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Atanu Ghoshray

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Abstract

This paper examines the extent to which increases in international food prices during the past few years have been transmitted to domestic prices in selected Asian developing countries. In analyzing the historical data, evidence on price transmission for important food commodities such as rice, wheat, and edible oil have been considered. The price transmission elasticity has been estimated using regression models coupled with recent econometric techniques such as unit root tests and error correction models with threshold adjustment. Finally, the paper draws some policy implications from the empirical results. This study provides the numerical estimates on the empirical relationship between international prices and domestic prices. The analysis uses commodity-specific monthly data rather than annual data during a period of substantial policy reforms in order to understand both long-run and short-run relationships between world and domestic prices.

I. Introduction

International agricultural commodity prices are set to rise considerably over the near future amid growing demand from emerging economies such as the People's Republic of China (PRC) and India and for biofuel production, according to an annual joint report by the Organisation for Economic Co-operation and Development (OECD) and the Food and Agriculture Organization (FAO). While agricultural commodity prices have fallen from their record peaks in 2008, it is believed that they are set to pick up again and are unlikely to drop back to their average levels of in the past decade.

In recent years economists have cited various reasons for this increase in agricultural commodity prices. Higher food demand in the emerging economies have led to reductions in food exports from these countries; and rising oil prices have led to increased demand for biofuel raw materials such as wheat, soybean, maize, and palm oil, which in turn have reduced the use of such crops for food production and animal feed. However, the implication of the impact of higher food prices on households, especially the poor in these countries, makes it necessary for policy makers to know whether and to what extent international commodity prices are transmitted to domestic prices and its impact on the economy.

Besides international prices, other factors such as supply shocks due to adverse weather conditions, political uncertainty, and an increase in domestic demand that may result from economic growth and/or flow of remittances can contribute to an increase in domestic prices in developing countries. The immediate adverse effect of higher domestic food prices is a fall in the real income of the households in real terms with a larger proportion of total expenditure being spent on food by the poor. In view of the importance of the issue, it is necessary for policy makers to design appropriate domestic economic and trade policies that can mitigate the adverse effects of international price changes on domestic prices based on credible knowledge of the "price transmission elasticity" between imported and domestically produced goods.

In theory if two markets are linked by trade in a free market regime, excess demand or supply shocks in one market will have an equal impact on price in both markets. However, prohibitively high import tariffs can obliterate opportunities for spatial arbitrage, while tariff rate quotas may result in international price changes not being proportionately transmitted to domestic prices at all points in time. The implementation of price support policies such as intervention mechanisms and floor prices may result in international and domestic prices being either unrelated to each other, or related in a nonlinear manner

depending upon the level of intervention or price floor relative to the international price. Changes in the international price will have no effect on the domestic price level when the international price lies on a level lower than that at which the price floor has been set. If the change in the international price is above the price floor level, it can be transmitted to the domestic market. The upshot is that price floor policies may result in the domestic price being completely unrelated to the international market below a certain threshold; or in the two prices being related in a nonlinear manner so that increases in international price are transmitted to the domestic level while decreases in international price are transmitted in a relatively slow or sluggish manner. Besides policies, domestic markets can be insulated by large marketing margins that may exist due to high transfer costs. This is particularly acute in developing countries where these large marketing margins may appear due to poor infrastructure, transportation, and communication services. These problems hinder the transmission of price signals as they may prohibit arbitrage. As a result, changes in international prices are not fully transmitted to domestic prices, resulting in countries adjusting partly to shifts in world demand and supply.

Nonlinearities and asymmetric adjustment remain an important issue to be addressed. This is true especially when the aim of the model is to take into account the above issues that call for a threshold mechanism, which causes a different adjustment to transmission of price signals. Such a discrete mechanism implies that movements toward long-run equilibrium do not take place at all points in time but only when the divergence from equilibrium exceeds the threshold. Rapsomanikis, Hallam, and Conforti (2006) urge for future research in international price transmission to be conducted in this area of two regime cointegration with threshold adjustment, which will be adopted in this paper.

Why is international price transmission important? Firstly, the price transmission mechanism can shed light on the stability of international prices. If a decrease in international prices is not fully transmitted to domestic prices, then reductions in world supply and increases in world demand that would have otherwise occurred will not take place. This would make the price decrease more acute and prolonged. Secondly, negotiations in agriculture since the 1986 Uruguay Round of the General Agreement on Tariffs and Trade have followed the objective of tariffication, that is, to convert all prevailing trade intervention instruments into trade taxes or subsidies. It is argued that this objective can lead to complete elimination of quantitative restrictions on trade. With a constant import tariff (tax or subsidy in the case of exports) international price signals would be transmitted to domestic prices. This study can help policy makers judge how far agricultural markets are from the tariffication objective.

The aim of this study is twofold. The first aim is to study the underlying trends for selected international agricultural commodity prices. Analyzing the trends in agricultural commodity prices is important as many developing countries rely on a small number of commodities to generate the majority of their export earnings. If commodity prices are found to contain declining trends over sustained periods of time, countries that are heavily

reliant on their commodity exports would have to export more to maintain their current level of imports. If their export revenues fall in relation to their import receipts over time, this would lead to a widening deficit of the current account. If increased borrowing is allowed to compensate for the deficit, a current and capital account deficit may appear, triggering a decline in international monetary reserves and currency instability.

