international price causes a $0.37 increase of retail price. On the other hand a $1 decrease of contemporary international price leads to $0.57 decrease of retail price. These short-run behaviors of Germany appear to have different shape, compared to France. The effect of international price on retail price is much more responsive for the case of price decrease than the case of price increase.[4] Similar to France, the response for exchange rate fluctuation becomes faster for positive change and slower for negative change in post ICA regime. Specifically, while one unit increase of exchange rate denoted by domestic currency (here German Mark) is associated with a $1.42 increase in retail price given that retail prices are converted into US dollars. However, the effect of exchange rate is much stronger in Germany than France which reflects the differences of unit currency values with US dollars between France and Germany. Our estimates imply that in pre-liberalized period of export the shock of international coffee price did not affect the retail market in France and Germany. But the change of exchange rate has a significant impact on the retail coffee prices in ICA period.

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

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[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. In economic sense, these results can be interpreted that the shock of international price may affect the retail price after one time period. The degree of price transmission for lagged international price change becomes stronger after the elimination of ICA. Based on the presence of asymmetries in post ICA period, positive change of international price has more impacts on the retail price from \$0.45 to \$0.94. Additionally, as an identification variable, the estimates of consumer price index for foods and beverages appear to be not statistically significant. The estimate results of the United States for post ICA period are in sharp contrast to Germany's results. In the United States, while increases of contemporary international prices have an impact on the retail price, decreases appear to have no effect on the retail price. However, in Germany, negative changes (price decreases) of international price seem to have a greater effect on retail price than positive change as stated above. 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. On the other side, the results of France also show different short-run behaviors from the United State's patterns. For a contemporary international price shock, while negative change has only effect on France's retail price, positive change does for the United States. In addition, while a \$1 increase of international price prior to one time period leads to \$0.23 increase in retail price of France, a \$1 increase of lagged international price results in \$0.94 increase of retail price of the United States.

[Table 8 here]

[Table 9 here]

[Table 10 here]

Symmetries and asymmetries of price transmission in ICA period and post ICA period are confirmed through the impulse response analysis. Considering the nonlinear characteristics of model, we employ Potter's (1995) approach. Potter (1995) points out while linear impulse responses model is independent from the history of time series and the sign and magnitude of shock have no effect on the time path of responses, in the case of the nonlinear model, the effect of shock of error terms on the time path of responses is affected by the magnitude and sign of the history of shock, that is history-dependent (Goodwin and Hold 1999; Abdulai 2002; Enders 2004). Potter (1995) suggests the modified representation of 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 price of each country to one positive and negative standard deviation shock of international price. In ICA period, the shock from international price has a symmetric effect on retail price in each country, regardless of the sign of shock. However, the responses to shock are slightly different across countries. While the responses of retail price in Germany die down after 8 months, those of the United States vanish after 5 months for positive and negative shock. Furthermore, while the responses to shock are rapidly dampened after one month in the United States, the responses of Germany and France are gradually diminished. In marked contrast to ICA period, the responses of retail prices to international shock seem to be asymmetrically affected in post ICA period. More specifically, while positive shock of international price persists by 4 months in Germany, negative shock disappears after 3 months. On the other hand, the responses of the U.S. retail price to shock demonstrate opposite results from Germany. While positive shocks are mostly absorbed within 2 months, negative shocks continue until 3 months. In case of France, negative shocks are more rapidly died out than positive shock in post ICA period even though both shocks last identically by 3 months. Moreover, we find in contrast to ICA period where shocks more last, the shocks from international market disappear faster in post ICA period regardless of the sign of shocks.

[Figure 2 here]

Finally Table 11 and 12 compare parameter estimates for short-run dynamics between threshold error correction model and standard error correction model[5] to confirm whether ignorance of the potential nonlinearity in error correction process cause the biased results. In ICA period where symmetric model is more fitted, deviations from equilibrium are generally faster adjusted in threshold model for all countries than standard model where nonlinearities are disregarded. In post ICA period where asymmetric model is adopted, while the estimates of threshold model for France show faster adjustments on the whole, the estimates of Germany and the United States show the mixed results which depend on variables.

