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# Agricultural Commodity Prices and Exchange Rates under Structural Change

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That exchange rates strongly influence agricultural commodity prices is a widely held belief. Observed divergences in price and exchange rate correspondence over time, however, have occasionally brought this conventional wisdom into doubt. We empirically test and find evidence to support hypotheses that key supply-use factors, such as low stocks and policy shifts, intermittently cause greater responsiveness of agricultural commodity prices to exchange rate changes because they give rise to more inelastic market demand. After accounting for these long-run effects, we also find that short-run price responsiveness to exchange rate changes is sometimes greater due to overshooting factors.

*Key words*: agricultural commodity prices, commodity booms, exchange rates, structural change, U.S. agricultural trade

## Introduction

During commodity booms a weak dollar typically signals price increases, not only for agricultural goods but also for other commodities (see figure 1 and Frankel, 2008). Ever since Schuh's (1974) classic article, the U.S. agricultural community has had a widely held belief that agricultural commodity prices are strongly influenced by exchange rates. Occasionally, deviations from this relationship occur, likely the result of specific supply-use or macroeconomic conditions. For example, the U.S. dollar steadily strengthened from May to December 2014, a factor often cited as contributing to observed commodity price declines. Agricultural commodity prices, like those of energy and metals, initially declined, but agricultural prices seemingly disconnected for part of that period. A similar phenomenon was observed during the 2005–2008 "commodity boom" and 2007–2008 "food crisis." Following a progressively weaker dollar, agricultural prices eventually rose with metals and energy prices, but with a long lag (Baffes and Haniotis, 2010). The apparent tendency for these variables to generally move in line with the stylized facts but to occasionally separate suggests that this important economic relationship warrants further analysis to explain why this is so.

Economic theory suggests full pass-through (i.e., unit elastic response) of agricultural commodity prices to exchange rate changes. Empirical measurement, however, typically finds incomplete pass-through to aggregate price levels (Goldberg and Knetter, 1997). Abbott (2010) argued that the United States is a large country (i.e., not a world price taker) in many global agricultural markets, so diminished impacts on prices and less than full pass-through should be expected. But empirical observations suggest greater-than-unit elastic responses. In the short run, these are possibly due to "overshooting" (Stamoulis and Rausser, 1988; Frankel, 2008) and, in the long run, because of collinearity between exchange rates and other macroeconomic variables. Additionally, changes in underlying market fundamentals can cause the agricultural commodity price-to-exchange rate relationship to change over time. In this paper, we identify specific factors

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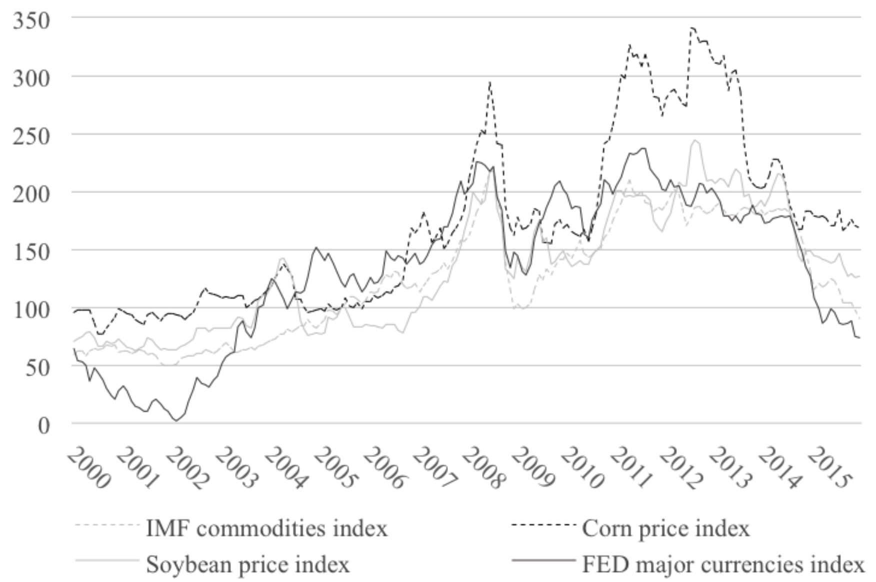


**Figure 1. Corn, Soybeans, and the IMF Commodity Price Index versus Exchange Rates, 2000–2015**
*Sources:* IMF Primary Commodity Prices and FRED.

conjectured to increase price responsiveness to exchange rate changes (i.e., cause structural change in this relationship), use economic fundamentals and logic to explain why, and test whether there is empirical evidence to support these hypotheses.

Previous attempts at empirical estimation of the agricultural commodity price and exchange rate relationship have proven difficult (Shane, Roe, and Somwaru, 2008), in part due to the apparent nonstationarity of these variables (see Frank and Garcia, 2010; Enders and Holt, 2012, for agricultural prices; and Engel and Hamilton, 1990, for exchange rates). While these series may be observed as individually nonstationary, the existence of a long-run equilibrium relationship can still be examined empirically via cointegration analysis (Engle and Granger, 1991). However, structural changes in relationship parameters may be mistakenly interpreted as nonstationarity (Perron, 1989) and lead to results seemingly contradicting cointegration.

The long-run supply-use factors that plausibly influence both market behavior and this relationship are 1) low stocks-to-use conditions, which follows from the theory of competitive commodity storage, and 2) policy-based factors. The theory of competitive storage indicates that demand for agricultural commodities has different levels and responsiveness parameters (i.e., different intercepts and slopes in a piecewise linear representation, Cafiero et al., 2011) under various stocks conditions: elastic under normal stocks-to-use ratios and inelastic for low stocks-to-use conditions (Wright, 2011).[1] Policy can also influence market behavior, including elasticity. Abbott, Hurt, and Tyner (2011) argued that policy changes in the United States—which resulted in biofuels industry expansion—and China—which adjusted soybean imports—caused demand for U.S. corn and soybeans to not only be higher (increasing price levels) but also more inelastic.

The macroeconomic factors that are proposed to cause changes in short-run price responsiveness to exchange rate changes and which drive overshooting are 1) "loose" monetary policy and 2) changes in the business cycle (e.g., recession). Dornbusch's (1976) overshooting hypothesis, adapted to commodities by Frankel (1984, 2008) and agricultural commodity prices by Stamoulis and Rausser (1988), provides an alternative explanation of parameter changes caused by monetary policy adjustments, business cycles, and inflationary expectations. In the overshooting framework,

[1] We treat the shift from one segment of the stocks demand function to another as a regime switch, hence, an apparent "structural change" in a linear cointegrating relationship (as in Wright, 2011).

agricultural goods and manufactures prices have different adjustment speeds, which causes greater than proportional short-run agricultural price responses to exchange rate changes, with long-run reversion to equilibrium.

In this study, we empirically examine corn and soybean price responsiveness to exchange rate changes from January 1990 to March 2013. We use a version of the Engle and Granger (1987) two-stage method tailored to accommodate these proposed structural changes in the cointegrating relationship parameters. The first-stage (linear levels) model includes indicator variables to account for structural changes to the means, by adjusting the intercept, and to the elasticity/slope, by interacting with the exchange rate. Supply-use structural change factors are examined in the levels models with indicator variables to represent low stocks-to-use conditions, policy-induced expansion of U.S. biofuels, and changes in Chinese soybean net import demand policy. Structural change in the short run is considered in a second-stage error correction mechanism (ECM) model, through interaction of the change in exchange rate variable with the indicator variables for macroeconomic factors: loose monetary policy and recession.

We provide theoretical results to show that greater inelasticity of market demand for an agricultural commodity (e.g., corn or soybeans) is expected to lead to higher responsiveness of prices to exchange rate changes, but only in large country cases. In the levels models, we find that agricultural commodity prices are more responsive to exchange rate changes (i.e., the parameters for this long-run relationship are nonconstant) under market conditions that cause demand to become more inelastic (e.g., low stocks-to-use, biofuels expansion, and a Chinese net import policy shift). In the second-stage ECM models, we find that overshooting factors (loose monetary policy and recession) cause increased price responsiveness in the short run. Results vary somewhat by exchange rate index used. The Federal Reserve Major Currency index had the best statistical fit of the indices or bilateral rates considered.