The second aim of the study is to test for international price transmission employing the method of cointegration in the presence of asymmetric error correction. An attractive method of estimation and testing follows the approach by Enders and Siklos (2001). Noting that there is a correspondence between error correction models (ECM) representing cointegrating relationships and autoregressive models of an error term, the method is an application of the Enders and Siklos (2001) threshold autoregressive (TAR) and momentum-threshold autoregressive (M-TAR) method of adjustment. Under the TAR model, the underlying price series might exhibit asymmetry where one price remains above the other price, but for short intervals, the lower price peaks above the higher price (Ghoshray 2002). For the M-TAR model, the underlying price series might display asymmetry of the following form: when prices are decreasing, the gap between the prices decreases at a faster rate as opposed to the case when both prices are increasing (Ghoshray 2002). In both cases, this type of asymmetric price behavior differs from the symmetric adjustment scenario where both prices move together and any transitory deviation in “levels” or “rates of change” is corrected in a symmetric manner.

This paper examines the extent to which increases in international food prices during the past few years have been transmitted to domestic prices in developing countries. In analyzing the historical data, evidence on price transmission for important food commodities (e.g., rice, wheat, and edible oil) have been considered. The price transmission elasticity has been estimated using regression models coupled with recent econometric techniques such as unit root tests and ECM with threshold adjustment (Enders and Siklos 2001). Finally, the paper draws some policy implications from the empirical results. This study provides the numerical estimates on the empirical relationship between international prices and domestic prices. The analysis uses commodity-specific monthly data rather than annual data during a period of substantial policy reforms in order to understand both long-run and short-run relationships between world and domestic prices.

II. Institutional Background

The subject of trends primary commodities has led to much debate as it has been used to explain the widening gap between developed and less developed countries leading to a large volume of studies that empirically test whether the long-term trends have

been declining over time. The evidence has been mixed, which leads to serious policy implications as to whether or not developing countries should continue to specialize in primary commodity exports. The World Bank for instance, has encouraged long-term primary commodity projects in developing countries (Ardeni and Wright 1992). Besides, free market solutions were provided to developing countries to deal with primary commodities instead of positive intervention (Maizels 1994). The upshot is that countries that rely on the exports of primary commodities must understand the nature of commodity prices in order to devise their development macroeconomic policy (Deaton 1999).

Besides, in the study of the nature of trends in commodity prices, price transmission has received a huge amount of interest from researchers and policy makers. Price transmission has important repercussions for food security. There is no doubt that price transmission can have adverse impacts on the poor, whose marginal propensity to consume on foodstuffs can be quite high. Cutting back on food consumption can lead to lowering of nutritional levels, which had been noted in some recent studies. While the world market price is an important indicator of food scarcity, food-consuming and food-producing households do not buy and sell directly on the world market. Thus, an important question for food security is the extent to which changes in world prices are transmitted, or passed through, to domestic consumers and producers (Dawe 2008a).

Based on recent literature on price transmission Conforti (2004) cites that transactions costs, trade policies, and exchange rates influence price transmission from world prices to domestic prices. Transactions costs act as a wedge between prices in world and domestic markets. If this wedge is greater than the transactions costs then it allows for possible arbitrage to take place, allowing the markets to be integrated. However, these transactions costs are assumed to be stationary and proportional to traded volumes to facilitate the construction of linear models. Relaxing this assumption calls for models that include thresholds (McNew 1996, Barrett and Li 2002, Brooks and Melyukhina 2003). Trade policies affect price transmission through intervention instruments such as nontariff barriers, tariff rate quotas, prohibitive tariffs, etc. Ad valorem tariffs are expected to behave as fixed or proportional transactions costs (Conforti 2004). Finally, exchange rates play an important role in influencing price transmission. For example, following the study on rice by Dawe (2008a), the Philippine peso (P) strengthened against the United States (US) dollar from P54.2 in 2003 to P46.1 in 2007. Thus, the world price as measured in Philippine pesos did not increase at the same rate as the world price as measured in dollars. Thus, for many Asian countries, this headline price increase was illusory because the US dollar was steadily depreciating against a wide range of regional currencies.

In the presence of differing inflation rates, price margins are inflated proportionally over time. Since the model is specified such that the change in the price margin is attributed to a change (reduction) in transaction costs over time, the presence of inflation may lead to a false interpretation of the parameter estimates (van Campenhout 2007). The effect of deflating by the average price level is to take common inflation out. To take account of

this issue, in this study, all international commodity prices are converted to the domestic country price and then adjusted for inflation. For example, in 2007–2008, Cambodia, India, and Viet Nam restricted or banned exports of rice, causing an increase in demand due to expectations that the supply of rice may become more restricted in the future. This fuelled the price of rice to soar (Dawe 2008a). All of these price changes were above and beyond the reported increases in the general consumer price index.

As discussed in the preceding paragraph, trade policies such as variable tariffs can help determine how world prices are transmitted to domestic markets. The relative increases (decreases) in domestic prices in relation to increases (decreases) in international prices can vary depending on whether prices are increasing and decreasing which in turn can be related to government policy and intervention (Ghoshray and Ghosh, forthcoming).

While deviations between international and domestic prices can persist as a result of government intervention, the case of Thailand presents a situation that runs contrary to this observation. While some degree of government intervention in terms of procurement and storage exists, domestic prices nevertheless have been observed to follow world prices closely. Thailand is a net exporter of soybean oil mostly to neighboring countries because of the proximity advantage. In the case of palm oil, although imports have been almost inexistent since 2000, palm oil is subject to an import quota system. These oil prices are controlled by the Ministry of Commerce, fixing a ceiling retail price. The import control system consists of a tariff rate quotas (TRQ) (under the World Trade Organization agreement) with an import quota of 2281 tons in 2009 subject to a 20% tariff rate and an out-of-quota tariff rate of 146%.

