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Testing Possible Causes of Asymmetric Price Transmission Behavior of Major Importers of U. S. Wheat

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Abstract

This study deals specifically with the international transmission of wheat prices wherein the effect of prices in one market impacts the prices of another. Specifically, it shows that import prices in some countries respond in an asymmetric fashion to changes in the export prices of U.S. wheat. Our results indicate that market concentration in the importing country influences price asymmetry and amount of price variability sends a sufficient clear signal to market participants. We also find that the 2008 financial and food price crisis changed the degree of asymmetry in most of the countries studied in this paper.

JEL Codes: F12, F14, Q13, Q17

Testing Possible Causes of Asymmetric Price Transmission Behavior of Major Importers of U. S. Wheat

Introduction

The 2019–2021 supply chain disruption reminded policy-makers and agricultural producers that volatile agricultural prices can, at times, become extreme and significantly disrupt global markets. During such disruptions, some observers claim prices will rise rapidly at first but fall more slowly afterwards (Baquedano and Liefert, 2014; Gopinath and Itskhoki, 2010; Peltzman, 2000). Such asymmetric responses occur whenever there are unforeseen economic shocks and/or market uncertainties in commodity markets.

A number of empirical studies discuss reasons why there may be an asymmetric price response to a change in some exogenous variable (Luckstead, 2018; Hong, 2014; Meyer and von Cramon-Taubadel, 2002; Pick, 1991; Abdulai, 2000; Kinnucan and Forker, 1987). These studies test for differences in a variable's reaction to rising price and the reaction to a falling price. Overall, the literature reveals that price asymmetries exist in many markets (Peltzman, 2000). Despite extensive discussion of asymmetry in commodity markets, there have been few empirical efforts to test which factors cause asymmetric price responses for agricultural commodities, especially in the international market for wheat. And, to our knowledge, the literature has yet to investigate whether a revealed asymmetry may emerge in a market in which there had been no evidence of a previous asymmetry in price response. Our objective is to examine factors affecting importing countries' asymmetric responses to changes in U.S. wheat prices.

To examine this issue, we investigate the asymmetric response of import prices of seven wheat-importing countries to changes in the export price of U.S. wheat to test their influence on the transmission of international prices. While price asymmetry has been previously studied extensively, we introduce several new sources of observed price asymmetry and examine their influence on the transmission of international wheat prices in the international commodity market. In addition, we evaluate the impact of market power, a measure of seller concentration and price uncertainty on the size of the observed asymmetric response in prices.

Price asymmetry studies in the literature cut across several industries and various levels of the supply chain. Some recent studies include assessments of price asymmetry in the food supply chain (e.g., Singh et al., 2022; Panagiotou, 2021; Ramsey et al., 2021; Hong and Lee, 2020; Rezitis et al., 2019; Dong et al., 2018; Abdulai, 2002 and 2000; Azzam**,** 1999) and price transmission across energy markets (e.g., Chen and Sun, 2021; Bragoudakis et al., 2020; Kang et al., 2019; Apergis and Vouzavalis, 2018). These studies, which test for differences in a variable's reaction to a price increase and decrease, have found that asymmetric responses of either quantities or prices to some initial price change are widespread (Peltzman, 2000). Such studies can provide cautionary insights for policy-makers and market participants who rely on price movements to make decisions.

Market power and industry concentration in production, purchasing, or transportation is often speculated to be a key resource for understanding price asymmetries (Rezitis and Tsionas, 2019; Meyer and von Cramon-Taubadel, 2004; Azzam, 1999; Kinnuncan and Forker, 1987). Firms that control a large share of the market react strongly to beneficial price changes and can be slow to react to less beneficial changes. With market share playing a key role, shippers and policymakers want to consider the nature of competition in import markets when formulating policy or setting prices. Peltzman (2000) suggests that loss aversion is prevalent among buyers and can create price asymmetries, which means that shippers and policy-makers may be more interested in making long-term trade agreements with key importers. Other authors identify uncertain price expectations (Kalayci, 2015; Tappata, 2009), adjustment costs (Meyer and von Cramon-Taubadel, 2004), sticky prices (Gopinath and Itskhoki, 2010), and kinks in the demand curve (Bailey and Brorsen, 1989) as possible reasons for asymmetry.

Despite these theoretical explanations concerning asymmetric economic relationships, there have been few efforts to empirically identify factors that create asymmetric price responses (Hong and Lee, 2020; Deltas, 2008; Verlinda, 2008). Thus, the asymmetry literature has had a limited policy impact in the absence of empirical evidence to help inform policy decision making. Furthermore, to our knowledge, the literature has not investigated whether the revealed asymmetry emerged in a market in which there was no evidence of a previous asymmetric response in an economic variable.

Empirically determining the sources of price asymmetry is important because it can provide insights for policymakers and market participants who rely on observing price movements to make decisions. Our study results indicate that 1) import price responses to changes in U.S. port prices are influenced by the direction of price changes at U.S. ports; 2) rising prices can elicit a difference response from importers than falling prices; 3) fluctuating levels of volatility, changes in market concentration, and market shocks (e.g., a financial crisis) influence the magnitude of the asymmetry, as represented by the spread between the import price's responses to a rising and falling U.S. wheat price; and 4) an increase in market concentration leads to large differences in an asymmetric reaction to rising and falling U.S. port prices in some countries, while in other countries the differences in reactions narrow.

Empirical Approach and Model

Toward Testing Asymmetric Transmission

Assume that the following long-run relationship exists:

1) $$V_{jt} = \beta_{j0} + \beta_g P_{gt} + \varepsilon_{jt}$$

where V_{jt} refers to country j's import unit values (IMUV) of U.S. wheat at time t. In this equation, j refers to each of seven importing countries: Colombia, Indonesia, Japan, South Korea, Mexico, the Philippines, and Taiwan.[1] The variable $P_{g,t}$ (g) is the U.S. Gulf price (mostly hard red winter) in time t. The coefficient β_j is a constant representing the average markup of import unit values over the Freight on Board (FOB) price. The variable V_{jt} contains information on tariff, non-trade barriers, unexpected transport cost, price markups or markdowns in the import market, and a host of other factors embodied in the IMUV. The coefficient β_g represents the pass through of port price to import unit value and ε_{jt} is an error term.