While recent time series studies have identified a break point in the agricultural commodity price and exchange rate relationship around the "commodity boom" of 2005–2008 (e.g., Frank and Garcia, 2010), they only identified a one-time break, which they did not attribute to underlying economic forces. We account for this one-time break in our framework, attribute it to policy changes, and explain the linkage between economic fundamentals and those policy adjustments. In addition, after we control for the policy adjustment, we find statistical evidence to support our hypotheses that both policy and low stocks-to-use conditions lead to greater price responsiveness to exchange rate changes. Had we included only one of these factors, our results indicate that low stocks-to-use better explains changes in exchange rate impacts than do the one-time breaks due to policy changes.

## Theory of the Agricultural Commodity Price and Exchange Rate Relationship

The starting point to establish a theoretical relationship between an agricultural commodity price and the exchange rate is the law of one price (LOP). Baffes (1991) defines LOP as the existence of a common price for a homogeneous commodity in spatially disparate markets. Equality of prices, expressed in terms of a common currency via exchange rate conversion, is achieved through spatial arbitrage within integrated markets, resulting in the direct effect of exchange rates on domestic prices.

Throughout this theory section, the United States is the domestic market (with prices in \$/metric ton (MT)), and the world market is the aggregate of exporters and importers of this globally traded commodity (see figure 2 for a graphical representation). In this framework, relevant exporter country excess supplies and importer country excess demands are aggregated into a composite world market net import demand, with a world price that is an average of prices in these countries. The relevant exchange rate is then a composite of exchange rates (expressed here in \$/foreign currency units) for those countries that are exporters or importers on global markets. The LOP relationship may then be

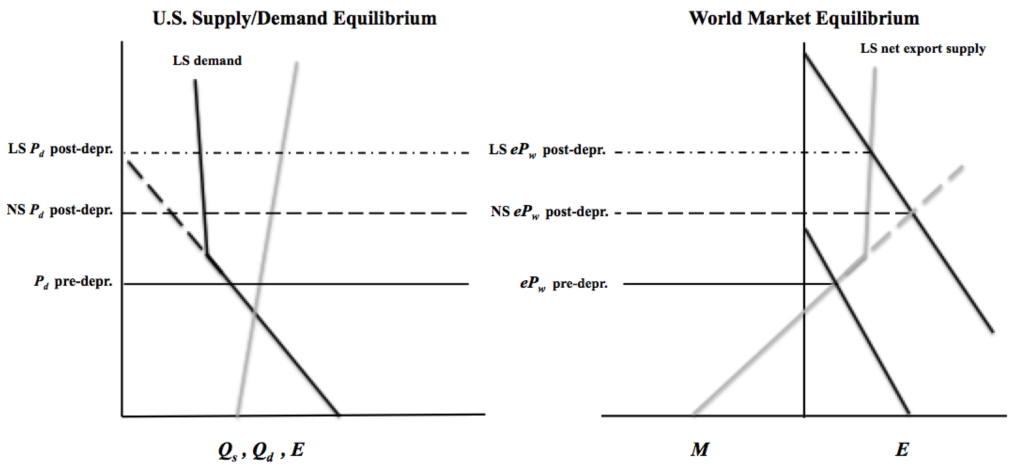


**Figure 2. Two Panel Large Country Open Economy Diagram under Depreciation**
*Notes:* Abbreviations are LS: Low stocks-to-use, NS: normal stocks-to-use, depr.: depreciation.

expressed as

(1) $$P_d = EP_w(1+\tau),$$

where $P_d$ is the U.S. equilibrium domestic price in \$/MT, $E$ is the exchange rate in \$/foreign currency units, $P_w$ is the world market equilibrium price in foreign currency units/MT, and $\tau$ are proportional transactions costs (and taxes).

Since the focus of this paper is on the responsiveness (or elasticity) of agricultural commodity prices to exchange rate changes, we provide theory relevant to the small and large country cases in order to help form expectations on the magnitudes of these elasticities.

### *Elasticity of Price Exchange Rate Changes: Small and Large Country Cases*

#### Small Country Case

Assume for now that our reference domestic country is a small country (i.e., a world price taker) exporter in global markets. In such a case, the world price adjusted for transactions costs is a constant plus white noise, resulting from variations in the world price, transactions costs, or both:

(2) $$P_w(1+\tau) = \alpha + u,$$

where $\alpha$ represents the constant portion and $u$ the random component of the (c.i.f.) world price (in foreign currency units/MT). The small country assumption means that the relevant net import demand slope that the U.S. faces is negative infinity ($\varepsilon_m = \infty$).[2] In this small country case, with all else held equal, a change in the bilateral exchange rate of the U.S. dollar to the currency of any participant country in the world market is fully passed through to the domestic U.S. price.

#### Large Country Case

We now make the contrasting (and more realistic) assumption that the United States is a large country (i.e., not a world price taker) exporter in global commodity markets. Changes in U.S. domestic

[2] In this initial small country case, the relevant net import demand curve in figure 2 is a horizontal line, which represents the world market net import demand curve and also the world price observed in the domestic economy of the small country, using the exchange rate to adjust for foreign currency units.

market behavior (e.g., a change in demand elasticity) affect the world price because the large country assumption implies that the U.S. contribution to world market net export supply is substantial. The applicable net import demand slope for a large country is not negative infinity ($\varepsilon_m < \infty$), so changes in the exchange rate are reflected in both the equilibrium domestic ($P_d$) and world prices ($P_w$).

In the linear form (as in figure 2), world market equilibrium can be represented as

$$[Q_{s0} + \varepsilon_s P_d] - [Q_{d0} - \varepsilon_d P_d] = [M_0 - \varepsilon_m P_w], \tag{3}$$

where $Q_{s0}$ is the constant and $\varepsilon_s$ is the slope in a U.S. domestic market supply function; $Q_{d0}$ is the constant and $-\varepsilon_d$ the slope in a U.S. domestic market demand function; and $M_0$ is the constant and $-\varepsilon_m$ the slope of a net import demand function for the rest of the world, hence faced by the United States.

We use equations (1) and (3) to obtain a price to exchange rate elasticity that is conditional on key market behavior parameters (notably the domestic market demand elasticity). We use this derived elasticity to show how this relationship is expected to adjust as these parameters change.

First, from elements in equation (3), we define terms $k_m = -\left(\frac{Q_{s0}-Q_{d0}-M_0}{\varepsilon_m}\right)$ and $\varepsilon = \left(\frac{\varepsilon_s+\varepsilon_d}{\varepsilon_m}\right)$. Next, we substitute these terms into equation (3) and solve for $P_w$. The redefined equation (3) is then inserted into equation (1), from which we obtain

$$P_d = E[k_m - \varepsilon P_d](1+\tau). \tag{1'}$$

We next take the total differential of the redefined equation (1') and factor. This total differential and rearranged form of equation (1) maintains the LOP equilibrium characteristics of equation (1) but allows for formation of elasticities:

$$dP_d(1 + E\varepsilon(1+\tau)) = dE[k_m - \varepsilon P_d](1+\tau). \tag{4}$$

We know that $\frac{P_d}{E} = [k_m - \varepsilon P_d](1+\tau)$ from equation (1'), so we insert it into equation (4), which yields the elasticity of an agricultural commodity price change with respect to the exchange rate:

$$\eta_{P_d,E} = \frac{dP_d}{dE}\frac{E}{P_d} = \frac{1}{1+E\varepsilon(1+\tau)} = \frac{1}{1+E\left(\frac{\varepsilon_s+\varepsilon_d}{\varepsilon_m}\right)(1+\tau)}. \tag{5}$$

$\eta_{P_d,E}$ depends on $\varepsilon_s$, $\varepsilon_d$, and $\varepsilon_m$ only in the large country case.

There are three notable results obtained from equation (5), which help to inform expectations on the sizes of responsiveness parameters in an empirical model:

1. Small country result: $\eta_{P_d,E} = 1$;
2. Large country result (#1): $\eta_{P_d,E} < 1$, with the size dependent on the magnitudes of the $\varepsilon_s$, $\varepsilon_d$, and $\varepsilon_m$ because $\varepsilon_s$, $\varepsilon_d$, $\varepsilon_m > 0$;
3. Large country result (#2): $\eta_{P_d,E}$ increases as $\varepsilon_d$ becomes more inelastic ($\varepsilon_d \to 0$).