Historically, trade policies in India have been characterized by tight controls over imports and exports through, among other things, an import licensing system, which amounted to an import ban for most agricultural commodities (exceptions were made for pulses, cotton, wool, and edible oil). Since 1965, the trade policy related to food grains has been characterized by export/import monopolies from the government agencies and important trade restrictions. Liberalization measures (consisting mainly of relaxation of domestic controls on intermediate and capital goods) started in the late 1970s and early 1980s and gained momentum in 1985–1989 under the government of Rajiv Gandhi; in 1991 following a strong devaluation (70% between 1991 and 1993);¹ in the late 1990s following the Uruguay Round Agreement (the import licensing system started to be dismantled in 1998); then again in April 2003 and in 2007. These reforms impacted manufactured and agricultural goods differently: although in 2007, 80% of industrial tariff rates were below 10%, in 2006, the unweighted average tariff protecting agricultural and processed food commodities was about 40%. Moreover these tariffs are characterized by high dispersion (with 15% above 50%) and redundancy. This translates into domestic prices for many commodities being lower than duty-inclusive import prices (therefore potentially indicating

¹ From 1993, a managed floating exchange rate regime against a basket of currencies was instituted whereby the real effective exchange is allowed to fluctuate within a 10% band around the postdevaluation level.

a lack of integration with world prices). According to Pursell et al. (2009), the philosophy behind India's trade policy concerning agricultural products is that world markets should only be used in case of domestic shortages (imports) and excess domestic surpluses to release stocks (exports). This may translate to highly opportunistic instruments to control trade flows such as variable tariffs and export subsidies (for example, cereals and sugar).

Malaysia follows a relative open trade policy regime. A free trade regime dominated subsidiary crops, horticultural products, livestock, and fishing (only round cabbages were subject to import restrictions and a few products such as forest products and crude oil were subject to export duties). Although export duties on rubber and palm oil were a significant source of government revenues, since the mid-1980s, they have been reduced drastically. Athukorala and Loke (2009) report an average annual duty rate of 4.7% for rubber and 1.1% for palm oil (although it may rise to between 5% and 10% for certain grades of crude palm oil to encourage domestic processing). The USDA Foreign Agricultural Service (2009a) reports export taxes reaching between 10% and 30% for crude palm oil and such practices as export tax exemptions for some large Malaysian companies with joint ventures in foreign countries.

The trade policy in the PRC has made import and export markets less restricted (and more decentralized), the foreign exchange rate regimes were transformed and there were reductions in nontariff barriers (NTBs) and quotas. However, commodities such as rice, wheat, and maize remain under the control of the government. Huang et al.'s (2009) analysis of nominal rates of assistance (NRA) of agricultural commodities (based on the difference between border and domestic prices) shows that a major change occurred in 1995 (for exportables) and more generally in the late 1990s (for import-competing goods). Before 1995, exportables were characterized by lower domestic prices than world prices. This gap became almost negligible for fruits, vegetables, poultry, and pig meat from the late 1990s while the NRA for rice went down from about 30% in the early 1990s to about 7% subsequently. The evolution of the NRA for cotton is similar to the one for rice. In the case of wheat and soybeans, distortions between domestic and border prices were significantly reduced from the early 2000s: the NRAs for wheat and soybeans went from about 30% in the late 1990s to 4% and 16%, respectively, in the early 2000s. In contrast, the NRAs increased over the period for sugar and maize. A floor price program applies to 13 grain-surplus producing regions accounting for 80% of the country's commercialized grains. The floor price is set annually and applies only during a designated period (harvest period, no more than 5 months per year) and rose for rice and wheat in the late 2000s, in part to offset rising input costs. In 2009, market prices for wheat and rice were higher than the floor prices. According to the USDA Foreign Agricultural Service (2009a), between 2006 and 2008, government purchases of wheat at the floor price amounted to about 34% of total wheat production; and in 2008–2009, the planned purchase of maize represented 25% of total output. This program was used to stabilize the prices of the three major grains, the government bearing the costs of storage and marketing. Trade policy tools implemented in the recent years consist of an export tax (2008-09) and a

value-added tax that occasionally makes the object of a rebate to stimulate exports (e.g., 13% rebate prior to 2008). According to the USDA Foreign Agricultural Service (2009a) the quality of wheat and rice produced in the PRC and exported was still comparatively low in the late 2000s, explaining the need to import grains of higher quality to cover the needs of the more affluent coastal cities. TRQs were instituted in the early 2000s (following the PRC's WTO membership) and were stabilized in 2004. State-owned enterprises exert significant amount of market power given that their percentage quota shares amount to 90% for wheat, 60% for maize, and 50% for rice.

Historically, the major agricultural commodities in the Philippines have been subject to a wide range of policy intervention including government monopoly control over imports and exports and domestic marketing operations, import and export bans, import licensing, imports and export tariffs, as well as quantitative trade restrictions on specific commodities. A notable reform effort was initiated in 1986 by the government of Corazon Aquino. The main measures were: (i) elimination of most export taxes and the coconut export ban; (ii) end of the National Food Authority² (NFA)'s monopolistic control over international trade of wheat, soybeans, and meat; and (iii) reduction of the NFA's domestic marketing operations. Quantitative import restrictions still apply to all major import-competing agricultural commodities and were reinforced by the Magna Carta of Small Farmers in 1991. The Uruguay Round Agreement on agriculture in 1994, which advocated the replacement of all NTBs to trade by a tariff equivalent and the reduction of tariff protection in general, led to limited changes. These include (i) rice being exempted from these measures until 2004; (ii) quantitative trade restrictions being removed in April 1996 and replaced by tariff rate quotas, which instituted in-quota tariffs at the same level as previously while out-of-quota were set significantly higher; and (c) tariffs being raised on products that are substitutes for other commodities (for example, wheat and barley for maize) and on processed products using these commodities as raw material.