[Table 11 here]

[Table 12 here]

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 nonlinearities 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 the existence of nonlinearities in the co-integrating vectors, 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

[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.

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 becomes faster during the post-ICA period in the United States.

The estimated threshold range became smaller in the post-ICA period, particularly for Germany. This indicates that retail prices became more responsive to changes in international prices, even if the change in the latter were of small magnitude. In the short-run, our parameter estimates suggest that during the ICA period, changes in international prices did not influence retail prices in France and Germany. In contrast, changes in international prices influenced retail prices in the United States, independent of the period. 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 nonlinearities 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 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	U.S.A
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 Headquaters 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 cointegration (*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**
U.S.A	**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**	**U.S.A**
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] ($\boldsymbol{\rho}_1^{(1)} = \boldsymbol{\rho}_1^{(2)}$)	ICA	2.721*	10.262**	23.641**
	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	U.S.A
$\Delta RP_{t-1}^{+} = \Delta RP_{t-1}^{-}$	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. Model fitness

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
U.S.A	TECM	ATECM	TECM	ATECM
AIC	-1269.25	-1252.46	-2327.11	-2391.95
SBC	-1224.92	-1180.42	-2271.55	-2301.67

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

Variables	ICA period	Variables	Post ICA period
	STECM		ATECM
Constant	-0.001 (0.004)	*Constant*	-0.001 (0.007)
$ECT_{t-1}^{(1)}$	-0.048*** (0.010)	$ECT_{t-1}^{(1)}$	-0.043*** (0.009)
$ECT_{t-1}^{(2)}$	-0.026 (0.021)	$ECT_{t-1}^{(2)}$	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)
Δz_{t}	-0.461*** (0.020)	Δz_{t}^{+}	-0.559*** (0.044)
		Δz_{t}^{-}	-0.423*** (0.039)
Δz_{t-1}	0.179*** (0.042)	Δz_{t-1}^{+}	0.013 (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 of Germany (Retail price equation)

Variables	ICA period	Variables	Post ICA priod
	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 of U.S.A (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 (ICA period)

Variables	France		Germany		U.S.A	
	ECM	**TECM**	**ECM**	**TECM**	**ECM**	**TECM**
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)
Δz_t	-0.268*** (0.042)	-0.461*** (0.020)	-1.217*** (0.175)	-1.415*** (0.100)	1.387 (2.839)	0.820 (2.365)
Δz_{t-1}	-0.195*** (0.040)	0.179*** (0.042)	-0.217 (0.163)	0.140 (0.202)	-0.343 (2.459)	0.129 (2.321)

Table 12. Comparison estimates from TECM with ECM (post ICA period)

Variables	France		Germany		U.S.A	
	ECM	TECM	ECM	TECM	ECM	TECM
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)
Δz_{t}^{+}	-0.409*** (0.057)	-0.559*** (0.044)	-2.630*** (0.345)	-2.464*** (0.304)	4.451** (2.058)	2.076 (1.699)
Δz_{t}^{-}	-0.311*** (0.053)	-0.423*** (0.039)	-1.608*** (0.336)	-1.405*** (0.278)	-1.310 (7.001)	-0.256 (6.828)
Δz_{t-1}^{+}	-0.163*** (0.046)	0.013 (0.067)	0.198 (0.276)	-0.463 (0.372)	-2.902 (2.550)	1.087 (1.695)
Δz_{t-1}^{+}	-0.100** (0.050)	0.246*** (0.046)	0.170 (0.318)	0.113 (0.309)	-0.147 (2.384)	0.702 (6.817)