If a country's IMUV and U.S. port prices follow a unit root process, Equation 1 could represent the long-run cointegrating relationship between IMUVs and the port prices of U.S. wheat. If so, Equation 1 represents an equilibrium relationship between an IMUV and the U.S. port price. Having an equilibrium relationship would ensure there is no long-run asymmetry between prices and IMUV. Next, we model the short-run disequilibrium equation in differences:

2) $$\Delta V_{jt} = \sum_{i=1}^{I} \beta_{ji} \Delta P_{g,t-i} + \rho\varepsilon_{j,t-1} + u_{jt}$$

where Δ represents time differences so that: $\Delta P_{g,t-1} = P_{g,t-1} - P_{g,t-2}$, and $\rho\varepsilon_{j,t-1}$=$\rho(V_{j,t-1} - \beta_{j0} - \beta_g P_{g,t-1})$. The variable u_{jt} is a normally distributed random error. Inclusion of the lagged error term, $\varepsilon_{j,t-1}$, from Equation 1, creates the well-established error correction model (Johansen and Juselius, 1990). The coefficient ρ represents the rate (or speed) at which IMUV's and port prices return to their equilibrium relationship (when $\rho < 0$). Lagged

Gulf prices are included to account for shipment delays and the effect of trade contracts negotiated in advance of the delivery date.

The most common way to account for asymmetries is to create two price variables to distinguish between rising and falling prices (Hong and Lee 2020; Kinnucan and Forker 1987; Pick et al., 1990). Following Abdulai (2000) and Borenstein et al. (1997), Equation 2 can be rewritten as:

$$\Delta V_{jt} = \sum_{i=0}^{I}\left(\beta_{jui}\Delta P_{g,t-i}^{+} + \beta_{jwi}\Delta P_{g,t-i}^{-}\right) + \rho\varepsilon_{j,t-1} + u_{jt} \tag{3}$$

where the superscript "+" refers to observations where prices have risen from the previous period and the superscript "–" refers to observations where prices have fallen from the previous period. In Equation 3, asymmetric responses for both the coefficients of the exogenous price variables and the adjustment rate coefficients are allowed.

Our Model

To test asymmetry sources, we expand Equation 3 to include two variables that interact with the right-hand side price variables. The first variable is the interaction of a *Hirschman–Herfindahl* market concentration index (Nauenberg et al., 1996) with prices, while the second interaction variable is between the prices and the standard deviation of prices. For the moment, we assume a model without lags, for which we write the IMUV equation for country *j* as:

$$\begin{aligned} \Delta V_{jt} = {} & \sum_{i=0}^{I}\left(\beta_{jui}\Delta P_{g,t-i}^{+} + \beta_{jwi}\Delta P_{g,t-i}^{-}\right) + \alpha_{ju}\left(Hf_{jt} * \Delta P_{gt}^{+}\right) + \alpha_{jw}\left(Hf_{jt} * \Delta P_{gt}^{-}\right) \\ & + \gamma_{ju1}\left(SD * \Delta P_{gt}^{+}\right) + \gamma_{jwn}\left(SD * \Delta P_{gt}^{-}\right) + \rho_{hu}\left(Hf_{jt} * \varepsilon_{j,t-1}^{+}\right) \\ & + \rho_{hw}\left(Hf_{jt} * \varepsilon_{j,t-1}^{-}\right) + \rho_{su}\left(SD * \varepsilon_{j,t-1}^{+}\right) + \rho_{sw}\left(SD * \varepsilon_{j,t-1}^{-}\right) + \rho_{u}\varepsilon_{j,t-1}^{+} + \rho_{w}\varepsilon_{j,t-1}^{-} + e_{jt} \end{aligned} \tag{4}$$

where Hf_{jt} is a Hirschman–Herfindahl concentration index for country *j*, calculated by summing the squared seller shares in market *j* at time *t*. The *Hf* index is used by economists to represent the degree of competition in a particular market (Nauenberg et al., 1996*)* and often serves as an explanatory variable in price markup equations (Azzam, 1997, Lopez et al., 2002). This *Hf* variable changes over time and across importing countries and represents the degree of competition among sellers to market *j*.[2]

The term *SD* in Equation 4 is a 3-period moving standard deviation of the U.S. wheat price representing uncertainty about U.S. prices. Higher price variability can create buyer confusion about the true price and indicates a higher level of risk. Such variability can alter market participants' expectations concerning the durability of a price change, which, in turn, could affect the degree of response to a price change that goes against an importer's interest. (This straightforward price variability means people expect a particular price may not stick.)

As specified, Equation 4 allows us to test whether market power, buyer confusion, or both factors play a role in price and adjustment rate asymmetry. If the coefficients for either (or both) of these interaction variables are significantly different for rising prices than for falling prices, either or both variables may be a source of price asymmetry. Similar arguments hold for adjustment rate asymmetry.

The Magnitude of a Price Asymmetry and Parameter Tests

In this analysis, the magnitude of a price asymmetry represents the difference in the impact of a rising and falling price on an import unit value (IMUV). The magnitude of prices impacts is calculated using the estimates from Equation 4, which gives us this difference:

5) $$\theta_{pr} = Abs\{(\hat{\beta}_{ju} - \hat{\beta}_{jw}) + Hf(\hat{\alpha}_{ju} - \hat{\alpha}_{jw}) + SD * (\hat{\gamma}_{ju1} - \hat{\gamma}_{jwn})\}$$

where *Abs* is the absolute value and $\hat{\beta},\ \hat{\alpha},\ \hat{\gamma}$ are the estimated values of $\beta, \alpha, and\ \gamma$, repsectively. To test whether interaction terms influence the magnitude of an asymmetric IMUV response, we undertake two parameter tests. First, we test if the degree of supplier concentration (*Hf* index) in a market influences the IMUV response to Gulf prices, which implies:

$\hat{\alpha}_{ju} = \hat{\alpha}_{jw} = 0$

Rejecting this restriction implies that market concentration (or the degree of competition among suppliers) can influence the IMUV response to a change in U.S. wheat prices. We apply this test at all lags (although not explicitly shown in Equation 4). Second, we test whether the degree of concentration influences the magnitude of price asymmetry by testing the less restrictive condition:

$\hat{\alpha}_{ju} - \hat{\alpha}_{jw} = 0$

We also test this possibility using all lags. Rejection of this restriction means that the degree of market concentration (*Hf* index) can influence the magnitude of a price asymmetry. A similar test is performed on the SD variable (a proxy for price confusion and/or risk). Similar tests are used if the magnitude of adjustment rate asymmetry (θ_{pr}) is influenced by market concentration and/or price variability.