Rice is the main staple food and a net import for Indonesia. Government policies have been driven by three main goals: to achieve self-sufficiency in food, mainly rice; stabilize food prices; and promote food processing. These policies translated into unprocessed exports being taxed and processing being subsidized; high rates of protection for two import-competing industries, that is, sugar and rice; and increases in the rates of protection for these two commodities from 2000 in the wake of democratic reforms following the Asian financial crisis. The evolution of economic policies in Indonesia was influenced by the worldwide shift from the mid-1980s characterized by a weakening of manufacturing protection and agricultural export taxation. The palm oil industry was protected by a tax on exports. From 1982 the price of oil began to decline, and the government responded by devaluing its currency to promote non-oil exports, introducing a value-added tax in 1983 and 1986 and increasing the protection of import-competing commodities. In 1985–1986 the government started a process of liberalization of the trade

² The National Rice and Corn Administration was first established in 1936. In the 1970s, it became the National Food Authority and its activities were expanded to allow for the tariff-free importation of other agricultural commodities against a context of high world commodity prices.

policy through the reduction of tariffs (1985) and NTB (1986), which were replaced by tariffs and the extension of the system allowing exporters to import inputs duty-free. In 1997–1998 the government eliminated the restrictions on imports of rice and sugar. Both tariffs and NTBs were reintroduced subsequently for these two commodities.

Government policies can be implemented to insulate the domestic country from international price shocks. For instance, depleting stocks can dampen price increases. However, this is a short-term measure as eventually stocks may run out. Another measure could be to lower import tariffs. However, this measure too is a short-term policy as the tariff rate could hit a lower bound at zero. At this stage though it is possible to subsidize imports, as these subsidies may become fiscally unsustainable in the face of increasing world prices. Market expectations can play a role in price transmission. When international prices are rising and the expectation is that they will continue to rise, farmers, traders, and households in response may hoard supplies. This will cause a decrease in domestic supply, forcing domestic prices to rise.

It is difficult to predict how international prices will change for the rest of 2010 and beyond. However, this report will make an attempt to understand the underlying trends and irregular behavior in commodity prices to enable policy makers to ascertain whether price increases will rise to a level that may substantially hurt the poor. This will present challenges for governments to manage these price fluctuations in order to minimize the impact on the poor.

III. Theoretical Background and Previous Research

Research in agricultural trade begins with a problem in need of explanation. Theory and experience are drawn on to list reasonable explanations. Models of trade are formulated during this deductive process. The next step in the research process is often quantitative specification and empirical testing. Using parameter estimates that have proven to be statistically significant, models are employed to predict future conditions, and to measure market impacts of policy decisions.

The continuing interest in trends of primary commodities stems from the fact that the central question is an empirical one. When considering the possibility of the existence of a trend in real commodity prices one is naturally led to the question of whether the prices are trend-stationary or difference-stationary. The standard econometric tool used to discriminate between the two alternatives is the unit root test. The unit root test is useful as it aids in distinguishing whether the real commodity prices are characterized by stochastic trends or not. If the price series is found to contain a unit root then the series is said to contain stochastic trends such that the effect of shocks to the underlying price series will be permanent. If however, the underlying price series is found to reject

a unit root, then the series is considered to be trend-stationary and the effect of shocks on the price series would be transitory in nature. Past studies that have analyzed trends assumed that the underlying data series is trend-stationary. While Sapsford (1985), Grilli and Yang (1988), Helg (1991), and Ardeni and Wright (1992) among others have advocated for commodity prices to follow a trend stationary model, Cuddington and Urzua (1989), Cuddington (1992), Bleaney and Greenaway (1993), and Newbold et al. (2005) recognized that commodity prices may be difference-stationary. The evidence on the trend function has been mixed. While Sapsford (1985) and Helg (1991) tend to support a negative trend, in contrast, Newbold and Vougas (1996) cannot provide compelling evidence as to whether the data is trend-stationary (TS) or difference-stationary (DS), while Kim et al. (2003) suggest that commodity prices generally display unit root behavior and that there is limited evidence for a negative trend in commodity prices.

Perron (1989) showed that if a structural break is ignored the power of the unit root test is lowered. His paper however, was criticized for the fact that he assumed that the date of the structural break is known. Zivot and Andrews (1992) developed a unit root test that allowed for the break to be unknown and determined endogenously from the data. Since the unit root test suffers from low power by ignoring a single structural break, it was argued by Lumsdaine and Papell (1997) that the loss of power is greater if there are two structural breaks. They formulated a method that is an extension of the Zivot–Andrews, which tests for a unit root under the null hypothesis against the alternative of two structural breaks determined endogenously from the data.

The issue of structural breaks applied to primary commodity prices has been of great interest to researchers. Leon and Soto (1997) applied the single break Zivot–Andrews test on primary commodity prices and found evidence of structural change. In contrast to Kim et al. (2003) their results show that most commodity prices can be described as trend stationary models and that the trend coefficients generally characterize a negative trend. Zanas (2005) and Kellard and Wohar (2006) have employed the Lumsdaine–Papell test to allow for two structural breaks. Zanas employs the test to the aggregate Grilli Yang index and concludes the data to be a trend-stationary process with two intercept shifts. The breaks are identified in 1920 and 1984, which cumulatively account for a 62% drop. However, Zanas (2005) considered the aggregate index, which has been questioned by recent studies, arguing that individual commodities behave in a different manner. The study by Kellard and Wohar (2006), KW hereafter, is a case in point. KW conduct a study of the disaggregated commodity prices over the period 1900–1998. Out of 24 commodity prices, their results indicate 14 are trend-stationary allowing for one or two structural breaks. Overall, they show that the deterioration of commodity prices has been discontinuous. Following Leon and Soto (1997), KW state that a single trend is a “summary measure” of several trends that may be positive or negative. Arguing that reliance on a single trend may be misleading to policy makers, KW create a measure to define the prevalence of a trend. A drawback of the past studies is the choice of a null hypothesis that allows for a linear unit root. This has been recognized with further

examination by Ghoshray (2010). Further, the approach of the prevalence of trends due to KW has been adopted to analyze the historical trends in commodity prices.