The small country result states that a percentage change in the exchange rate will be fully passed into an equal percentage change in the domestic price (because $\varepsilon_d$ and $\varepsilon_s$ drop from this expression when $\varepsilon_m = \infty$, and so $\eta_{P_d,E} = 1$). The first large country result implies that the ability of a large country to influence the world price will result in less than full exchange rate pass-through, consistent with the imperfect competition model results of Dornbusch (1987). The third result justifies expectations that the magnitude of the commodity price to exchange rate elasticity is greater (only in the large country case) when domestic market demand ($\varepsilon_d$) becomes more inelastic.

### *Theoretical Results and Conjectured Structural Changes*

#### Long-Run Supply-Use Factors

The supply-use use factors conjectured to affect the long-run responsiveness of commodity prices to exchange rate changes are those that cause the market elasticity of demand ($\varepsilon_d$) and, therefore, net export supply to be more inelastic (i.e., decrease in absolute magnitude). The first factor that causes the market demand elasticity to become more inelastic ($\varepsilon_d \to 0$) is low stocks-to-use conditions, which follows from competitive storage theory (Wright, 2011). The persistence of the policy mandates for the Renewable Fuel Standard (for corn) and Chinese imports of soybeans also caused U.S. market demand for each crop to become more inelastic ($\varepsilon_d \to 0$) (Abbott, Hurt, and Tyner, 2011). Based on the third large country result, each of these factors is expected to cause greater responsiveness of agricultural commodity prices to exchange rate changes ($\eta_{P_d,E}$ increases).

#### Short-Run Overshooting Factors

Two macroeconomic factors, loose monetary policy and recession, are also conjectured to change the short-run agricultural commodity price to exchange rate elasticity. Adapted to agricultural commodity prices by Stamoulis and Rausser (1988), the Dornbusch (1976) overshooting model states that, because of inflationary expectations, an exchange rate change causes greater than proportional changes in agricultural prices ("flex" price goods) to make up for the relatively slower adjustment of manufactured goods ("fix" price goods) in the short run.[3] Similarly, in a sticky price business cycle model (see Ravenna and Walsh, 2011, for example), during a recession, expectations of future inflation cause relatively large "flex" price adjustments to occur because "fix" price goods are particularly resistant to adjustment (i.e., sticky). We argue that short-run exchange rate effects on agricultural commodity prices will be greater during loose monetary policy periods and/or recessions because of these changes in price adjustment during these periods.

## Empirical Models

### *Long-Run Levels Models*

Rather than propose a full structural model, we use a reduced form along the lines of equation (1) to examine whether parameter estimates are consistent with the small or large country theoretical results above. We then introduce interaction variables that account for the factors that plausibly influence the long- and short-run agricultural commodity price to exchange rate relationship.

We begin with the LOP relationship in equation (1). We next substitute our small country theoretical assumption in equation (2) into equation (1) in order to obtain a direct relationship between the domestic U.S. agricultural commodity price and the exchange rate:

$$P_d = E(\alpha + u). \tag{6}$$

We take the natural logs of the variables in in equation (6) to obtain an estimable form of the variable levels:

$$p_t = \alpha + \beta e_t + v_t, \tag{7}$$

where $p_t$ is the natural log of the U.S. agricultural commodity price, $e_t$ is the natural log of an exchange rate or exchange rate index, $\alpha$ is the intercept term, $\beta$ is the long-run elasticity of agricultural commodity price to the exchange rate, and $v_t$ is the error component in each period ($t$).

[3] Saghaian, Reed, and Marchant (2002) used the Stamoulis and Rausser (1988) overshooting framework to argue that convergence paths between agricultural prices and the exchange rate might be asymmetric.

The levels model in equation (7) is consistent with LOP. Therefore, the agricultural commodity price and exchange rate are expected to maintain a long-run cointegrating relationship (Ardeni, 1989). Based on our first two theoretical results, if the United States were a small country then $\hat{\beta} = 1$ would be expected, but if the United States were a large country then $\hat{\beta} < 1$ would be expected.

Our primary hypothesis is that plausible structural change factors can cause both the levels of prices (captured in the intercept, $\alpha$, in equation 7) and responsiveness of prices to exchange rates (captured in the elasticity term, $\beta$, in equation 7) to be nonconstant over time. In order to account for these structural changes, we introduce indicator variables constructed to proxy for those factors and include them in an expanded version of the base levels model (7). A similar adjustment was shown by Gregory and Hansen (1996) to be consistent with cointegration theory. In such a framework, indicator variables ($I_t$) adjust both the intercept ($\alpha$) and elasticity ($\beta$). The resultant parameters are $\alpha_I$, which represents the increase in levels of prices, and $\beta_I$, which represents the increase in the elasticity of prices to exchange rate changes when structural change is relevant. The resulting levels model that accounts for the long-run structural change factors is

$$p_t = \alpha + \alpha_I I_t + \beta e_t + \beta_I (e_t \times I_t) + u_t, \tag{8}$$

where the definitions of $p_t$, $e_t$, $I_t$, and the parameters are as above and $u_t$ is the error component in each period ($t$). Equation (8) can be expanded to account for both factors through inclusion of two indicator variables and interactions, accounting for each factor's simultaneous influence on the price levels and price responsiveness to exchange rates.

Our theoretical results provide expectations for the sizes of parameter magnitudes in an estimated equation (8). The expectation from the small country result is $\hat{\beta} + \hat{\beta}_I = 1$. The first large country result leads to the expectation that $\hat{\beta} + \hat{\beta}_I < 1$. If our hypotheses that the structural change factors change both the levels (to be higher) and the responsiveness of prices to exchange rates (to be greater) are true, then it is expected that $\alpha_I > 0$, and $\beta_I > 0$ (consistent with the second large country result).

### *Short-Run Error Correction Mechanism (ECM) Models*

Since the macroeconomic factors associated with the overshooting hypothesis—loose monetary policy and recession—are conjectured to change the responsiveness of agricultural commodity prices to exchange rates in the short run, a dynamic model is necessary. We therefore estimate a version of the Engle and Granger (1987) ECM model. In our short-run framework, we allow for the possibility of nonconstant responsiveness parameters in certain periods by interacting the first differenced exchange rate variable with indicator variables that represent our short-run structural change phenomena. The resultant ECM model has the form

$$\Delta p_t = \gamma_0 \hat{u}_{t-1} + \delta_0 \Delta e_t + \delta_I (I_t \times \Delta e_t) + \sum_{i=1}^{12} \phi_i \Delta p_{t-i} + \sum_{i=j}^{12} \theta_j \Delta e_{t-j} + \varepsilon_t, \tag{9}$$

where $\Delta p_t$ is the first difference of the price as defined in the levels models, $\hat{u}_{t-1}$ are the lagged residuals from a levels model in equation (8) (with either one or both indicator and interaction terms included), $\Delta e_t$ is the first difference the exchange rate or exchange rate index as defined in the levels models, $\gamma_0$ is the error-correction term, $\delta_0$ is the instantaneous response parameter that measures the short-run change in prices to a change in the exchange rate, and $\delta_I$ is the parameter that captures the greater instantaneous response during the overshooting periods.

Equation (9) is a general form that can be expanded to accommodate both overshooting factors through inclusion of an additional interaction variable and associated parameter as well as more lags of the first differenced price and exchange rate variables (and associated parameters) to account for any additional relevant dynamics. Our theoretical results suggest that, for a small country subject to LOP, the instantaneous response of prices to an exchange rate change is expected to be $\hat{\delta}_0 + \hat{\delta}_I = 1$. The first large country result leads to the expectation that the response is $\hat{\delta}_0 + \hat{\delta}_I < 1$.

If our conjectures that overshooting factors cause the response of prices to exchange rates to be greater during periods of loose monetary policy and/or recession are correct, then the interaction parameter(s) would be $\hat{\delta}_I > 0$ (consistent with the overshooting theory logic outlined above).

### *Estimation Issues*

Least squares assumptions are violated when one cannot reject nonstationarity (or fail to reject stationarity) for the random variable(s) of interest (Granger and Newbold, 1974). Engle and Granger (1987, 1991) highlighted that, in unique cases, a linear combination of nonstationary random variables is stationary and the linear combination represents a stable long-run "cointegrating" relationship that can be exploited for explanation. Our version of the Engle and Granger (1987) two-stage cointegration method is, therefore, able to accommodate nonstationary series in estimation.

