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Asymmetric Price Transmission in the U.S. Soybean Exports

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ABSTRACT

The purpose of this study was to estimate asymmetric price transmission in the U.S. soybean exports and establish an indicator of the asymmetric price transmission that will make it possible to analyze its movement. The asymmetric price transmission from U.S. domestic soybean prices to export prices is estimated using repeated threshold autoregressive model which takes into consideration possible changes in asymmetric price transmission or changes in the parameters of the threshold autoregressive model. Asymmetric price transmission can be positive or negative: Positive (negative) asymmetry indicates that the squeezed margin between domestic and export prices are restored more quickly (slowly) than the stretched margin. According to the estimation results, the conclusion is that the price transmission was positively asymmetric from 1967 to 1977 and then became neutral or negative and remained so until 1988. Then the asymmetry transitioned back to positive until the latter half of the 1990s and became negative again afterwards. The U.S. has lost long-term excess profits from the export of soybeans through asymmetric price transmission, possibly due to the emergence of Brazil and Argentina as major soybean exporters. The contributions of this study include that it estimates repeated threshold autoregressive model using U.S. soybean prices with various numbers of subsamples and that it establishes an indicator of asymmetric price transmission by accumulating the repeated threshold autoregressive estimation results.

Key words: Asymmetric price transmission, TAR model, cointegration, time-varying parameter, market power

INTRODUCTION

Asymmetric Price Transmission (APT) is a popular research topic, as mentioned in the survey study by Meyer and von Cramon-Taubadel (2004). Price transmission is said to be asymmetric if the speed of adjustment of the output price is different after the input price increases or decreases. Following the method of Enders and Granger (1998) many empirical studies have been conducted using a Threshold Autoregressive (TAR) model to estimate APT with cointegration tests.

Previous empirical studies of APT that use the TAR model for agricultural products include that of Abdulai (2002) who analyzed the Swiss pork market and showed that retail prices adjust more rapidly when producer prices increase than when they decrease, which indicates the downward rigidity of price transmission. Ghoshray (2002) estimated APT for wheat export prices in major wheat-producing countries and detected multiple instances of APT using the momentum TAR (M-TAR) model. Gonzales *et al.* (2003) analyzed the marine products market in France and identified downward rigidity in price transmission from the landing prices for cod to their retail prices and upward rigidity from the import prices for salmon to their retail prices.

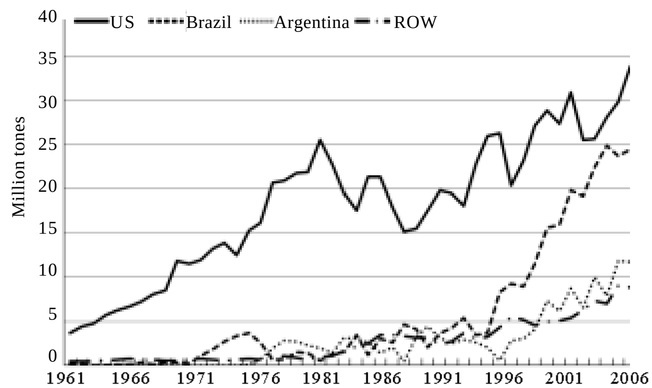

Fig. 1: Soybean exports from the U.S., Brazil, Argentina and the rest of the world. Source: FAOSTAT

Hassan and Simioni (2001) considered price transmission from shipping prices to retail prices in the vegetable market in France, showing the downward rigidity of price transmission in the chicory market and the upward rigidity of price transmission in the tomato market.

One of the problems associated with these previous studies that employ the TAR model is the sample period selected is rather arbitrary. Some studies mention potential structural change that and divide the sample period into subsamples, after which they estimate the TAR model using each subsample (Enders and Siklos, 2001). Because the choice of sample period affects the values of the parameters in a given model, TAR estimation should be conducted based on the sensitivity of the different sample sizes (Bermejo *et al*., 2011).

The purpose of this study was to estimate APT for U.S. soybean exports and establish an indicator for APT so as to analyze movement via the TAR model which takes into consideration the possible changes in APT or changes in the parameters of the TAR model.

Figure 1 shows the soybean exports of three major exporters and the Rest of the World (ROW) using the data from Food and Agriculture Organization Corporate Statistical Database (FAOSTAT). U.S. exports increased from the 1960s until the early 1980s and they reached 25 million tons in 1982. However, exports decreased to 15 million tons in the 1980s. Since the 1990s, U.S. exports have increased with some fluctuations and they reached 34 million tons in 2008. The U.S. has been the major exporter of soybeans to date since the 1960s and particularly dominated the world soybean export market until the middle of the 1990s. In the latter half of the 1990s, however, exports from Brazil surged from 5 to 25 million tons, matching those from the U.S. In the late 1990s, exports from Argentina also increased from 3 to 12 million tons.