Price transmission can be defined as the relationship between prices in two related markets; for example, international and domestic markets. The principle has its roots in the law of one price (LOP), which states that the difference between the two prices in spatially separated markets should not exceed the cost of transportation of the commodity in question from one market to the other market. Theoretically, the price transmission elasticity of international to domestic prices is an application of the LOP. According to the LOP, the price of a traded good will be the same in both domestic and foreign economies, when expressed at a common currency.

Price transmission refers to the effect of prices in one market on the prices of another market. It is generally measured as the price transmission elasticity, which is the percentage change in price of one market to a given percentage change in price of another market. If such relationship between two prices, such as international and domestic prices, holds in the long run, the markets can be said to be integrated. This relationship may not hold in the short run. On the other hand, the two prices may be completely independent, leading one to conclude that there is no market integration or price transmission. The long-run relationship between international and domestic prices that implies market integration lends itself to a cointegration interpretation with its existence tested by cointegration methods. If two prices are found to be cointegrated, there is a tendency for both prices to co-move over time in the long run. In the short run there may be deviations that can be driven by shocks in one price not being transmitted to the other price; however arbitrage would make these deviations transitory and prices are brought back to their long-run equilibrium over time.

So far the assumption is that the price in one market influences the price of the other market in a symmetric fashion. However, when analyzing the price relationship between international commodity prices and domestic prices in Asia, there may be a case of asymmetry in price transmission. Minot (2010) argues that domestic prices are unlikely to have a noticeable impact on world prices, but world prices in turn can exert a substantial influence on domestic prices. Besides, the share of world exports can cause a “small” domestic country to have very little measurable impact on world commodity prices due to the small share of exports of the domestic small country relative to a “large” country (Ghoshray 2006).

The literature on horizontal price linkages, that is, links between prices at different locations, has been typically concerned with spatial price relationships. The underlying empirical literature has studied these price relationships under the concept of the LOP. The study of price transmission has been motivated largely by the view that co-movement of prices in different markets can be interpreted as a sign of efficient markets, while the absence of price co-movement can be viewed as a sign of market failure. A large number

of studies have examined price co-movement within a country (see Ravallion 1986, Palaskas and Harriss-White 1993, Abdulai 2000, Ghosh 2003, Kuiper et al. 2003, Meyers 2008); as well as the transmission of prices from world markets to domestic markets in developing countries (see Mundlak and Larson 1992, Quiroz and Soto 1995, Conforti 2004, Minot 2010).

Price transmission studies are seen as an empirical exercise. These studies attempt to test the predictions of economic theory to demonstrate how changes in the prices of one market are transmitted to another, which provides information on the extent of the market and whether markets are efficient. Inference can be drawn in determining the magnitude and distribution of welfare effects of trade policy scenarios. Modifying and finetuning these mechanisms of price transmission are therefore important in addressing these policies.

Early studies of price transmission used simple correlation coefficients of contemporaneous prices. A high correlation coefficient is evidence of co-movement and was often interpreted as a sign of an efficient market. In an integrated market system, the coefficient of correlation should be positive, indicating a common positive relationship over time between the price series. That is the prices tend to move in the same direction. The higher the coefficient, the stronger is the linear association of the prices. Lele (1971) for instance, used the correlation coefficient approach to study weekly sorghum prices in Western India. Timmer (1987) calculated the spatial correlation coefficients between paddy prices in various provinces in Indonesia and found the coefficients to be uniformly high, even on the outer islands. Stigler and Sherwin (1985) used the correlation coefficients on the logarithm of prices to determine the range of product markets and substitution effects between products. However, the correlation coefficients approach has important weaknesses as a method to test for market integration. The correlation coefficient does not provide a reliable indication to illustrate the extent to which markets are integrated. More generally, the correlation coefficient is simply a measure of linear association among stationary processes. Nonstationarity and nonlinearity reduce its efficiency. Besides, another important shortcoming of the approach is that it involves the influences of common components such as inflation, climatic patterns, and population growth (Abdulai 2006).

Following the criticism levelled at the correlation approach, regression-based approaches were employed. Mundlak and Larson (1992) estimated a regression on contemporaneous prices, with the regression coefficient being a measure of the co-movement of prices. Mundlak and Larson estimated the transmission of international food prices to domestic prices in 58 countries using annual price data from the FAO. Their findings showed estimates of price transmission elasticity to be approximately 0.95, implying that 95% of any change in world markets was transmitted to domestic markets. However, the static regression approach has been criticized for assuming instantaneous response in one market to changes in other markets. In fact, there is generally a lag between the

price change in one market and the impact on another market due to the time it takes traders to notice the change and respond to it. A change in world prices may take more than a month to be reflected in domestic prices. These dynamic effects can be captured by including lagged world prices as explanatory variables in the regression analysis (Ravallion 1986).

However, a major shortcoming of these studies is the nonstationarity of variables in these regressions and the inappropriate application of transformations on these variables. After the seminal paper by Granger and Newbold (1974) on spurious regressions and formal tests for cointegration due to Engle and Granger (1987), standard linear regressions came under scrutiny. Standard regression analysis assumes that the mean and variance of the variables are constant over time. This implies that the variable tends to return toward its mean value, so the best estimate of the future value of a variable is its mean value. However, in the analysis of time-series data, prices and many other variables are often non-stationary, meaning that they drift randomly rather than tending to return to a mean value. One implication of this “random walk” behavior is that the best estimate of the future price is the current price. When standard regression analysis is carried out with nonstationary variables, the estimated coefficients are unbiased but the distribution of the error is non-normal, so the usual tests of statistical significance are invalid. In fact, with a large enough sample, any pair of non-stationary variables will appear to have a statistically significant relationship, even if they are actually unrelated to each other (Granger and Newbold 1974, Phillips 1987). Following the research into nonstationary variables, Quiroz and Soto (1995) replicated the analysis of Mundlak and Larson (1992) with similar data but using the error-correction model. They found no relationship between domestic and international prices for 30 of the 78 countries examined, and even in countries with a relationship, the convergence was very slow in many of them.