Lagged Asymmetries

Asymmetry can occur in the timing of a price response to a change in an exogenous variable (Meyer and von Cramon-Taubadel, 2004). For example, markets may immediately react to a rising price but may be slow to respond to a falling price. The specification in Equation 2 allows IMUVs to respond to U.S. wheat prices at up to *n* lags. We estimate an expanded version of this equation, which includes interaction terms (*Hf* concentration and price risk variables) at each lag.

This broader specification enables us to test coefficient differences on rising prices and falling prices, test interaction terms, and measure the magnitude of asymmetry at each lag.

Data and Estimation

In this study, the U.S. FOB Gulf price of wheat represents the export price of U.S. wheat, which strongly correlates with the prices of all classes of U.S. wheat.[3] When Gulf or Pacific prices serve as the originating prices of wheat from the U.S., the magnitude of each importing country's reaction to U.S. gulf prices can differ (Figure 1). FOB prices contain information on international supply and demand for wheat, as well as the cost of shipping to a port. IMUVs for wheat (Table 1), which represents the price we use for importing countries, contain information on additional shipping costs, shipping uncertainty, importer policies, and the competitiveness of a county's import market (Beckman et al., 2022).

Monthly Gulf wheat prices from January 2000 to January 2020, mostly representing hard red winter wheat, were obtained from the United States Department of Agriculture (USDA) Economic Research Service (2021)[4]. IMUVs[5] for wheat were assembled for seven countries, which are among the top 15 importers of U.S. wheat from the Trade Data Monitor (2021).[6] In addition, each country was chosen for analysis because data were consistently available. Data for five countries (Colombia, Japan, Mexico, Taiwan, and South Korea) were available from January 2000 to January 2020. Data for two countries (the Philippines and Indonesia) were available for shorter periods. Changes in Gulf prices provide a representative signal of changes in any class of U.S. wheat (see Footnote 3).[7]

A three-period moving standard deviation of Gulf prices was used to represent the variability of port prices. Longer period averages would have reduced movement in this measure which, as calculated, varied considerably across time. Examining all seven countries combined,

Figure 2 shows that over the years U.S. exports have been facing increased competition from other wheat exporters. Since 2012, the seven countries in our study have imported more wheat from U.S. competitors.

Prior to Estimation

We apply unit root tests (Dickey and Fuller, 1979) to determine whether there is stochastic trend in the U.S. Gulf prices and each country's IMUV. Testing monthly data over the period (January 2000–January 2020), revealed that at the 0.05 (95%) confidence level the null hypothesis of unit roots could be rejected for Gulf wheat prices and the IMUV of Taiwan but not for other countries IMUVs.[8] Repeated unit root equations were then estimated and incrementally (sample sizes) tested for different data break points. Significant differences in both the coefficient of lag price and the unit root test statistic revealed break points existed within the year 2008 for every variable. Specifically, a small break occurred in May and a large break occurred in September for the U.S. wheat price. This was the year of the financial crisis when commodity prices fell dramatically.

After finding a 2008 break, additional unit root tests were applied to two samples of data prior to and after 2008. These tests revealed that we could not reject the null hypothesis of a unit root for Gulf Wheat prices and the IMUVs of each tested country in the sub-period prior to 2008. In the second sub-period (which included Indonesia and the Philippines) we could reject the null that unit roots existed for Gulf (and Portland) prices and IMUVs of every country except for Japan.

Given the initial unit roots test, we estimated an error correction model (ECM). The ECM makes it possible to deal with non-stationary wheat data series and helps to capture the true dynamics of a system (Granger and Weiss, 1983). We estimated ECM on data for three separate time periods: 1) a sample representing all observations. 2) a sample representing observations before September 2008, and 3) a sample representing observations after 2008.

Estimating the Price Transmission Equations

Each country's model was estimated separately using Granger's two step method. First, the long-run equilibrium, Equation (1), was estimated by regressing each country's IMUV on the Gulf price. Lagged errors from the long-run equations were then used as an explanatory variable in the second-stage difference (disequilibrium), Equation (4). The long-run "equilibrium" equation was estimated for each country without lags since equilibrium relationships are stable by definition. For similar reasons, no effort was made to distinguish IMUV response to a rising and falling U.S. port price in the equilibrium equation.

For the estimation of the second-stage difference equation, three lags of the Gulf (and Portland) prices were allowed (fourth lags were consistently insignificant).[9] Further, the lagged errors of the long-run equations were included. Distinct coefficients on the rising and falling prices and the rising and falling values of the lagged long-run error term were specified, allowing us to test for asymmetry in price response at each lag and at the sum of the lags as well as asymmetric adjustment rates. The inclusion of an *Hf* index interacting with each lag U.S. wheat price allowed us to test if market concentration influenced the potentially asymmetric IMUV response. The inclusion of a moving standard deviation of price interacting with each lag of the U.S. wheat price further allowed us to test if price variability influenced the possible asymmetric deviation of the IMUV response.

Empirical Results

First Stage

Estimated coefficients on U.S. wheat prices variables ranged from a high of 1.39 in Mexico before the financial crisis of 2008 to a low of 0.72 for Indonesia. All t-statistics were significant on the wheat price variable at the 0.05 confidence level. The equations' R-square measure ranged from Mexico's 0.96 in the post-2008 sample to the Korea's 0.97 value for the full sample. However, the R-square of the Indonesia IMUV regression was only 0.17. The constant of the long-run equation representing a price markup ranged from a high 130 in Japan's equation (approximately a 60% increase over the average Gulf price) to a low of -9.92 in Mexico's equation prior to the 2008 crisis.[10] In general, constant terms were high, perhaps owing to the influence of tariff and non-tariff barriers contained within the left-hand side IMUV variable.[11]

Testing Interaction Terms in the Second Stage

We use the ECM difference equation (using ordinary least squares (OLS)) to estimate a model for five wheat-importing countries (Colombia, Japan, South Korea, Mexico, and Taiwan) across the entire time period. Each country equation was estimated three times over the full sample period, the period before October 2008, and the period after October 2008. Following a similar format, the Indonesian and Philippines equations were estimated once using monthly data from February 2010 to January 2019 and October 2006 to January 2019, respectively.