Since our primary focus is on whether the model parameters are constant over time, after establishing cointegration, we test whether there is a structural break in the commodity price and exchange rate relationship. The Bai and Perron (1998, 2003) method is used to test the existence of, and to count the number of, estimated structural breaks in this relationship. After both confirming cointegration and establishing the existence of at least one structural break in the agricultural commodity price and exchange rate relationship, we move forward with the estimation of the levels and associated ECM models.[4] These cointegration estimation methods follow the Engle and Granger (1987) two-stage method but test for nonconstant long-run parameters via inclusion of indicator variables in levels models using methods similar to those used in Gregory and Hansen (1996). In the second-stage ECM models, we account for nonconstant short-run parameters, as demonstrated empirically by Beckmann, Belke, and Kühl (2015) and Arnade, Kuchler, and Calvin (2011). We expand on their methods through interaction of the response variable rather than the error correction term, which is more consistent with our theory.

## Price Data and Exchange Rate Indices

### *Prices and Analysis Period*

Our case markets in this study are U.S. corn and soybeans. Monthly corn and soybean price data are from the International Monetary Fund (IMF) Primary Commodity Prices database, for January 1990 to March 2013. January 1990 was chosen as the start date because major changes in U.S. agricultural and monetary policy just before this time are argued to have had a potentially dramatic impact on corn and soybean prices and their relation to exchange rates.

Within the 1985 U.S. farm legislation, dual provisions of reduced price supports and cutting off of government subsidized on-farm storage ended the accumulation of sizable government grain stocks and subsequently reduced U.S. official stocks (Westcott and Hoffman, 1999). Since our results depend on private stockholder behavior (following from the competitive storage theory of Wright and Williams, 1982, and Wright, 2011), we focus on the period since 1990, when stocks were privately held.

Additionally, toward the end of the 1980s, the U.S. Federal Reserve (Fed) Federal Open Market Committee (FOMC) made an adjustment to the method of monetary policy implementation. Specifically, the FOMC started to target the federal funds rate—the interest rate banks charge for overnight interbank lending (Federal Reserve Bank of New York)—rather than the borrowed reserve level and has since maintained this policy regime (Meulendyke, 1998). The manner and timing of bank and other financial institution behavioral responses to monetary policy adjustments (e.g., asset trading) likely changed around this time and impacted the exchange rate.

[4] Specific tests for stationarity, cointegration, and structural breaks and results are in the empirical results section.

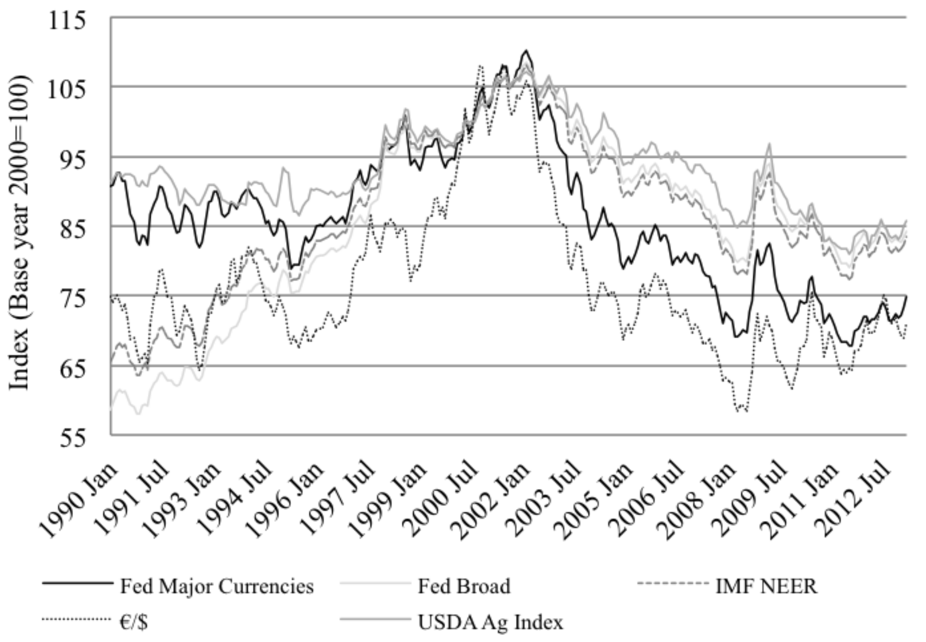


**Figure 3. Exchange Rate Indices, January 1990 to March 2013**

*Sources:* FRED, IFS, Bank of England, and USDA.

### *Exchange Rate Indices*

Using exchange rate data in empirical analysis necessarily requires a choice of bilateral rates versus exchange rate indices. Indices are preferred to bilateral rates because they better capture multilateral trade and balance of payments effects, but institutional variation in index formulation leads to differences in empirical results.

We chose to examine four exchange rate indices: 1) Fed trade-weighted Major Currencies (Fed MC), 2) Fed trade-weighted Broad (Fed Broad), 3) IMF Nominal Effective Exchange Rate (NEER), and 4) the USDA agricultural trade-weighted exchange rate index (USDA Ag Index).[5] Figure 3 shows the four indices and the bilateral Euro to USD rate (Bank of England) plotted over time with 2000 as the base year. The broad indices are less correlated than the Fed MC index with the Euro to USD bilateral rate.[6,7]

## Structural Change Indicator Variables

### *Long-Run Supply-Use Structural Change Indicator Variables*

There are two long-run supply-use structural change indicator variables for corn and soybean prices. Low stocks-to-use ratios are relevant for both corn and soybeans. One policy mandate is relevant for

[5] The Fed Major Currencies index includes bilateral rates of the U.S. dollar (USD) to the currencies of the Euro Area, Canada, Japan, the United Kingdom, Switzerland, Australia, and Sweden. The Broad Index includes bilateral rates of the U.S. dollar to currencies of twenty-six countries. These series are available from Federal Reserve Economic Data (FRED) of the Federal Reserve Bank of St. Louis. The IMF NEER index includes the bilateral rates of the USD to the currencies of twenty-six countries and is provided in International Financial Statistics (IFS). The USDA index includes bilateral rates of the USD to the currencies of seventy-nine countries and is provided by the USDA Economic Research Service (ERS).

[6] Correlation coefficients are Fed MC: €/ $ = 0.90; USDA Ag Index: €/ $ = 0.81; IMF NEER: €/ $ = 0.73; Fed Broad: €/ $ = 0.63.

[7] Loretan (2005) argued that the Fed MC index is a relatively more accurate reflection of international financial flows than the broader indices.

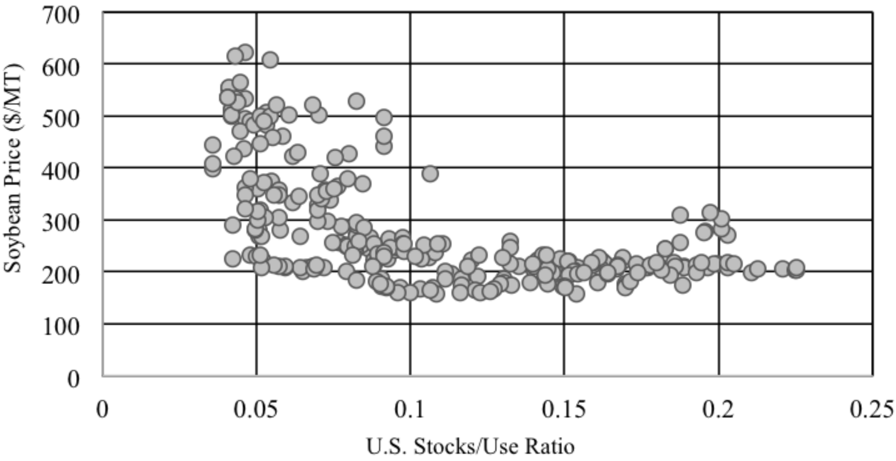


**Figure 4. Monthly Soybean Prices with Associated WASDE U.S. Stocks-to-Use Ratios**
*Sources:* USDA WASDE reports (various issues) and IMF Primary Commodity Prices.

each crop: the Renewable Fuel Standard (RFS) for corn and the adjustment to Chinese net import policy for soybeans.