Morisset (1998) examines the spread between international and domestic prices. The methodology is based on that of Mundlak and Larson (1992), which goes a step further allowing for asymmetries to include upward or downward changes in world prices by introducing two zero-one dummy variables in the model. The results show that when world prices are declining the price changes were not transmitted to domestic prices. However, increases in world prices were transmitted to domestic prices.

Baffes and Gardner (2003) also employ the ECM to examine the responsiveness of domestic prices to international prices. The authors analyze a total of 31 country/commodity pairs from the mid-1980s to the early 1990s. The analysis was conducted with and without structural breaks. Their results conclude that the price transmission elasticity is generally low.

Conforti (2004) examined price transmission in 16 countries, including three in sub-Saharan Africa, using the ECM. The countries chosen span three different continents being Asia, Latin America, and Africa. The study concluded that price transmission was

relatively complete in Asia, mixed in Latin America, and low in Africa. In Ethiopia, he found statistically significant long-run relationships between world and local prices in four out of seven cases, including retail prices of wheat, sorghum, and maize. In Ghana, there was a long-run relationship between international and local wheat prices, but no such relationship for maize and sorghum. In Senegal, he found a long-run relationship in the case of rice, but not maize. In general, the degree of price transmission in the sub-Saharan African countries was less than in the Asian and Latin American countries.

This report builds on the dynamic approach of Baffes and Gardner and the asymmetric mechanism of transmission (Morisset 1998, Abdulai 2000 and 2006, Ghoshray 2002 and 2008). An interesting question is whether both world price increases and decreases are transmitted in a symmetric manner to domestic prices.

Statistical models that take into account nonstationarity face another problem. The lack of price integration does not necessarily imply inefficient markets or policy barriers to trade. One econometric approach to deal with this situation is threshold autoregressive (TAR) models and M-TAR models (Enders and Siklos 2001). These models take into account thresholds caused by transactions costs that deviations must exceed before provoking equilibrating price adjustments that lead to market integration. When shocks to prices occur such that the magnitude of the shock exceeds a certain threshold, the response to that shock will be different if the shock had been smaller and below the threshold. This model allows a researcher to investigate the degree to which the market violates the spatial arbitrage condition, as well as the measure of the speed with which it eliminates these violations. Studies that have employed such models include Abdulai (2000), Zapata and Gauthier (2003), Abdulai (2006), Ghoshray (2007), Ghoshray (2008). The manner in which international prices of commodities affect final domestic prices is central to recent macroeconomic analysis as the issue is closely related with globalization and economic integration. The degree and timing of price transmission is important for developing countries, in terms of formulating monetary policy in response to shocks in international prices.

IV. Econometric Methodology

This section describes the econometric methods employed to analyze the nature of the underlying trends of the commodity prices in this study as well as the methods used to measure international price transmission.

A. Econometric Methodology: Estimating Trends

When considering unit root tests that allow for structural breaks, there may be a case of size distortion that leads to spurious rejection of the null hypothesis of a unit root when

the actual time series process contains a unit root with a structural break (Nunes et al. 1997). The test developed by Lumsdaine and Papell (1997) suffers from the same size distortion problems and consequent spurious rejection of the null hypothesis (Lee and Strazicich 2003). The minimum two-break Lagrange Multiplier (LM) test developed by Lee and Strazicich (2003) allows for structural break under the null hypothesis and does not suffer from the spurious rejection of the null hypothesis. Besides, the minimum LM test possesses greater power than the Lumsdaine–Papell test. The upshot is that under the LM test setting, rejection of the null hypothesis can be considered as genuine evidence of stationarity. A further disadvantage of the Zivot–Andrews test and the Lumsdaine–Papell test is that the tests tend to estimate the breakpoint incorrectly where bias in estimating the unit root test is the greatest (Lee et al. 2006). This leads to size distortion, which increases with the magnitude of the break. This size distortion does not occur when using the LM test as it employs a different detrending method (Lee et al. 2006).

The commodity price series under consideration may be trend-stationary. If the underlying commodity price series were to be trend-stationary, then to test the nature of the underlying trends, one typically estimates the following log-linear time trend model:

$$P_t = \alpha + \beta t + \varepsilon_t \quad (1)$$

where P_t is the log ratio of commodity price series to the consumer price index and the errors denoted by ε_t is a process integrated of order zero, or $I(0)$ and is assumed to follow an autoregressive moving average (ARMA) process to allow for cyclical fluctuations of relative commodity prices around their long-run trend. If the commodity price series P_t were to contain a unit root, or in other words were to be integrated of order one, that is $I(1)$, then estimating the trend stationary model given by (1) will generate spurious estimates of the trend, by concluding that the trend is significant when it is actually not. A negative estimate of β , which is statistically significant, implies that the underlying price series follows a negative trend. An appropriate strategy for estimating the trend is to adopt the following difference-stationary model:

$$\Delta P_t = \beta + \eta_t \quad (2)$$

where η_t is a stationary error process. In the difference-stationary model, if we find a negative estimate of β that is statistically significant, we may conclude that the underlying price series follows a negative trend. An important point to note is that if the real commodity price series is a trend-stationary process but is treated as a difference-stationary process, then tests based on equation (2) are inefficient, lacking power relative to those based on equation (1).

If structural break(s) are ignored, the power of the unit root test is lowered. In this paper we consider the unit root test subject to two endogenously determined structural breaks (Lee and Strazicich 2003) and a single structural break (Lee and Strazicich 2004). The

Lee–Strazicich unit root test determines the break points endogenously by utilizing a grid search (details of the test can be found in the Appendix).