Table 2 reports a series of F statistics for the five country models that tested whether the interaction terms, (*Hf* indices and price variability) terms belong in the model. Tests were applied to sample prior to and after 2008. We also report the results of Chow tests we conducted to determine whether the model differs before and after the year 2008.[12] The test results indicate that at least one interaction variable was significant in each model in at least one of the three time periods: the whole sample period, before the 2008 crisis, and after the 2008 crisis. In Colombia, Japan, South Korea, and Indonesia (not shown), the market concentration (*Hf* index)-price interaction variable was significant at the 0.05 or 0.01 significance level. In the Japan, South Korea, Mexico, and Taiwan models, the variability-price interaction variable was significant. In Colombia and Japan, the interaction term between the *Hf* index and the adjustment rate was found to be statistically significant at the 0.01 level before the crisis. In Japan, South Korea, Mexico, and Taiwan, the interaction term between variability and adjustment was found to be statistically significant at the 0.05 confidence level or higher. Only in the single period Philippines model was neither interaction term significant.

Asymmetric Price Impacts and Adjustment Rates.

Table 3 reports F-tests for asymmetry in both the IMUV response to Gulf price changes and the adjustment rate of the ECM equations. A significant F statistic on price tests indicates that an equal IMUV response to rising and a falling U.S. port prices can be rejected. Similarly, a significant F statistic on the lagged error term tests indicates equality of adjustment rates to a positive and negative long-lagged errors terms; thus, error terms that represent the degree of displacement for equilibrium can be rejected.[13]

In imposing the restriction that IMUV response to rising and falling prices (lagged error terms) were equal, interaction terms were accounted for if previous test showed they were significant. Restrictions were imposed at the means of the *Hf* and price variability index data. Tests for asymmetry were applied at 1) each of the three price (differences) lags, 2) the sum of the price lags, and 3) each of the error correction terms. Each test was applied over the entire period sample, the pre-2008 crisis, and the post-2008 crisis. The results in Table 3 show that IMUV response to price changes and lagged long-run errors is often asymmetric. The results from an F-test that jointly tests all price impacts for asymmetry over the entire period were significant at 0.05 level for Colombia, Japan, South Korea, and Philippines; that is, the restriction that the IMUV response to rising prices and falling prices were equal could be rejected in four out the seven countries.

Testing the period before the crisis, symmetric IMUV response to price changes could be rejected for all countries except for South Korea. The post-crisis symmetric response to price changes could be rejected for Colombia and South Korea. The F statistics show a mixed result when a symmetric response to prices was tested at specific lags. For example, when testing the whole period, a symmetric price response in Colombia could be rejected at the third lag but not at the first and second lag. The same restriction of an equal response to price changes could be rejected in Japan at the second and third lag, in South Korea at the first lag, and in Mexico at the first and third lag.

Table 3 shows that, when tested over the entire period, symmetric adjustment rates — adjustment rates representing the speed which IMUV returned to their equilibrium relationship with port prices — were rejected the 0.05 significance level in all countries except for South Korea. When tested pre-2008 crisis, symmetric adjustment rates were rejected in every country except for South Korea and Mexico. When tested post-crisis, symmetric adjustment rates were rejected in both South Korea and Taiwan. One interesting feature of the Philippines model is that symmetry of price and adjustment rates were rejected even though interaction terms were not included in these tests as they were shown to be insignificant in the previous model test. In sum, IMUV responses were influenced by the direction of price changes at U.S. ports and the adjustment rates were influenced by the sign on the lagged errors of the long-run equilibrium equation.

Measuring Asymmetric Impacts

Tables 4 and 5 report estimates of each country's magnitude of asymmetry for prices and adjustment rates. This price and adjustment rate spread was calculated using the formula in Equation 5, evaluated at the means of the *Hf* and price variation (*PV)* variables. Asymmetry magnitudes are reported for before and after the 2008 financial crisis. Table 5 reports the same measure without the impact of any interaction terms.

Differences between Tables 4 and 5 show that the interaction terms had a mixed impact on asymmetry. In Table 4, Colombia's *Hf* interaction terms reduced both the spread between rising and falling prices and the adjustment rate impacts before the crisis. However, the *Hf* index raised the magnitude of the asymmetry after the crisis. In Japan, the impact of both the concentration and price variability indices slightly increased the price and adjustment rate spread before the 2008 crisis but lowered it after the crisis.

In South Korea, the sum of price variability terms increased the magnitude of asymmetry before and after the crisis. After 2008, the spread between the higher and lower adjustment rates was wider when the interaction affects were considered. In Mexico, price variability decreased the magnitude of price asymmetry but raised the spread between the adjustment rates. In Taiwan, price variability raised the price spread before the crisis and lowered it after the crisis. A post-crisis analysis of Indonesia indicated that price variability reduced the magnitude of the asymmetry of adjustment rates. This comparison of tables implies that price variability may play two roles. On the one hand, too much price variability can generate buyer confusion creating IMUV responses. On the other, too little price variability may send muted signals that could also produce buyer confusion — particularly in markets with less sophisticated technology for evaluating price differences. That is, a certain level of price variability may be required to send a sufficiently clear signal to market participants (Tomek and Gray, 1970). Indonesia and the Philippines may fit best into this particular category of markets.

Our results also showed that the magnitude of asymmetry was quite different before and after the 2008 financial crisis. If a crisis can change the magnitude of an existing asymmetry, it is possible a singular event may be capable of creating economic asymmetry when before there was none.

Testing Concentration Effects Across Different Markets.

Tests using specific country equations show that a changing market concentration widens the asymmetric reaction to rising and falling U.S. prices in some countries and narrows it in others. It is more likely that variation of concentration across markets increases asymmetry. To test for a cross-country *Hf* effect, we estimated a system of seemingly unrelated regression (SUR) equations for the five countries where data was available over the same time period (Colombia, Japan, South Korea, Mexico, and Taiwan). In other words, our system consisted of five distinct equations, each representing the disequilibrium version of the transmission of U.S. Gulf prices to a particular importing country. We imposed the cross-equation restriction that the coefficients on wheat prices were equal across markets and made the coefficients on the price and price variability interaction terms the same across all markets. However, we allowed *Hf* interaction coefficients to vary across markets and compared these estimates to those derived when the *Hf* interaction coefficients were made the same across countries. We also used the same method to measure cross-country *Hf* impacts on adjustment rates.