## Low Stocks-to-Use Ratios

A low stocks-to-use indicator variable was created to identify periods in which stocks move near a critical (nonzero "stock-out") level as measured by the expected stocks-to-use ratio reported in USDA's monthly World Agricultural Supply and Demand Estimates (WASDE) outlook reports. In those periods the responsiveness of agricultural commodity prices to exchange rate changes is expected to be higher because market demand becomes more inelastic (Wright, 2011). A typical negative relationship, with high prices associated with low stocks and vice versa (Good and Irwin, 2015a,b), between monthly soybean prices and the associated expected stocks-to-use ratio is shown in figure 4.[8]

The hypothetical benchmark is the low stocks-to-use critical level, at which the conjectured change in price levels, price responsiveness to exchange rates, or both occurs. This ratio level was adjusted at plausible critical values to create a set of associated indicator variables. Each indicator variable received a value of 1 if stocks-to-use was below its associated critical level and 0 otherwise. Fifteen variables in the set were created with associated ratio levels ranging from 0.20 and 0.05, in increments of 0.01. Each levels model in this set, in the form of equation (8), was estimated with these indicator variables sequentially included. The indicator variable used in the empirical analysis was that with the lowest associated sum of squared errors.

The estimated critical stocks levels identifying low stocks-to-use conditions periods varied somewhat across exchange rate index. The Fed Broad, IMF NEER, and USDA Ag Indices had the same critical values of 0.10 for corn and 0.08 for soybeans. The Fed MC index had somewhat higher critical values of 0.125 for corn and 0.10 for soybeans.

## U.S. Biofuels Expansion

The massive expansion in demand for corn in biofuels production associated with the Renewable Fuel Standard (RFS) implementation is unique in two ways (Abbott, Hurt, and Tyner, 2011): 1) the large size of demand shift was unprecedented, causing higher price levels and 2) demand persistence

[8] The plot for corn is virtually identical. See Good and Irwin (2015a) for a plot of the corn price to the stocks-to-use ratio for this same period.

due to policy mandates caused the U.S. corn market demand to be more inelastic. Abbott (2013) argued that the dual facets of RFS mandates and ethanol use as a methyl tertiary butyl ether (MTBE) substitute due to state bans[9] led to the large expansion of the U.S. biofuels industry, thus causing structural change in U.S. corn market demand. The result of these policy and regulation factors was higher price levels and greater responsiveness of corn prices to exchange rate changes than in the period preceding these phenomena (Abbott, 2013). We include this RFS structural change variable in our econometric model to empirically test this conjecture.

The RFS indicator variable takes a value of 1 for the month in which the conjectured RFS-related structural change occurred and all subsequent months and a 0 for months preceding the change. Empirical search for the break month was done as before with sequential least squares estimation of a set of models in the form of equation (8). Each model in the set was assigned a RFS indicator variable associated with various plausible start months. The model with the lowest sum of squared errors was used in the reported empirical analysis. This model had an RFS break month of October 2006 across all four exchange rate indices, similar to the September 2006 time series estimate of Frank and Garcia (2010).

#### Soybean Exports to China

Abbott, Hurt, and Tyner (2011) argued that Chinese demand for U.S. soybeans accelerated rapidly after the 2007–2008 "food crisis" due to a Chinese government policy choice to rebuild soybean stocks through imports (much of which originated in the United States). The resultant import demand increase caused a dramatic rise in soybean price levels. This policy mandate also led U.S. soybean market demand to be relatively more inelastic in the period after the mandate. The size of the demand increase and persistence of demand are conjectured to have both increased price levels and made soybean prices more responsive to exchange rate changes.

To construct the soybean net import demand indicator variable, an initial indicator variable was constructed and assigned a value of 1 for a plausible start date and 0 for all preceding months. A set of indicator variables was then constructed in the same fashion for a range of plausible start dates. Each indicator variable in the set was included in sequential estimation of the econometric model in equation (8). The start date chosen for the indicator variable included in this analysis was that associated with the model that minimized the sum of squared errors. The identified critical break month varied somewhat across exchange rate indices. The break month with the lowest sum of squared errors was September 2007 for the Fed MC index, June 2007 for the Fed Broad and the IMF NEER, and February 2007 for the USDA Ag Index.

### *Short-Run Macroeconomic Structural Change Indicator Variables*

The two overshooting macroeconomic structural change indicator variables are applicable for both corn and soybean prices. Loose monetary policy and recession indicator variables interact with the first difference of the exchange rate in our corn and soybean ECM models.

#### Loose Monetary Policy

The loose monetary policy indicator variable is assigned a value of 1 during periods when the federal funds rate, as set by the Fed FOMC, was reduced or remained constant after a reduction. It receives a value of 0 for periods in which there was a rate rise or was unchanged after a previous rise. We used monthly historical values for the effective federal funds rate (from the Federal Reserve System Board of Governors), to identify these periods.

[9] See Rausser and de Gorter (2013) for a detailed discussion of ethanol substitution for MTBE after MTBE was banned in some states. Ethanol price effects of this began to be observed by July 2006.

**Table 1. Stationarity of Price and Exchange Rate Series and Associated First Differences**

| | ADF (null: nonstationary) | PP (null: nonstationary) | KPSS (null: stationary) | ADF Δ (null: nonstationary) | PP Δ (null: nonstationary) | KPSS Δ (null: stationary) |
|---|---|---|---|---|---|---|
| Corn price | −0.67 | −0.58 | 1.03*** | −7.70*** | −12.66*** | 0.17 |
| Soybean price | −0.58 | −1.11 | 1.07*** | −6.90*** | −12.14*** | 0.10 |
| Fed MC | −1.40 | −1.21 | 0.89*** | −10.82*** | −11.08*** | 0.09 |
| Fed Broad | −2.08 | −2.44 | 0.85*** | −10.33*** | −10.63*** | 0.15 |
| USDA | −1.38 | −1.34 | 0.45* | −11.36*** | −12.13*** | 0.15 |
| IMF NEER | −1.85 | −2.09 | 0.65** | −10.78*** | −10.95*** | 0.47** |

*Notes:* Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. The null hypothesis for the ADF and PP tests is that the series is nonstationary. For the KPSS test, the null hypothesis is that the series is stationary. The critical values for the ADF and PP tests are the same and are provided in Fuller (1996). The tests for levels stationarity include an intercept. The critical values for the tests of levels stationarity at the 1%, 5%, and 10% significance levels and $n = 250$ are, respectively, −3.45, −2.88, and −2.58. The stationarity tests for difference stationarity do not include an intercept. The critical values for the tests of differences stationarity at the 1%, 5%, and 10% significance levels are, respectively, −2.58, −1.95, and −1.62. The KPSS stationarity test critical values are provided in Kwiatkowski et al. (1992) and are 0.74, 0.46, and 0.35 at the 1%, 5%, and 10% significance levels.

### Recession

To capture recession periods, we referred to the National Bureau of Economic Research (NBER) U.S. Business Cycle Expansions and Contractions calendar. The created indicator variable takes a value of 1 for every month during which the U.S. economy was in recession and 0 otherwise.

These indicator variables should be similar because loose monetary policy is typically invoked to combat recession, but loose monetary policy has often persisted after a recession has ended (according to NBER). Our loose monetary policy indicator variable corresponds with the extended periods of loose monetary policy in the early 1990s, early 2000s, and since the beginning of 2007 and so captures the inertia that characterizes monetary policy (Coibion and Gorodnichenko, 2012). Recessions were less persistent over this period.[10]

## Empirical Results

### *Stationarity of Series*

The first task in implementing the cointegration framework is to determine whether the variable series included in the econometric models are stationary in levels (I(0)) and/or first differences (I(1)). To test stationarity of both levels and first differences, we implemented the Augmented Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979, 1981), the Phillips and Perron (PP) test (Phillips and Perron, 1988), and the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) tests (Kwiatkowski et al., 1992).[11] The results of these tests are listed in table 1. The results show evidence that the two price series and four exchange rate series are nonstationary in levels but stationary in first differences, meaning that all of the series appear to be integrated of order 1 (i.e., I(1)). Since these series may be nonstationary in levels, least squares assumptions would be violated unless the variables are cointegrated.