If we find that we can identify up to two structural breaks in any price series, in the next step we attempt to determine whether the sign of the trend is negative or positive and whether it is significant or not. Besides, it would be of interest to observe whether the trend changes signs in the different regimes that are outlined by the structural breaks. In order to determine how the trend has shifted over time we proceed to model the data generating process in the manner conducted by KW. For the trend-stationary process the logarithm of the commodity price series is regressed against a constant and a time trend and one or two intercept and slope dummy variables corresponding to the results of the structural break tests. The error structure is modeled as an ARMA(p,q) process (details provided in the Appendix).

Following KW, the trend function for price is reparameterized to facilitate the estimation of the trend coefficient for up to three different regimes. For the three-regime model, P_t^* is defined as:

$$P_t^* = \begin{cases} P_t & \text{if } 1 \leq t \leq TB1 \\ P_t - \delta_1(TB1 - TB0) & \text{if } TB1+1 \leq t \leq TB2 \\ P_t - \delta_1(TB1 - TB0) - \delta_3(TB2 - TB1) & \text{if } t > TB2 \end{cases} \quad (3)$$

In the above model (3), δ_1 and δ_3 denote the slope coefficients of price in different regimes. $TB0$ denotes the starting point of the data, whereas $TB1$ and $TB2$ denote the break points selected by the Lee-Strazicich test. To facilitate comparison with KW, the estimation was conducted by exact maximum likelihood and the ARMA order (p,q) was selected through the Schwartz Bayesian Criteria (SBC) allowing all possible models with $p + q \leq 6$.

B. Econometric Methodology: Price Transmission

To formalize the theory, the Engle and Granger (1987) two-step method to test for cointegration between two prices, say world prices, P_t^W and domestic prices, P_t^D entails using ordinary least squares (OLS) to estimate the long-run relation of the two prices. This is given by the equation below:

$$P_t^D = \alpha + \beta P_t^W + \varepsilon_t \quad (4)$$

where P_t^W and P_t^D are nonstationary I(1) prices, α and β are the estimated parameters. α is an arbitrary constant that accounts for the differential (transfer costs and quality differences), β denotes the price transmission elasticity and ε_t is the error term that may

be serially correlated. To test the hypothesis, equation (4) is estimated using OLS. The second step advocates a Dickey-Fuller test on the estimated residuals of equation (4) of the following kind:

$$\Delta \varepsilon_t = \gamma \varepsilon_{t-1} + \omega_t \quad (5)$$

where ω_t is a white noise error term. If ω_t is not white noise, an Augmented Dickey Fuller (ADF) test may be used where lagged values of $\Delta \varepsilon_t$ may be added to equation (5). Rejecting the null hypothesis ($H_0 : \gamma = 0$) of no cointegration implies that the residuals of equation (5) are stationary. Thus equation (4) is like an attractor such that its pull is strictly proportional to the absolute value of ε_t .

However, given the arguments made earlier about price transmission that may be asymmetric, an alternative method of adjustment may be proposed. Besides, from an econometric point of view, Enders and Siklos (2001) argue that the test for cointegration and its extensions are mis-specified if adjustment is asymmetric. They consider an alternative specification, called the threshold autoregressive (TAR) model, such that equation (5) can be written as:

$$\Delta \varepsilon_t = I_t \gamma_1 \varepsilon_{t-1} + (1 - I_t) \gamma_2 \varepsilon_{t-1} + \omega_t \quad (6)$$

where I_t is the Heaviside indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } \varepsilon_{t-1} \geq \tau \\ 0 & \text{if } \varepsilon_{t-1} < \tau \end{cases} \quad (7)$$

This specification allows for asymmetric adjustment. If the system is convergent, then the long-run equilibrium value of the sequence is given by $\varepsilon_t = \tau$, where τ is the estimated threshold. The sufficient conditions for the stationarity of ε_t are $\gamma_1 < 0$, $\gamma_2 < 0$ and $(1 + \gamma_1)(1 + \gamma_2) < 1$ (Petrucelli and Woolford 1984). In this case if ε_{t-1} is above its long-run equilibrium value, then adjustment is at the rate γ_1 and if ε_{t-1} is below long-run equilibrium then adjustment is at the rate γ_2 . The adjustment would be symmetric if $\gamma_1 = \gamma_2$. However, if the null hypothesis $H_0 : (\gamma_1 = \gamma_2)$ is rejected then using the TAR model we can capture signs of asymmetry. If for example, $-1 < \gamma_1 < \gamma_2 < 0$, then the negative phase of the deviation between the two prices will tend to be more persistent than the positive phase.

In the TAR model it is necessary to estimate the value of the threshold that will be equal to the cointegrating vector. A method of searching for a super-consistent estimate of the threshold was undertaken by using a method proposed by Chan (1993); see Appendix for details.

In equation (7) the Heaviside Indicator depends on the level of ε_t . Enders and Siklos (2001) suggest an alternative such that the threshold depends on the previous periods

change in ε_t . The Heaviside Indicator given by equation (7) can be set to the Momentum-Heaviside Indicator as:

$$I_t = \begin{cases} 1 & \text{if } \Delta\varepsilon_{t-1} \geq \tau \\ 0 & \text{if } \Delta\varepsilon_{t-1} < \tau \end{cases} \quad (8)$$

In this case the series ε_t exhibits more momentum in one direction than the other. The model given by equation (6) along with equation (8) depicts the M-TAR model. The M-TAR model can be used to capture a different type of asymmetry. If for example, $|\gamma_1| < |\gamma_2|$, the M-TAR model exhibits little adjustment for positive $\Delta\varepsilon_{t-1}$ but substantial decay for negative $\Delta\varepsilon_{t-1}$. In other words, increases tend to persist, but decreases tend to revert quickly back to the attractor irrespective of where disequilibrium is relative to the attractor. To estimate the threshold, Chan's methodology is employed (details are in the Appendix).