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

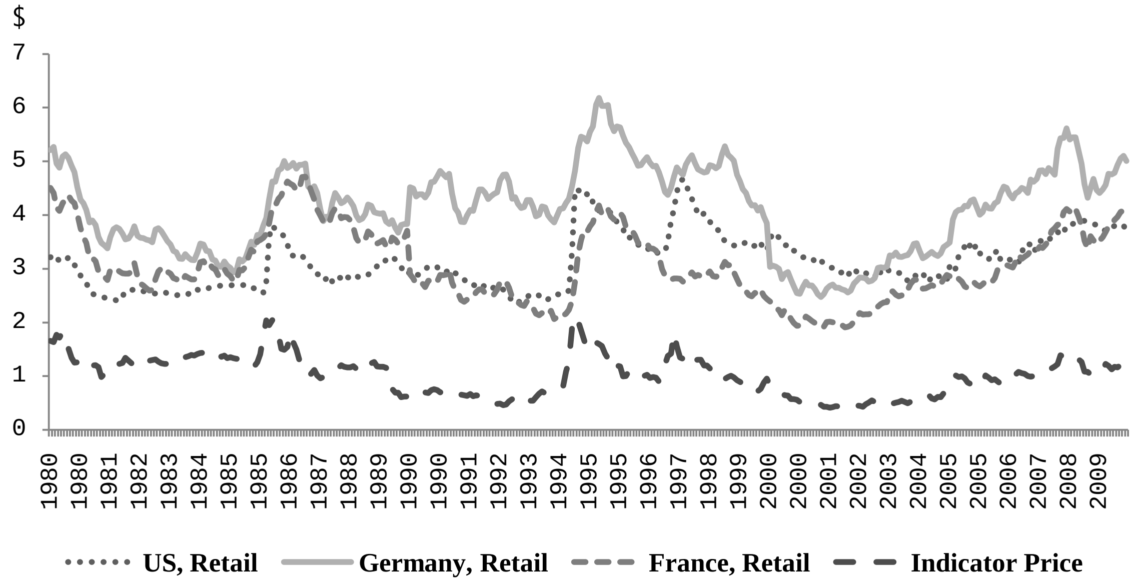
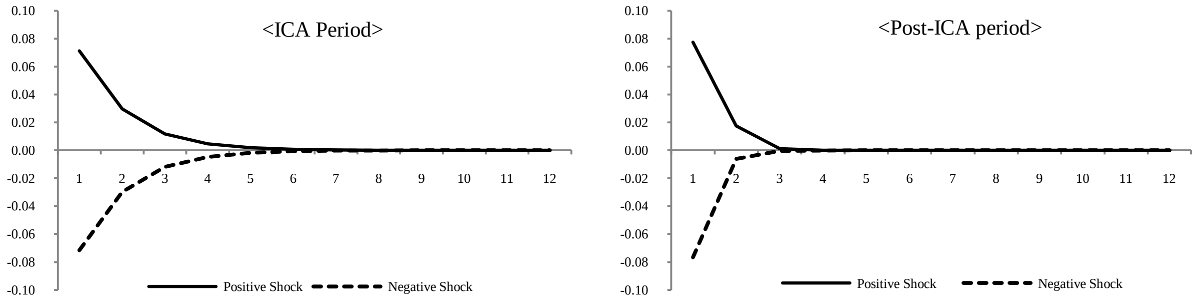
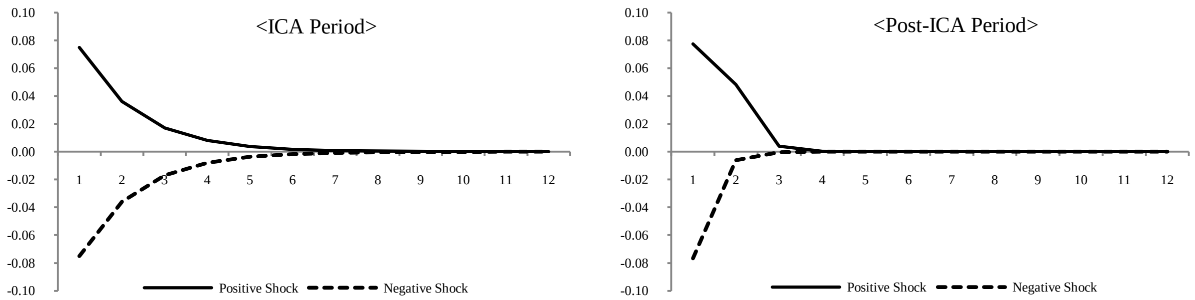


Figure 2. Responses of retail price to the change of international price

A. France



B. Germany



C. U.S.