MATERIALS AND METHODS

TAR model: Below, the TAR model based on Enders and Siklos (2001) is explained. Denote p_{it} and p_{ot} as the input and output prices at time t. The long-run relationship between p_{it} and p_{ot} is represented using Ordinary Least Squares (OLS) regression as follows:

$$p_{ot} = \alpha + \beta\, p_{it} + \mu_t \quad (1)$$

where, α and β are parameters and μ_t is the disturbance term which may be serially correlated. According to Engle and Granger (1987), if p_{it} and p_{ot} are part of a non-stationary process and Δp_{it}

and Δp_{ot} are part of a stationary process (that is, if they are first-difference stationary (I(1)) variables), then Eq. 1 may indicate a spurious regression. If the residual series $\{\hat{\mu}_t\}$ is stationary, however, then p_{it} and p_{ot} are said to be cointegrated. Therefore, it is necessary to conduct unit root tests and cointegration tests on p_{it} and p_{ot} to avoid a spurious regression.

In a TAR model, a cointegration test is performed using $\{\hat{\mu}_t\}$ from Eq. 1 in the following equations (Enders and Siklos, 2001):

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{T}\gamma_i\Delta\mu_{t-i} + \varepsilon_t \tag{2}$$

$$I_t = \begin{cases} 1 & \text{if} \quad \mu_{t-1} \geq \tau \\ 0 & \text{if} \quad \mu_{t-1} < \tau \end{cases} \tag{3}$$

where, I_t is the Heaviside indicator function and τ is the super-consistent estimator of threshold μ_{t-1} calculated following Chan (1993). ε_t is the white noise disturbance term and satisfies the following conditions:

$$E(\varepsilon_t) = 0,\ E(\varepsilon^2_t) = \sigma^2,\ E(\varepsilon_t\varepsilon_j) = 0\ (t \neq j) \tag{4}$$

The necessary and sufficient condition for $\{\hat{\mu}_t\}$ to be stationary is as follows (Petrucelli and Woolford, 1984):

$$\rho_1 < 0,\ \rho_2 < 0 \text{ and } (1 + \rho_1)(1 + \rho_2) < 1 \text{ for any } \tau \tag{5}$$

T is the lag order that satisfies the conditions of Eq. 4, 5 and minimizes the BIC (Bayesian Information Criteria).

A cointegration test is performed by testing $\rho_1 = \rho_2 = 0$; i.e., if the null hypothesis of $\rho_1 = \rho_2 = 0$ is rejected, then p_{it} and p_{ot} are said to be cointegrated. APT can be tested in the same model to compare the absolute values of ρ_1 and ρ_2. If $\rho_1 = \rho_2$ is rejected and $|\rho_1|<|\rho_2|$, then the negative discrepancies from the equilibrium error adjust more rapidly than the positive discrepancies. The implication is that a shock that decreases the margin adjusts more rapidly than a shock that increases the margin. That is, price transmission exhibits downward rigidity, called positive APT. However, if $\rho_1 = \rho_2$ is rejected and $|\rho_1|>|\rho_2|$, then the positive deviations adjust toward the equilibrium error more rapidly than do the negative deviations. A shock that increases the margin adjusts more rapidly than does one that decreases the margin. This results in negative APT which indicates upward rigidity.

The M-TAR model is the same as in Eq. 2, 3 except that μ_{t-1} in Eq. 3 is replaced with $\Delta\mu_{t-1}$. The TAR model and M-TAR model correspond to the two asymmetric adjustment processes, Deepness and Steepness (Sichel, 1993). In both models, however, $|\rho_1|<|\rho_2|$ indicates positive APT and $|\rho_1|>|\rho_2|$ indicates negative APT.

Repeated TAR: One of the problems with the TAR model based on Enders and Siklos (2001) is that the parameters may change depending on the sample period. One possible way to overcome this problem is to conduct a structural change test. A structural change test may be conducted on the residual series $\{\hat{\mu}_t\}$ in Eq. 1 using Bai and Perron (2003) method. The sample is divided into subsamples according to the test results and then the TAR model is estimated using each subsample. It seems reasonable to use $\{\hat{\mu}_t\}$ to test possible structural changes in APT but there is no in-depth theoretical background indicating that it is appropriate to test structural changes using $\{\hat{\mu}_t\}$. Additionally, if a subsample is too short to estimate the TAR model, only limited or no estimation results may be obtained.