Table 6 reports the percent increase or decrease in the magnitude of asymmetry (price spread) arising from country differences in market concentration. These are reported at each lag, but what matters most is the magnitude of asymmetry summed impacts across all price lags. The results show a large cross-country concentration effect before and after the 2008 financial crisis. For example, the differences from the common base for Columbia, Japan, and Mexico were 325.9%, 33.7%, 19.7% higher, respectively. Taiwan's difference (the country with the highest average concentration index of 0.89) was 540.9% higher. Only in South Korea, the country with lowest average concentration index (0.33), did the degree of market concentration decrease the magnitude of asymmetry (by 88.6%). We found an even larger impact from market concentration on asymmetry after the 2008 crisis. When differences in country concentration measures were considered, Colombia's asymmetry spread increased 767.0%, Japan's increased 935.0%, Mexico's increased 82.0%, and Taiwan's increased by over 1,000.0%. After the crisis, even Korea's market concentration increased asymmetry spread by 383.0%.

Table 7 presents related results for market concentration impacts on adjustment rate asymmetry. Prior to the crisis, differences in market concentration across countries reduced the adjustment rate spread for three countries (Colombia, Japan, and Taiwan) and increased it for two countries (South Korea and Mexico). However, after the 2008 crisis, concentration differences accounted for a large part of the asymmetry spread in the systems model. For example, Japan's asymmetry increased by over 3,000%. Table 7 also reports the adjustment rates estimates of the SUR system of equations. What is notable is that, prior to 2008, three of five countries show the wrong sign adjustment rates for positive long-run errors but not for negative long-run errors, a sign of a market bubble prior to the 2008 crisis.

In general, the estimated systems of equations reveal an asymmetric price response at impact and across all time lags. Second, the magnitude of the asymmetric response varies across different time lags. There are differences not only in the size of the IMUV response across countries but also in the direction of change. Thus, both types of asymmetry discussed by Meyer and von Cramon-Taubadel (2004) exist.

Conclusion

This study examined asymmetry price response in seven markets that purchase U.S. wheat. First, the magnitude and speed of import price (represented as IMUV) response to changes in the port price of U.S. wheat exports to seven countries were tested for asymmetry. Second, tests were carried out to determine whether any of three factors influence importer asymmetry: market concentration (representing market power), price variability (representing buyer confusion and risk), and the 2008 financial crisis.[14] Third, the magnitude of importers' asymmetric price responses was measured by calculating the difference in importers' estimated IMUV response coefficients to a rising and a falling price of U.S. wheat price. Finally, using lags, differences in the timing of IMUV response to rising and falling prices of U.S. wheat exports were detected**.**

All three factors were found to influence the magnitude of asymmetry, as represented by the spread between the responses of various importing countries' IMUV to a rising and falling U.S. wheat price. Equations were estimated on data from seven countries (Colombia, Japan, Indonesia, South Korea, Mexico, the Philippines, and Taiwan) and revealed that a reduction in seller competition across time can either increase or decrease the magnitude of price asymmetry depending on the country and the period analyzed. We also found the variability of U.S. port prices of wheat impacted the magnitude of price asymmetry, the gap between an importer's IMUV response to a rising and falling exporter prices, in some countries. Estimating a system of equations, we discovered countries facing less seller competition exhibit a wider spread between the IMUV response to rising prices and falling prices. Our empirical tests also showed this spread between estimated coefficients changed after the 2008 financial crisis. Thus, an economic crisis can change the magnitude of either price or adjustment rate asymmetry.

This study leaves an opportunity for future work to expand the analysis by including other domestic prices when reliable data become available. We recognize that IMUV is imperfect data, but they are the most available data we could find given the absence of other reliable domestic import prices. At the very least, our study reveals that there are reasons for asymmetric price responses other than market power. Importantly, price variability, which can increase when there is uncertainty about trade and shipping, can lead to asymmetric transmission of prices and send distorted market signals to key importers of U.S. agricultural products.

We also note that smaller markets that are more volatile than the major ones covered in our study can be a goal for future research. For example, the Russia-Ukraine crisis will have a major impact on some markets in North Africa because Russia and Ukraine are major suppliers to those markets. Russia’s military invasion of Ukraine in the spring of 2022 has affected the global wheat market. The invasion came at a critical time when there was a spike in global demand for wheat, leading to a spike in wheat prices. Sanctions on Russian exports of commodities reduced its supply to the global market while Ukraine was unable to export for several months owing to the conflict.

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[1]Countries were selected based on availability of data and the volume of wheat trade with the U.S.

[2]The market concentration measures the level of competition among export countries in the importing market. We assume that a less concentrated market (i.e., a market with high number of competitors) is more price competitive. Dufwenberg and Gneezy (2000) find that prices depend on the number of competitors in a market. Shares are wheat import quantities from one country relative to total imports of wheat. Seasonal production patterns ensure country Shares vary across time. Importers primarily buy wheat from Canada, Argentina, Australia, the European Union (EU), Russia, Ukraine, and the U.S. (Trade Data Monitor, 2021).

[3] Among all classes of wheat, the price of Portland soft white wheat had the lowest correlation coefficient (0.927) with U.S. Gulf prices. Models were estimated and tested a second time using the Portland Soft white price. The results from the Portland models can be provided to interested readers upon request.

[4] In carrying out this study with monthly prices, we were aware that a similar study using data of higher (daily or weekly) or lower frequency (quarterly, annual) could have led to different set of results. However, trade quantities, which are used to obtain IMUVs, are only consistently available on a monthly basis. Studies that use more volatile daily price data may find that importers are quick to response to rising prices and slow to respond to falling prices, or the opposite, depending on whether a price change works for or against their benefit. In any case, our use of monthly data should be taken into account when viewing the results in the next section.

[5] Economists often use IMUVs in trade models because other forms of prices paid by importers are not available (e.g., Pattichis, 1999; Deyak et al., 1989).

[6] Because detailed breakdowns of the IMUVs from importing countries were unavailable, we initially included tariff data in our estimations but discounted their importance for two reasons: 1) tariffs were—except Korea, very small or zero and 2) would change once every 12 or 15 years. Therefore, there was insufficient variation in tariffs to have any impact on an econometric model.

[7] A statistically significant coefficient on Gulf wheat price provides an additional indication that the price is representative. Despite this, a second set of models were estimated using Portland soft white wheat.

[8] Data for Indonesia were only available from January 2010 to January 2019, and data for the Philippines were available from January 2006 to January 2019.

[9] In any given month, imports can consist of wheat that had that been contracted for export months in advance and imports made in response to current spot prices.