### *Levels Models without Indicator Variables, Cointegration Tests, and Structural Break Tests*

In order to test whether corn and soybean prices are cointegrated with each of the exchange rate indices, we implement two types of tests: 1) residual-based tests and 2) the Pesaran, Shin, and Smith (2001) "bounds" test.

[10] Between January 1990 and March 2013, the NBER identified thirty-four months during which the U.S. economy was in recession; our loose monetary policy indicator applies in 192 months.

[11] The null hypothesis for the ADF and PP tests is that the test variable is nonstationary. The KPSS test null hypothesis is that the test variable is stationary.

The residual-based tests involve tests of the stationarity of the estimated residuals from a levels regression model (Engle and Granger, 1987; Phillips and Ouliaris, 1990). Stationarity of the residuals is tested using the same tests as the series (ADF, PP, and KPSS), but without an intercept for the ADF and PP tests (which influences critical values). Rejection of the null hypothesis of stationarity for the ADF and PP tests, and/or failure to reject the null hypothesis of stationarity in the KPSS test implies cointegration.

The bounds testing approach accommodates the possibility of uncertainty with regard to the order of integration of the levels of variables, tests for error serial dependence, and determines dynamic stability (Pesaran, Shin, and Smith, 2001). If the critical value of the bounds test is within a set of critical bounds (determined by whether a series is I(0) or I(1)), then there is evidence of a long-run relationship between the two variables.

### Levels Models without Indicator Variables Results and Cointegration Tests

The first step in cointegration testing was to estimate equation (7), the levels model without indicator variables, for corn and soybean prices with each of the four exchange rate indices.[12,13] The results from estimation of this set of models are reported in table 2. Estimated statistics vary across the exchange rate indices. Based on the residual-based cointegration tests, only the Fed MC and the USDA Ag Index were found to be cointegrated with corn and soybean prices. The Fed MC index, however, is the only index with which corn and soybean prices have a long-run relation as determined by the Pesaran, Shin, and Smith (2001) bounds test.[14] The Fed MC index models also have substantially higher adjusted $R^2$ values than those associated with the other exchange rate indices. Failure to find cointegration with some exchange rate indices may be due to the structural changes we subsequently identify (Perron, 1989), but these tests suggest that the relationships found are consistently stronger when the Fed MC index is used.

Given these results, we chose to proceed with subsequent estimation using only the Fed MC index. With regard to the Fed MC results in table 2, the long-run response parameter of prices to exchange rate changes is statistically different from 0, which means that the exchange rate, corn, and soybean prices correspond with each other over time. It is notable that, for both the corn and soybean models, the magnitude of the long-run response parameter is also well above 1 ($\hat{\beta} > 1$) at the 1% significance level. These results are inconsistent with the small and large country assumptions presented in the theory section and imply that the exchange rate carries more information, such as expectations of global macroeconomic performance, than just direct effects of exchange rate changes on prices.

### Structural Break Test

We performed the Bai and Perron (1998, 2003) sequential tests for structural breaks in the (now established) cointegrating relationship between corn prices and the Fed MC index as well as for

[12] Some readers may be concerned with the normalization choice of the corn/soybean price as the dependent variable in equation (7). In addition to our justification that the normalization follows directly from LOP theory, we also conducted Granger causality tests, which Maddala and Kim (1998) argued are useful for normalization choice in a cointegration framework. The results for the Granger causality F-test statistics with two lags (determined by lowest Akaike Information Criterion (AIC)) were 4.38 and 4.16 for corn and soybeans with the exchange rate (Fed MC index) as the independent variable. These were statistically significant at the 5% significance level. With the exchange rate (Fed MC index) as the dependent variable and corn and soybean prices as the independent variable, the F-statistic was 0.84 for corn and 0.80 for soybeans. Neither was statistically significant.

[13] Some readers may also be concerned with endogeneity. We defer to Chen and Rogoff (2003), who argued that endogeneity is not an issue in a cointegration framework.

[14] Associated with the Pesaran, Shin, and Smith (2001) bounds test are preliminary tests for error serial correlation and a check for the dynamic stability of AR roots. The p-value for the Breusch-Godfrey Serial Correlation LM Test was 0.35, so we failed to reject the null hypothesis of no error serial correlation. The value for the inverted AR roots was 0.31, which is far from 1, and so there appears to be no unit root.

**Table 2. Levels Models Results for Models with No Indicator Variables**

| | Corn | | | |
|---|---|---|---|---|
| | **Fed MC** | **Fed Broad** | **USDA Ag Index** | **IMF NEER** |
| $\hat{\alpha}$ | 16.16*** | 6.32*** | 23.14*** | 8.54*** |
| | (31.77) | (9.74) | (22.67) | (11.22) |
| $\hat{\beta}$ | 2.53*** | 0.32** | 3.99*** | 0.81*** |
| | (22.22) | (2.25) | (17.91) | (4.84) |
| ADF $\hat{v}$ | −2.26** | −0.68 | −1.78* | −0.70 |
| PP $\hat{v}$ | −2.58*** | −0.65 | −2.14** | −0.92 |
| KPSS $\hat{v}$ | 0.36* | 1.19*** | 1.39*** | 1.32*** |
| Pesaran, Shin, and Smith (2001) bounds test | 4.84* | 0.65 | 2.21 | 0.56 |
| Adjusted $R^2$ | 0.64 | 0.01 | 0.53 | 0.07 |
| | **Soybeans** | | | |
| | **Fed MC** | **Fed Broad** | **USDA Ag Index** | **IMF NEER** |
| $\hat{\alpha}$ | 16.18*** | 6.74*** | 22.22*** | 8.75*** |
| | (36.19) | (11.28) | (23.47) | (12.47) |
| $\hat{\beta}$ | 2.38*** | 0.25** | 3.64*** | 0.70*** |
| | (23.76) | (1.97) | (17.60) | (4.54) |
| ADF $\hat{v}$ | −2.58*** | −0.65 | −1.85* | −0.77 |
| PP $\hat{v}$ | −2.90*** | −1.16 | −2.05** | −1.15 |
| KPSS $\hat{v}$ | 0.40* | 1.21*** | 1.37*** | 1.34*** |
| Pesaran, Shin, and Smith (2001) bounds test | 4.91* | 0.94 | 2.25 | 1.00 |
| Adjusted $R^2$ | 0.67 | 0.01 | 0.53 | 0.07 |

*Notes:* $\hat{\alpha}$ is the estimated intercept term; $\hat{\beta}$ is the estimated elasticity of the corn/soybean price to the exchange rate. Values in parentheses are t-statistics. Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. The null hypothesis for the ADF and PP tests is that the residuals are nonstationary. For the KPSS test, the null hypothesis is that the residuals are stationary. The critical values for the ADF and PP tests are the same, and are provided in Fuller (1996). The tests for residual stationarity do not include an intercept. The critical values for the residual stationarity tests at the 1%, 5%, and 10% significance levels and $n = 250$ are, respectively, −2.58, −1.95, and −1.62. The KPSS stationarity test critical values are provided in Kwiatkowski et al. (1992) and are 0.74, 0.46, and 0.35 at the 1%, 5%, and 10% significance levels. The critical value bounds for the Pesaran, Shin, and Smith (2001) "bounds test" for [I(0),I(1)] and are [6.84, 7.84], [4.94, 5.73], and [4.04, 4.78] at the 1%, 5%, and 10% significance levels.

soybeans. Clear evidence was found for the existence of breaks in these relationships.[15] We believe that these breaks can be explained by economic factors, captured by our created indicator variables.

### *Levels Models with Indicator Variables*

The empirical results for estimation of the corn and soybean sets of long-run levels models that account for the conjectured structural change in the agricultural commodity price and exchange rate relationship via the supply-use indicator variables, in the form of equation (8), are presented in table 3. Our hypotheses, tested through estimation of equation (8), are that both the long-run levels of prices will be higher ($\alpha_I > 0$) and the long-run responsiveness of prices will be greater ($\beta_I > 0$) when the conjectured supply-use factors are relevant. Each set includes models with the low stocks-to-use indicator variable tailored for each crop, the RFS and Chinese net import change policy variables for corn and soybeans, respectively, as well as a composite model with both the respective low stocks-to-use and policy indicator variables.