Another possible way to count the potential time-varying parameters of the TAR model is through repeated TAR. In the repeated TAR methodology, regressions are repeated by increasing the starting and ending periods one by one using a given subsample size. For example with a sample size of 100, the regression is conducted using periods 1-100 of the dataset and then using periods 2-101, 3-102 and so on. This methodology has rarely been applied to the TAR model. Bermejo *et al.* (2011) used recursive estimation with a different type of TAR model. In this study, repeated TAR estimation is conducted by setting the subsample size at 100 and increasing the starting and ending periods by one, as in the example above. A subsample size of 100 is large enough to eliminate the problem of small sample bias. Subsample sizes of 150 and 200 are also used in the estimation below but a large subsample may have a bias of possible structural changes. The results achieved using 150 and 200 subsamples are referred to in the discussion. Furthermore, for each subsample estimation, a cointegration test is conducted and whether conditions (Eq. 4, 5) are satisfied is tested. If these conditions are satisfied and it is confirmed that the variables are cointegrated, then the optimal lag length using BIC is selected. Under these conditions ρ_1 and ρ_2 are stored in each subsample estimation.

Using this repeated TAR estimation, time-varying ρ_1 and ρ_2 are obtained and the changes of these parameters emerge. This information, however, is not time-specific. That is, it is not possible to determine the value of the parameters in each period. To determine a specific value for a parameter in each period, the ρ_1 and ρ_2 obtained should be accumulated during each period. Let APT_{it} be a value of APT in subsample i in period t. APT_{it} is the same value for the same i and the value depends on whether the APT is significantly positive, significantly negative, or insignificant as follows:

$$APT_i = \begin{cases} 1/w & \text{if APT is significantly positive at least at the 10\% level,} \\ -1/w & \text{if APT is significantly negative at least at the 10\% level,} \\ 0 & \text{otherwise.} \end{cases} \tag{6}$$

where, w is the subsample size. The value of APT at t (APT_t) is derived by averaging the sum of APT_t using the number of subsample estimations (denoted by s) for each t as the weight. Because the subsample size is fixed, s depends on t. When the total sample size is T (T>w), s = min (t, w,T-t + 1). For example, when w = 100 and T = 525, s = 10 for t = 10, s = 100 for t = 200 and s = 26 for t = 500.

RESULTS

Data: The data used for input prices were the average U.S. soybean prices received, obtained from USDA-NASS. The output prices were the average U.S. soybean export prices for all countries obtained from USDA-GATS. Both datasets included monthly data from January 1967 to September 2010. The total sample size was 525. These price series were transformed into natural logarithmic form as is always done in empirical analyses of TAR models (Ben-Kaabia and Gil, 2007).

Table 1: Unit root test results

	ADF		PP		KPSS	
Price	Level	1st diff.	Level	1st diff.	Level	1st diff.
Domestic	-2.48 (7)	-9.45 (7) ***	-2.52 (6)	-19.18 (11) ***	1.10 (17) ***	0.04 (7)
Export	-2.28 (1)	-16.26 (0) ***	-2.03 (2)	-16.00 (8) ***	1.24 (17) ***	0.06 (1)

Values are statistics for each unit root test. Values in parentheses for the ADF tests indicate lag order based on the Akaike Information Criteria (AIC). Values in parentheses for the PP and KPSS tests indicate Newey-West bandwidth using the Bartlett kernel. In the equations for all tests, the intercept (no trend) is included. ***, ** and * represent 1, 5 and 10% significance, respectively

Unit root tests: To test whether the price series are I (1) variables, Augmented Dickey-Fuller (ADF) tests, Phillips-Perron (PP) tests and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests were conducted. The results are shown in Table 1. Based on the ADF and PP tests, the null hypothesis that the series have a unit root is not rejected for the level series but it is for the first-difference series. For the KPSS tests, the null of the stationary series is rejected for the level series but is not for the first-difference series. Therefore, both price series can be said to be I (1) variables. The test statistics shown in Table 1 are those achieved by including intercepts (but not trends) in the test equations. Similar results were obtained by including intercepts and trends in the test equations, which are not shown here to save space.