[10] The negative constant in Mexico equation can be interpreted as a situation where if all independent variables are set to zero the country would export rather than import. For an example, see Eisenhauer (2003) and Frost (2013).

[11]Low Durbin Watson statistics in the long-run equations implied error terms were serially correlated, thus potentially biasing standard error estimates. We did not correct for this as we did not apply hypothesis tests to the long-run equations. This allows the actual lagged error terms to be represented in the second stage ECM equations.

[12] Unit root structural break tests inform us about the underlying data, while the Chow tests for differences in price transmission before and after 2008 inform us about the model itself.

[13] Adjustment rates indicate the speed at which IMUV values return to their equilibrium relationship with port prices when displaced by a price change.

[14] The 2008 financial crisis led to a global reduction in the demand for U.S. agricultural products and a decline in agricultural prices (Shane et al., 2009).

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Tables and Figures

Table 1: **Annual average Gulf wheat price and import unit values ($US per metric ton)**

Year	Gulf Price[1]	Import Unit Values						
		Colombia	Indonesia	Japan	Korea	Mexico	Philippines	Taiwan
2000	114	150	Na [2]	164	116	117	Na	140
2001	127	159	Na	176	123	123	Na	146
2002	148	167	Na	180	136	148	Na	157
2003	146	191	Na	200	159	158	Na	180
2004	157	205	Na	222	159	169	Na	192
2005	152	208	Na	218	157	166	Na	185
2006	192	208	Na	230	181	198	187	183
2007	255	286	Na	304	243	262	229	282
2008	326	491	Na	540	334	382	383	498
2009	224	299	Na	307	225	257	300	251
2010	224	273	235	300	228	243	279	258
2011	316	392	402	419	319	329	338	343
2012	313	328	307	354	306	317	310	340
2013	312	375	430	363	312	325	329	343
2014	285	339	342	337	264	295	308	307
2015	204	312	334	297	234	250	268	287
2016	167	260	258	244	200	214	225	244
2017	177	242	244	255	198	214	233	221
2018	215	269	245	278	219	240	249	250

Notes: 1) Hard red winter wheat; 2) No data not available
Source: TDM (2022)

Table 2**: Selected F statistics: testing *Hf*[1] and *PV* interaction term**

InterVar[2]	Period	Test Variable	Colombia F-stat	Japan F-stat	Korea F-stat	Mexico F-stat	Taiwan F-stat
Dprc	wh-p[4]	3-PV	1.85*[3]	4.69***	3.66***	5.03***	2.12*
Dprc	bcrs	3-PV	1.89*	14.60***	1.33	7.31***	6.80***
Dprc	acrs	3-PV	1.22	1.62	2.25**	4.02***	0.68
Adj	wh-p	3-PV	1.60	9.78***	2.43*	4.46**	4.06**
Adj	bcrs	3-PV	1.38	20.00***	0.35	3.82**	6.08***
Adj	acrs	3-PV	2.09*	1.56	3.62**	2.37*	2.05*
Dprc	wh-p	*Hf*	7.22***	5.98***	0.76	2.10*	0.57
Dprc	bcrs	*Hf*	7.22***	11.50***	5.05***	0.92	1.08
Dprc	acrs	*Hf*	6.15***	1.90*	0.48	0.93	1.57
Adj	wh-p	*Hf*	7.31***	2.80*	0.77	1.75	1.30
Adj	bcrs	*Hf*	13.52***	4.78**	2.32	0.79	0.53
Adj	acrs	*Hf*	2.41*	0.20	0.75	1.11	0.32
Chow test at 2008[5]			f(16,219)	f(22,219)	f(22,219)	f(16,219)	f(14, 219)
			5.10***	2.96**	36.41***	3.39***	2.47**

Notes: 1) *PV* is a three-period standard deviation of gulf wheat prices and *Hf* is the Herfindal index of import concentration. **2)** InterVa is the interacted variable is the model variable with interaction of Hf and PV. Dprc is the differenced Gulf prices of wheat and Adj is the lag long-run error term. Response to this term provides the rate of adjustment to equilibrium. **3)** * p<0.10, ** p<0.05, *** p<0.01 (p-values). Variables were included in the subsequent model only if the significance was equal to or greater than .05. F-tests jointly tested all three prices, creating f(6,219) degrees of freedom. **4)** Wh-p represents whole period, bcrs means before 2008 (October) crisis, and acrs represents after crisis. **5)** A Chow test was employed to determine whether the model estimated before the 2008 financial crisis is significantly different from the model estimated after the crisis.

Table 3: **Test of asymmetric IMUV response to price differences and adjustment rates[1,2]**

Test prices[3]	Period[4]	Colombia[5]	Japan	Korea	Mexico	Taiwan	Indonesia	Philippines
all prices	all obs	10.77***[6]	4.22***	3.47***	2.76	1.43	1.5	3.12**
	bf crisis	39.7***	7.25***	0.86	1.35***	18***		
	aft crisis	2.43**	1.69	2.74**	1.22	1.6		
1 lag	all obs	2.05*	1.67	3.27**	3.06**	1.03	0.069	0.33
	bf crisis	1.46	2.72**	0.2	0.92	18.2***		
	aft crisis	2.39*	1.33	4.02**	0.39	1.8		
2 lag	all obs	2.29	4.62***	0.13	1.07	2.1	0.76	7.98***
	bf crisis	4.17**	3.2**	0.98	0.74	7.2***		
	aft crisis	3.42**	1.96	0.87	0.45	0.16		
3 lag	all obs	10.3***	5.26***	3.78	4.46**	1.3	2.5	0.21
	bf crisis	55.2***	5.12***	1.24	18.09***	23***		
	aft crisis	4.94**	2.2*	0.41	1.19	1.99		
Test adj								
adj rate	all obs	3.9**	29.99***	0.51	3.03**	5.35***	3.36	4.6**
adj rate	bf crisis	17.05***	20.00***	0.5	0.61	18.5***		
adj rate	aft crisis	0.66	1.52	6.3***	1.8	2.8**		

Notes: 1) F-tests report whether imposing the restriction that price (adjustment rate) impacts for IMUVs are the same for rising values of prices (lagged errors) and falling prices (lagged errors). Significant F statistics indicate the presence of price asymmetry. **2)** Impact tests include significant interaction variables, which were tested at the means of the data. **3)** All prices were jointly tested using 1 lag, 2 lag, and 3 lag prices. **4)** The variable "all obs" represents all observations, "bf crisis" indicates before the 2008 crisis, and "aft crisis" indicates after the 2008 crisis. **5)** The data for Indonesia and the Philippines were not long enough to break into sub-periods. **6)** The degrees of freedom of each test varies according to sample size and number of interaction terms involved in the restriction. **7)** * $p<0.10$, ** $p<0.05$, *** $p<0.01$ (p-values).