It is clear in the corn model estimation results that our conjectured supply-use variables are useful to explain why the corn price and exchange rate relationship changes over time. While the

[15] Eight and sixteen structural breaks were sequentially identified for corn and soybeans, respectively. Notably, the first break dates identified for both corn and soybeans were those found for the RFS and Chinese imports dates searches, October 2006 and September 2007, respectively, using our alternative least squares method.

**Table 3. Levels Models Results for Models with Indicator Variables**

| | Corn | | |
|---|---|---|---|
| | Low Stocks-to-Use | RFS | Low Stocks-to-Use & RFS |
| $\hat{\alpha}$ | 10.64*** | 8.42*** | 7.59*** |
| | (20.94) | (13.28) | (15.21) |
| $\hat{\beta}$ | 1.33*** | 0.83*** | 0.66*** |
| | (11.75) | (5.92) | (5.97) |
| $\hat{\alpha}_{s/u}$ | 8.24*** | – | 3.59*** |
| | (10.97) | | (4.72) |
| $\hat{\beta}_{s/u}$ | 1.79*** | – | 0.75*** |
| | (10.60) | | (4.43) |
| $\hat{\alpha}_{rfs}$ | – | 13.09*** | 8.68*** |
| | | (7.69) | (5.86) |
| $\hat{\beta}_{rfs}$ | – | 2.91*** | 1.92*** |
| | | (7.43) | (5.66) |
| ADF $\hat{u}$ | −2.16** | −3.31*** | −3.57*** |
| PP $\hat{u}$ | −4.96*** | −2.92*** | −4.88*** |
| KPSS $\hat{u}$ | 0.23 | 0.12 | 0.19 |
| Adjusted $R^2$ | 0.83 | 0.80 | 0.88 |

| | Soybeans | | |
|---|---|---|---|
| | Low Stocks-to-Use | Low Stocks-to-Use & Chinese Net Import Policy | Chinese Net Import Policy |
| $\hat{\alpha}$ | 10.79*** | 10.17*** | 10.60*** |
| | (18.88) | (19.35) | (19.68) |
| $\hat{\beta}$ | 1.21*** | 1.06*** | 1.17*** |
| | (9.54) | (9.09) | (9.77) |
| $\hat{\alpha}_{s/u}$ | 5.60*** | – | 2.00** |
| | (8.38) | | (2.32) |
| $\hat{\beta}_{s/u}$ | 1.19*** | – | 0.40** |
| | (7.95) | | (2.09) |
| $\hat{\alpha}_{imp}$ | – | 6.21*** | 3.71** |
| | | (3.53) | (2.42) |
| $\hat{\beta}_{imp}$ | – | 1.33*** | 0.80** |
| | | (3.27) | (2.28) |
| ADF $\hat{u}$ | −6.44*** | −4.56*** | −5.58*** |
| PP $\hat{u}$ | −6.44*** | −3.81*** | −5.71*** |
| KPSS $\hat{u}$ | 0.16 | 0.06 | 0.07 |
| Adjusted $R^2$ | 0.86 | 0.82 | 0.88 |

*Notes:* $\hat{\alpha}$ is the estimated intercept term; $\hat{\beta}$ is the estimated elasticity of the corn/soybean price to the exchange rate; $\hat{\alpha}_{s/u}$ is the estimated intercept in low stocks-to-use periods (both corn and soybeans); $\hat{\beta}_{s/u}$ is the estimated elasticity of the corn/soybean price to the exchange rate in low stocks to use periods; $\hat{\alpha}_{rfs}$ is the estimated intercept in the RFS period; $\hat{\beta}_{rfs}$ is the estimated elasticitiy of the corn price to the exchange rate in the RFS period; $\hat{\alpha}_{imp}$ is the estimated intercept in the Chinese soybean imports policy adjustment period; $\hat{\beta}_{imp}$ is the estimated elasticity of the soybean price to exchange rate in the Chinese soybean imports policy adjustment period. Values in parentheses are t-statistics. Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. The null hypothesis for the ADF and PP tests is that the residuals are nonstationary. For the KPSS test, the null hypothesis is that the residuals are stationary. The critical values for the ADF and PP tests are the same, and are provided in Fuller (1996). The tests for residual stationarity do not include an intercept. The critical values for the residual stationarity tests at the 1%, 5%, and 10% significance levels and $n = 250$ are, respectively, −2.58, −1.95, and −1.62. The KPSS stationarity test critical values are provided in Kwiatkowski et al. (1992) and are 0.74, 0.46, and 0.35 at the 1%, 5%, and 10% significance levels.

direct price to exchange rate elasticity remains statistically significantly greater than 0 at the 1% significance level ($\hat{\beta} > 0$) in all models, the responsiveness parameter during both low stocks-to-use and RFS periods is also statistically significantly greater than 0 at the 1% significance level

($\hat{\beta}_{s/u} > 0, \hat{\beta}_{rfs} > 0$). This is clear empirical evidence in support of our conjecture that responsiveness of corn prices to exchange rate changes is higher during low stocks-to-use periods and since the RFS and associated regulations were implemented. Additionally, the estimated intercept parameters associated with the low stocks-to-use and RFS indicator variables were also statistically greater than 0 at the 1% significance level ($\hat{\alpha}_{s/u} > 0, \hat{\alpha}_{rfs} > 0$). This implies that prices are higher during low stocks-to-use periods and the RFS period, consistent with the results of other authors (e.g., Good and Irwin, 2015a; Abbott, Hurt, and Tyner, 2011). The cointegration tests remain robust after the addition of the indicator variables. Adjusted $R^2$ values for all models are substantially higher than for those without an indicator variable. The adjusted $R^2$ for the model with only the low stocks-to-use indicator variable is somewhat higher than that with only the RFS indicator, but the model with both indicator variables performs best.

The finding that both responsiveness parameters associated with low stocks-to-use and RFS are statistically significantly different from 0 at the 1% significance level is a key result of this paper. This result implies that time series models that only account for a one-time break around 2005–2008 (e.g, Frank and Garcia, 2010) due to policy do not fully capture important economic fundamentals that explain why the corn price and exchange rate correspondence changes intermittently.

Our results for the corn set broadly apply for the soybean set. The soybean model results further support our conjectures that supply-use factors can help explain why soybean prices more closely correspond with exchange rates during some periods than others. The magnitudes of the direct response parameters of soybean prices to exchange rates ($\hat{\beta}$), still statistically significantly different than 0 at the 1% significance level for all soybean models, are broadly higher for soybeans than for corn. Low stocks-to-use periods have higher soybean price responsiveness to exchange rate changes, with $\hat{\beta}_{s/u} > 0$ at the 5% significance level for all models. There is clear evidence that soybean price responsiveness to exchange rates was higher during the period since the Chinese adjustment to their net import policy, with $\hat{\beta}_{imp} > 0$ at the 5% significance level for all models. However, the low stocks-to-use periods and the Chinese net import policy adjustment appear to be relatively more correlated with each other than are the low stocks-to-use and RFS periods for corn. This is shown in the decline in the associated t-statistics for the intercept and response interaction term parameters in the model that included both the low stocks-to-use and Chinese import policy change indicator variables. The adjusted $R^2$ of the model that only included the low stocks-to-use indicator variable was higher than that with only the Chinese net import policy change indicator variable, but the model with both had the highest adjusted $R^2$ value. We also found strong statistical support for the conjecture that low stocks-to-use periods are associated with greater soybean responsiveness to exchange rate changes while simultaneously accounting for the one-time Chinese import policy shift.

Across all corn levels models, the magnitudes of the combined responsiveness parameters were statistically significantly greater than 1 ($\hat{\beta} + \hat{\beta}_I > 1$) at the 1% significance level; the same was found for soybeans at the 5% significance level. That the combined exchange rate effects are statistically significantly above 1 at high confidence levels is further evidence that the exchange rate captures more information than its direct price effect (e.g., expectations for global macroeconomic growth).

### *Short-Run ECM Models*

The ECM models were tailored to account for the two hypothetical macroeconomic overshooting factors—loose monetary policy and recession—and take the form of equation (9).[16] The results from the estimation of this set of ECM models are reported in table 4. Since the long-run levels models with both of the supply-use factor indicator variables in them had the best statistical fit, the residuals from these models were included in the sets of second-stage ECM models. If our hypothesis that higher corn and soybean prices are more responsive to exchange rate changes during

[16] Each ECM model was estimated with one to twelve lags of the first-differenced variables. For all ECM models, the model with one lag always had the lowest AIC. This AIC was also lower than the less general model with no lags. Thus, only results from models with one lag are reported.