TAR and repeated TAR results: According to the methodology mentioned above, TAR and repeated TAR estimations were conducted. The results of the TAR estimations conducted using the total samples are shown in Table 2. The lag order for each model was determined by minimizing the BIC when the conditions of Eq. 4, 5 are satisfied. Based on both the TAR model and the M-TAR model, the conclusion is that p_{it} and p_{ot} are cointegrated because the Φ statistics in both models are much larger than those at the critical 1% significance level in Enders and Siklos (2001). According to that study, the Φ statistics at a 1% significance level for 250 observations and four lagged changes are 10.18 for TAR and 8.47 for M-TAR. As shown in Enders and Siklos (2001), the Φ statistic tends to decrease as the number of observation increases and it tends to increase as the number of lags increases. Therefore, it is reasonable to conclude that the null of no cointegration is rejected at the 1% level in both the TAR and the M-TAR model. In fact, the Φ statistics obtained in this study are far beyond the values shown above. Although the null hypothesis that $\rho_1 = \rho_2$ is not rejected at the 10% level based on the TAR model, the null is rejected at the 10% level based on the M-TAR model and $|\rho_1|<|\rho_2|$. It follows that in the total sample, price transmission from U.S. domestic soybean prices to export prices is symmetric in the TAR model and that there is positive APT in the M-TAR model.

The estimation results produced using the total sample remain inconclusive because the results achieved using the TAR and M-TAR models were different. Next, the estimations were repeated using a subsample with a fixed size to trace the changes in $|\rho_1|$and $|\rho_2|$. The results are shown in Fig. 2. Fig. 2a depicts the changes in $|\rho_1|$ and $|\rho_2|$ based on repeated TAR estimations and Fig. 2b depicts those of repeated M-TAR estimations. In the repeated TAR results, $|\rho_1|<|\rho_2|$ from period 1 to 70, $|\rho_1|>|\rho_2|$ from 70 to 120, $|\rho_1|\approx|\rho_2|$ from 120 to 250, $|\rho_1|<|\rho_2|$ from 250 to 380 and $|\rho_1|>|\rho_2|$ from 380 to 426. In the repeated M-TAR results, $|\rho_1|<|\rho_2|$ from period 1 to 70, $|\rho_1|>|\rho_2|$ from 70 to 130, $|\rho_1|\approx|\rho_2|$ from 120 to 150, $|\rho_1|>|\rho_2|$ from 150 to 220, $|\rho_1|\approx|\rho_2|$ from 220 to 370 and $|\rho_1|>|\rho_2|$ from 370 to 426. In both results, it seems that positive APT has gradually become negative, with several upward and downward (positive and negative) movements.

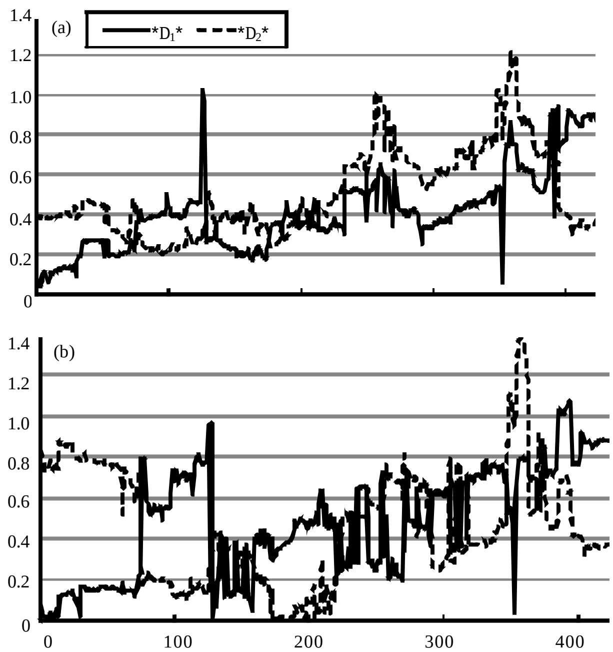

Fig. 2: Changes in $|\rho_1|$ and $|\rho_2|$ based on repeated TAR (a) and repeated M-TAR (b)

Table 2: TAR results using total sample

Model	ρ_1	ρ_2	lags	Φ	Asym.	Q (6)	BIC	τ
TAR	-0.361 ***	-0.289 ***	2	35.29***	1.13	4.90	-4998.4	0.051
	(0.058)	(0.044)			[0.29]	[0.56]	3rd	
M-TAR	-0.271 ***	-0.394 ***	2	36.51***	3.28 *	4.26	-5000.5	-0.031
	(0.044)	(0.058)			[0.07] +	[0.64]	3rd	

ρ_1 and ρ_2 are the adjustment coefficients in Eq. 2. 'lags' is the lag length in Eq. 2. Φ is the F statistic for the test of the null hypothesis $\rho_1 = \rho_2 = 0$. The rejection regions are based on Enders and Siklos (2001). Asym. is the F statistic for the test $\rho_1 = \rho_2$. + indicates significant positive APT. Q (6) represents the Q statistics from the Portmanteau test for white noise, whose null hypothesis is that the error term is white noise up to 6 lags. In BIC, 3rd indicates that the value is the 3rd smallest. τ is the threshold in Eq. 3. For each result, the values in ( ) denote standard errors and the values in [ ] denote p values. ***, ** and * represent 1, 5 and 10% significance, respectively