Table 4**: Spread between up and down impacts on IMUVs with interaction variables**[1]

	Colombia		Japan		Korea	
	Before[2]	**After**	**Before**	**After**	**Before**	**After**
1 lag[3]	0.54	0.56	0.63	1.37	0.27	0.84
2 lag	0.47	0.25	0.63	0.52	0.5	0.1
3 lag	1.61	0.02	1.03	0.56	0.8	0.2
SUM	0.29	0.29	2.26	2.49	0.61	1.14
adj.rate[4]	0.21	0.1	0.61	0.19	0.28	0.6
	Mexico		**Taiwan**		**Indonesia**	**Philippines**
	Before	**After**	**Before**	**After**	**After**	**After**
1 lag	0.14	0.11	0.724	1.84	.ns[6]	.ns
2 lag	0.05	0.01	0.676	0.385	.ns	.ns
3 lag	0.08	0.24	0.83	1.244	.ns	.ns
SUM	0.35	0.27	2.23	0.211	.ns	.ns
adj.rate[5]	3.19	0.61	0.78	0.098	0.02	.ns

Notes: 1) This indicates the difference in the absolute value of the impact effects of rising and falling prices and lag long-run error. **2)** This denotes before and after the 2008 financial crisis. **3)** One lag means the first lag price difference. The sum is the impact of the sum price effects. Note the sum spread is not the spread of the sum of the reported absolute value of spreads across all lags because some lag differences were negative, and some were positive. **4)** Adjust rate is the spread in the adjustment rate to equilibrium. **5)** Similar results for Portland wheat price models are available upon request. **6)** Ns means the interaction terms were not significant in the model.

Table 5: **Spread between up and down impact without interaction variables[1]**

	Colombia		Japan		Korea	
	Before	**After**	**Before**	**After**	**Before**	**After**
1 lag[2]	1.88	0.81	1.80	3.67	0.40	0.84
2 lag	0.66	2.35	0.17	0.95	0.73	0.36
3 lag	4.30	3.57	3.2	0.03	0.70	0.27
SUM	3.29	0.33	1.52	4.66	0.37	0.76
adj.rate[3]	1.37	0.02	0.57	0.27	0.19	0.09
	Mexico		**Taiwan**		**Indonesia**	**Philippines**
	Before	**After**	**Before**	**After**	**After**	**After**
1 lag[4]	0.03	0.03	0.03	2.25	0.45	0.29
2 lag	0.26	0.13	1.41	0.55	1.58	1.29
3 lag	1.14	0.19	0.69	0.78	2.69	0.23
SUM	0.61	0.83	2.00	0.92	4.71	0.23
adj.rate	1.38	0.39	0.79	0.62	1.86	1.07

Notes: 1) This is the difference in absolute value of the impact effects of rising and falling prices and lag long-run error. **2)** One lag is the first lag price difference. Sum is the impact of the sum price effects. Note the spread is not the spread of the sum of the reported absolute value of lag spreads because some differences were positive, and some were negative. **3)** Adjust rates the spread in the adjustment rate to equilibrium. **4)** Similar results for Portland wheat price models are available upon request.

Table 6: **Cross-country influence of concentration on price impact spread**

	Colombia	Japan	Korea	Mexico	Taiwan
			Before 2008 crisis		
1 lag[1]	-81.2	129.8	-56.0	-31.4	233.1
2 lag	-49.2	33.7	-91.6	19.6	-26.8
3 lag	37.1	33.7	-77.9	19.6	36.5
Sum[2]	325.9[3]	33.7	-88.6	19.6	540.9
			After 2008 crisis		
1 lag	52.4	-45.2	183.1	3.5	30.1
2 lag	2042.3	3012.9	15.7	566.3	3281.5
3 lag	1222.7	631.2	-93.5	81.1	1006.2
Sum	767.3	935.5	383	82.3	1334.9

Notes: 1) 1 lag, 2 lag, 3 lag are the first, second, and third price lag. Sum is the total effect across all lags.
2)Note the sum adds both positive and negative effects; thus do not use the sum % as each lag.
3)For example, the differences in the *Hf* index across markets and time increased the cross-section spread between rising and falling price impacts in Colombia by 325.9% before the crisis and 767.3% after crisis. The only country where the concentration lowered the spread was South Korea (before the crisis), the country that had the lowest *Hf* index of 0.33.

Table 7: **Cross-country influence of concentration on adjustment rate spread**

		Adjust rate			
		Up	**Down**	**Spread**	**%change**
Before 2008 crisis					
w/o *Hf*[1,2]		0.62	-0.38	1.06	
wth *Hf*	Colombia	-0.24	-0.13	0.11	-89.60[3]
wth *Hf*	Japan	-0.20	0.03	0.16	-77.60
wth *Hf*	Korea	0.24	-3.59	3.90	283.80
wth *Hf*	Mexico	1.47	-1.26	2.79	173.20
wth *Hf*	Taiwan	-0.56	-0.35	0.15	-79.10
After 2008 crisis					
w/o *Hf*		-0.14	-0.18	0.04	
wth *Hf*	Colombia	-0.32	-0.27	0.05	24.00
wth *Hf*	Japan	-1.93	-0.41	1.52	3802.30
wth *Hf*	Korea	-0.43	-0.74	0.30	676.90
wth *Hf*	Mexico	-0.37	-0.19	0.18	368.60
wth *Hf*	Taiwan	-0.78	-1.05	0.27	597.90

Notes: 1) Hf is the concentration index, which in the system model varies across time and countries. **2)** W/o hf represents adjustment rates without the Hf interaction effect, and with Hf represents adjustment rates with the Hf interaction effect. **3)** This represents the % change in the spread between up and down adjustment rates when Hf is included. For example, prior to the 2008 crisis, the concentration reduced the spread by 89.6% in Colombia and after the crisis increased the spread by 24%. **4)** Note that after the crisis, the concentration greatly increases the adjustment rate spread. Note also the wrong sign adjustment rates as lag errors rise, which is an indication of a market bubble.