**Table 4. ECM Models Results**

| | Corn | | |
|---|---|---|---|
| | **Interaction Fed Funds Only** | **Interaction Recession Only** | **Interaction Both Fed Funds and Recession** |
| $\hat{\gamma}_0$ | $-0.12^{***}$ | $-0.13^{***}$ | $-0.13^{***}$ |
| | (−4.58) | (4.82) | (−4.87) |
| $\hat{\delta}_0$ | −0.56 | 0.02 | −0.57 |
| | (−1.35) | (0.09) | (−1.37) |
| $\hat{\delta}_{ff}$ | $1.16^{**}$ | – | $0.81^{*}$ |
| | (2.48) | | (1.69) |
| $\hat{\delta}_{rec}$ | – | $1.53^{***}$ | $1.32^{***}$ |
| | | (3.31) | (2.75) |
| $\hat{\phi}_1$ | $0.30^{***}$ | $0.30^{***}$ | $0.29^{***}$ |
| | (5.31) | (5.29) | (5.21) |
| $\hat{\theta}_1$ | −0.26 | −0.27 | −0.28 |
| | (−1.27) | (−1.32) | (−1.38) |
| AIC | −2.99 | −3.00 | −3.01 |
| (AIC range: 1 to 12 lags) | (−2.99 to −2.89) | (−3.00 to −2.91) | (−3.01 to −2.92) |
| Adjusted $R^2$ | 0.15 | 0.16 | 0.17 |

| | Soybeans | | |
|---|---|---|---|
| | **Interaction Fed Funds Only** | **Interaction Recession Only** | **Interaction Both Fed Funds and Recession** |
| $\hat{\gamma}_0$ | $-0.11^{***}$ | $-0.10^{***}$ | $-0.11^{***}$ |
| | (−3.96) | (−3.64) | (−3.88) |
| $\hat{\delta}_0$ | −0.27 | $0.47^{**}$ | −0.28 |
| | (−0.66) | (2.13) | (−0.70) |
| $\hat{\delta}_{ff}$ | $1.29^{***}$ | – | $1.04^{**}$ |
| | (2.79) | | (2.19) |
| $\hat{\delta}_{rec}$ | – | $1.25^{***}$ | $0.99^{**}$ |
| | | (2.72) | (2.10) |
| $\hat{\phi}_1$ | $0.33^{***}$ | $0.32^{***}$ | $0.31^{***}$ |
| | (5.62) | (5.42) | (5.23) |
| $\hat{\theta}_1$ | $-0.48^{**}$ | $-0.46^{**}$ | $-0.47^{**}$ |
| | (−2.34) | (−2.25) | (−2.33) |
| AIC | −3.05 | −3.05 | −3.06 |
| (AIC range: 1 to 12 lags) | (−3.05 to −2.95) | (−3.05 to −2.95) | (−3.06 to −2.96) |
| Adjusted $R^2$ | 0.17 | 0.17 | 0.18 |

*Notes:* $\hat{\gamma}_0$ is the estimated error correction parameter associated with the levels model with both indicator variables included; $\hat{\delta}_0$ is the estimated short-run price to exchange rate responsiveness parameter; $\hat{\delta}_{ff}$ is the estimated short-run responsiveness parameter for loose monetary policy periods; $\hat{\delta}_{rec}$ is the estimated short-run responsiveness parameter for recession periods; $\hat{\phi}$ is the estimated parameter associated with the lagged first difference of the price; $\hat{\theta}$ is the estimated parameter associated with the lagged first difference of the exchange rate. Values in parentheses are t-statistics. Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level.

loose monetary policy periods, recession periods, or both is correct, then ($\hat{\delta}_I$) for the respective interaction variable will be statistically significantly greater than 0 ($\hat{\delta}_I > 0$). Similar to the estimation of the levels models, the ECM models were estimated sequentially, with one of the respective overshooting factors included in individual models and then both included in a composite model.

For the corn set we found evidence that corn prices have a higher response to exchange rate changes during both loose monetary policy and recession periods. The short-run response parameter associated with the loose monetary policy interaction variable is statistically significantly different from 0 ($\hat{\delta}_{ff} > 0$) at the 10% significance level for both models and at the 5% significance level in the model in which it was the lone interaction variable. There was somewhat stronger evidence of higher short-run corn price responsiveness during recession periods. The short-run corn price responsiveness to exchange rates parameter associated with the recession interaction variable was statistically significantly different from 0 ($\hat{\delta}_{rec} > 0$) for all models at the 1% significance level. The model with both interaction variables had the lowest associated Akaike Information Criterion (AIC) value.

Like the results for the corn set, results for soybeans showed clear evidence that soybean price short-run responsiveness to exchange rate changes is higher during loose monetary policy and recession periods. The short-run soybean price responsiveness parameter to exchange rate changes during loose monetary policy periods is statistically significantly different than 0 ($\hat{\delta}_{ff} > 0$) for all models at the 5% significance level. The same is true for the parameter associated with the recession interaction variable ($\hat{\delta}_{rec} > 0$) at the 5% significance level for all models. The soybean models, therefore, show more balanced evidence than the corn models that both loose monetary policy and recession are important to explain greater soybean price responsiveness to exchange rate changes (in contrast to more robust results for recession for corn). The model with the lowest AIC value was again the model with both interaction variables, which suggests (as with corn) that both the loose monetary policy and recession interaction variables help explain short-run dynamics.

## Conclusions

This paper presented clear econometric evidence that the long-run and short-run responsiveness of agricultural commodity prices to exchange rates not only varies over time (i.e., the agricultural commodity prices and exchange rate relationship undergoes structural change) but also that the changes in responsiveness can be explained with economic theory.

We conjectured that that supply-use factors (low stocks-to-use for corn and soybeans, RFS policy and associated regulation implementation for corn, and Chinese net import policy change for soybeans) explain observed changes in long-run correspondence between prices and exchange rates. These hypotheses were formed based on theory and empirical observations indicating that market demand would be more inelastic when such factors are relevant. We outlined theoretical results to show the linkage between changes in market demand elasticities and agricultural commodity price to exchange rate responsiveness. The theory of overshooting was called upon to explain why the short-run responsiveness of agricultural commodity prices to exchange rates is likely higher during periods of loose monetary policy and recession.

Results from the long-run levels models showed that both corn and soybean prices are more responsive to exchange rate changes under low stocks conditions. For corn, higher responsiveness was also found for the period in which the RFS and associated regulatory measures became relevant. For soybeans, evidence of higher responsiveness was especially strong for the period in which China adjusted its soybean import policy. Results for the short-run ECM models showed evidence that overshooting mechanisms can help explain higher short-run response of corn and soybean prices to exchange rates.

Numerous time series studies (e.g., Frank and Garcia, 2010) have identified structural breaks in the agricultural price relationships (with exchange rates, oil prices, and other variables) around the time of the 2005–2008 “commodity boom” and the 2007–2008 “food crisis.” Our analysis showed that economic fundamentals are more powerful in explaining when and why these breaks occur. Since these economic fundamentals are intermittently relevant, structural change is more likely to recur regularly rather than be a one-time phenomenon.

Empirical results varied dramatically based on the exchange rate index (or bilateral rate) used. We found that both corn and soybean prices are strongly correlated and cointegrated with the Federal Reserve Major Currencies (Fed MC) index. Our long-run levels models results showed consistently estimated price to exchange rate response elasticities statistically significantly greater than 1 for both corn and soybeans, even after accounting for our conjectured higher responsiveness factors. These estimates vary greatly from theory (large country theory suggests a response elasticity of less than 1) and imply that the exchange rate carries more information than simply its direct effect on prices, trade, or capital flows. We believe a key correlated factor is expectations on global macroeconomic performance. We (and others) have found it difficult to identify a variable that could capture these expectations, especially at monthly frequency. Indeed, the exchange rate itself is often used as a proxy to measure global macroeconomic performance.

Our results will be of interest to stakeholders who follow and wish to explain corn and soybean price movements over time. Accurate forecasts on agricultural commodity price responses to exchange rate changes will rely on the ability to account for expectations on relevant supply-use fundamentals, policy, and macroeconomic conditions, which can change quickly.

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