Although the changes in ρ_1 and ρ_2 are shown, the results above do not indicate the movement of APT itself and there is no information about dates. To overcome these problems, the indicator for APT (APT_t) was calculated according to the previously mentioned method. Fig. 3 shows the indicator for both the TAR and the M-TAR results. Based on the repeated TAR, APT_t was positive from January 1967 to June 1977, neutral until April 1988, positive again until the end of 2000 and then negative to date. Based on the repeated M-TAR, APT_t was positive from January 1967 to April 1977 but then became negative after that and remained so to date, although it increased from January 1988, became neutral in the middle of 1993 and then was slightly negative until the end of 2003. Although there are some discrepancies between TAR and M-TAR, the movement and the direction at any given time are quite similar.

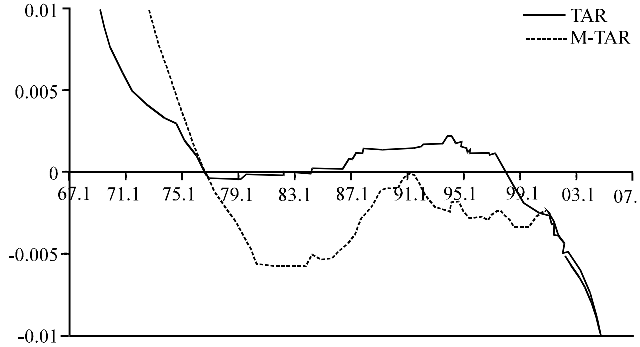

Fig. 3: Changes in APT

DISCUSSION

The findings based on the empirical analysis above suggest that the U.S. has lost long-term excess profits from soybean exports because the positive APT changed to the negative APT. As Meyer and von Cramon-Taubadel (2004) have suggested, most previous studies of APT have focused on the relationship between market power and positive APT. Hence, the U.S. might have enjoyed market power related to soybean exports until the middle of the 1970s and lost that power in the middle of the 1990s after the surge in soybean exports from Brazil and Argentina. In this study, however, only APT was estimated and an estimation of market power itself remains to be conducted. Because the explanations of the relationship between positive APT and market power lack a rigorous theoretical foundation (Meyer and von Cramon-Taubadel, 2004), it is difficult to conclude that there is a clear relationship between changes in APT and changes in U.S. market power.

The empirical findings in the late 1970s through the 1980s surprisingly correspond to the conclusion in Pick and Park (1991) where the U.S. had no market power in soybean exports over any major importers except for the Netherlands in 1978-1988 periods. The findings in the 1990s through the 2000s in this study are also close to the conclusion in Song *et al.* (2009) which showed that import companies in China had more market power than did exporters in the U.S. from 1999 to 2005 using the two-country partial equilibrium model. Because there are no other studies regarding market power in the U.S. soybean exports, it is not possible to compare previous researches with the empirical results of this study during the periods when the U.S. had market power. However, it may be quite mesleading to conclude that the U.S. had no market power according to the previous studies, because they did not consider possible changes of the parameters of market power and the data used in their studies are shorter than those in this study.

Although the results above were obtained from the estimations using 100 subsamples, repeated TAR estimations were also conducted with 150 and 200 subsamples. The movements of APT in the TAR and M-TAR models were quite similar to those in the 100 subsample case, although APT_t was almost 0 in the 1960s and 1970s in the TAR model. However, the p values from the F tests for the null of $\rho_1 = \rho_2$ in these periods were very close to 0.1, especially in earlier periods, which indicates that the results are not very different from those achieved using 100 subsamples. This is one of the issues associated with using an APT indicator such that $APT_t = 0$ when the p values are more than 0.1.

CONCLUSION

This study estimated APT for U.S. soybean exports using TAR and repeated TAR models. The conclusion is that APT was positive from 1967 to 1977 and became neutral or negative until 1988. After 1988, APT moved back toward positive until the latter half of the 1990s and became negative afterwards. The U.S. lost long-standing excess profits from soybean exports, possibly because of the emergence of Brazil and Argentina as major soybean exporters. The contributions of this study include an estimation of repeated TAR using U.S. soybean prices with various numbers of subsamples and the establishment of the APT indicator by accumulating the repeated TAR estimation results. Meanwhile, there is room to improve the APT indicator and further related studies should be conducted in the future.

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