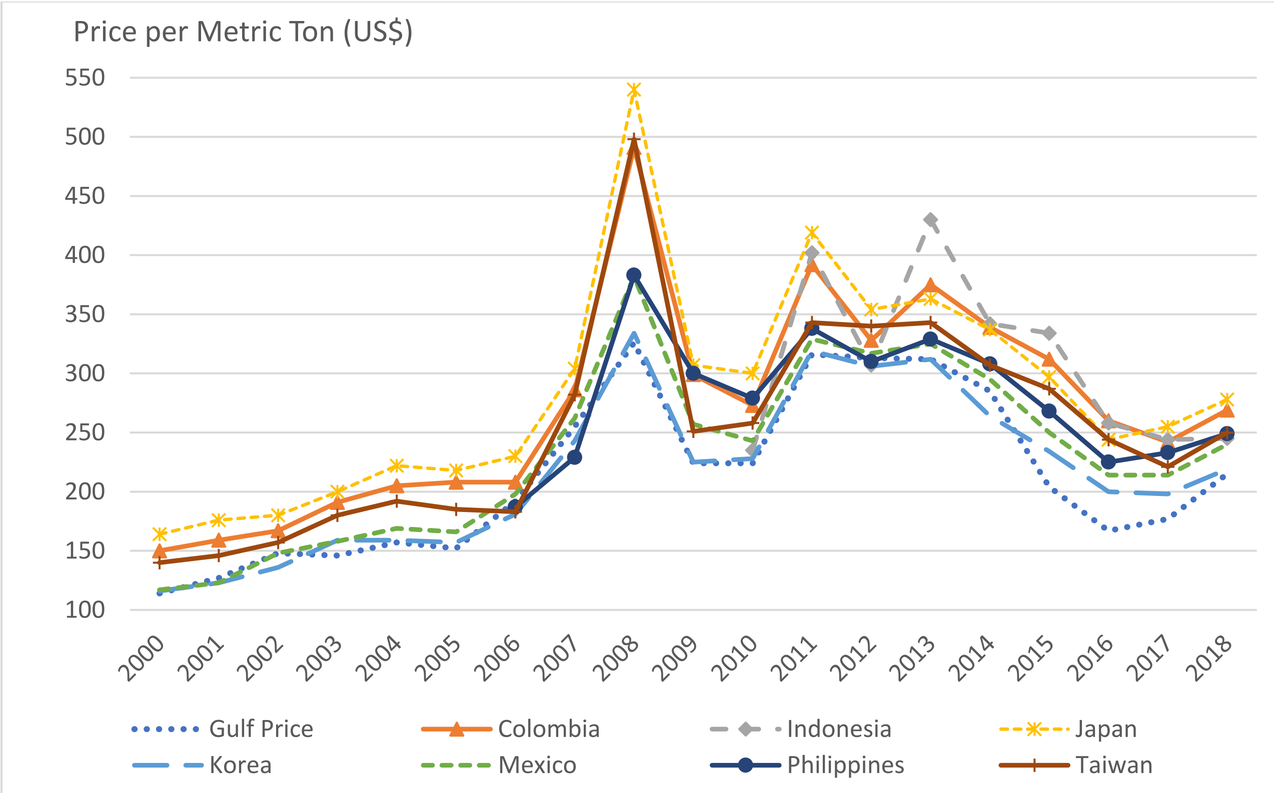

Figure 1: U.S. Gulf wheat port prices and import unit values from select countries.

Source: USDA, Economic Research Service using data from Trade Data Monitor.

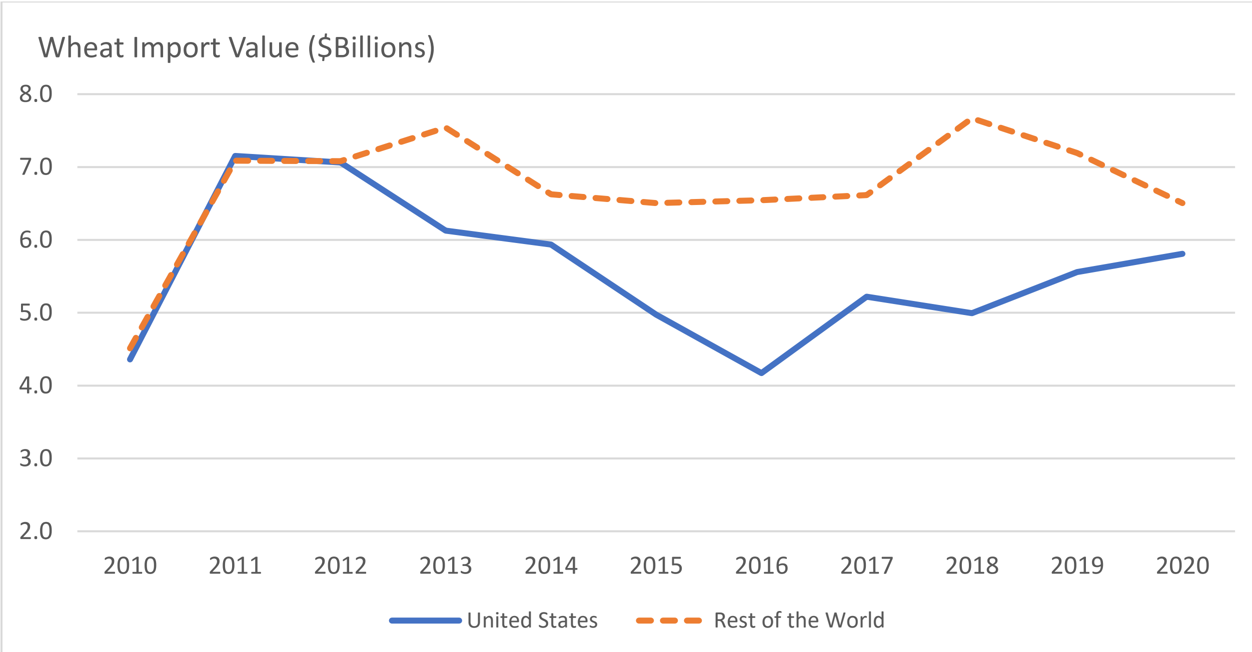

Figure 2: Total value of wheat imports from U.S and rest of the world by all selected countries in this study.

Source: USDA, Economic Research Service using data from Trade Data Monitor.