To implement in this test the case of the TAR or M-TAR adjustment the Heaviside Indicator function is set according to equation (7) or equation (8), respectively, and equation (6) estimated accordingly. The Φ -statistic for the null hypothesis of stationarity of ε_t i.e., $H_0 : (\gamma_1 = \gamma_2 = 0)$ is recorded. The value of the Φ -statistic is compared to the critical values computed by Enders and Siklos (2001). If we can reject the null hypothesis, it is possible to test for asymmetric adjustment since γ_1 and γ_2 converge to multivariate normal distributions (Tong 1990). The F statistic is used to test for the null hypothesis of symmetric adjustment, that is, $H_0 : (\gamma_1 = \gamma_2)$. Diagnostic checking of the residuals are undertaken to ascertain whether the ω_t series is a white noise process using the Ljung-Box Q tests. If the residuals are correlated, equation (6) needs to be re-estimated in the form:

$$\Delta\varepsilon_t = I_t\gamma_1\varepsilon_{t-1} + (1-I_t)\gamma_2\varepsilon_{t-1} + \sum_{i=1}^p \phi_i \Delta\varepsilon_{t-i} + \omega_t \quad (9)$$

Equation (9) is comparable to (6) except that it incorporates lagged first differences of the dependent variable to correct for the autocorrelation in the error term ω_t . The SBC is used to determine the lag length.

The positive finding of cointegration with threshold adjustment justifies the estimation of the following threshold error correction model, which takes the form:

$$\Delta P_t^D = \mu_1 ECM_{t-1}^+ + \mu_2 ECM_{t-1}^- + \pi \Delta P_t^W \sum_{i=1}^p \psi_i \Delta P_{t-i}^W + \sum_{i=1}^p \delta_i \Delta P_{t-i}^D + u_t^W \quad (10)$$

where P^W and P^D denote the world price and the domestic price, respectively, ECM refers to the error correction term, and both u_t^W is a white noise error. Equation (10) describes the dynamics of world prices by nesting the short-run and long-run adjustments. The parameters μ_1 and μ_2 denote the error correction coefficients. If there is a deviation from long-run equilibrium, and the deviation happens to be positive, depending on the Heaviside Indicator, then the speed of adjustment is given by μ_1 . Similarly for negative

deviations defined by the Heaviside Indicator, the speed of adjustment is given by μ_2 . The parameter estimate π denotes the short-run price transmission elasticity, that is, the immediate impact of a change in world price on domestic price. Lagged differenced variables of P^W and P^D are included to ensure that the errors are white noise and the model is well specified. The number of lags p may be determined using the SBC.

V. Data and Empirics

This section describes the data used in this study and the empirical results obtained from the econometric analysis.

A. Description of the Data

Table 1 provides a description of the basic statistics of commodity prices employed in this study.³ Domestic price series for India (except for sugar), Philippines (except for rubber), the PRC (except for beef and rubber), and Thailand (except for rubber) do not seem to exhibit a marked trend over the period considered. However it should be noted that the rubber price series covers a shorter period, usually from the late 1990s to 2010. World prices span a much longer period of around 50 years and exhibit overall a slight decreasing trend. Since the 1990s, deflated world prices do not show any obvious trend, except maybe for a slight decreasing trend in the beef price.

The coefficients of variation of domestic prices vary from 0.09 to 0.48 (average 0.21) and those for the world prices from 0.39 to 0.95 (average 0.55). Although they do not correspond to the same period of time, world prices exhibit therefore a much more significant monthly volatility than domestic prices in real terms. Indian prices tend to have a lower variability, as well as commodities such as rice, wheat, sugar, and soybean oil. The volatility of the monthly prices of rubber, coffee, and palm oil is higher (between 0.30 and 0.48) and is reflected in the presence of spikes in the graphs of the variables (given below). Although most of the coefficients of variation of world prices are between 0.40 and 0.60, the prices of sugar and coffee stand out as being highly volatile (0.95 and 0.70, respectively). The high volatility of world prices of sugar may explain the high rates of protection of domestic sugar industries in most of these countries, which in turn may explain the low relative volatility of the domestic prices of sugar. The world price of beef appears to be the least volatile (0.39).

³ Further description of the data can be found in the Appendix.

Table 1: Descriptive Statistics

| | Mean | Coefficient of Variation | Skewness | Kurtosis |
|------------------------|--------|--------------------------|----------|----------|
| Domestic Prices | | | | |
| Rice (India) | 116.14 | 0.10 | 0.18 | 1.64 |
| Wheat (India) | 96.04 | 0.09 | 1.10 | 5.10 |
| Sugar (India) | 187.44 | 0.18 | 0.42 | 3.29 |
| Beef (Thai) | 869.50 | 0.11 | 0.08 | 1.59 |
| Rubber (Thai) | 352.80 | 0.43 | 1.24 | 4.15 |
| Soybean (Thai) | 332.45 | 0.18 | 0.04 | 2.50 |
| Palm Oil (Thai) | 184.75 | 0.31 | 0.67 | 2.79 |
| Coconut Oil (Thai) | 234.60 | 0.26 | 0.12 | 3.78 |
| Rice (Philippines) | 229.24 | 0.11 | 1.01 | 3.73 |
| Corn (Philippines) | 113.51 | 0.12 | 1.16 | 5.84 |
| Rubber (Philippines) | 208.72 | 0.48 | 0.56 | 2.18 |
| Coffee (Philippines) | 535.99 | 0.35 | 1.21 | 4.52 |
| Rubber (PRC) | 141.61 | 0.29 | 0.27 | 2.04 |
| Beef (PRC) | 194.79 | 0.21 | 0.41 | 1.39 |
| Rice (PRC) | 27.82 | 0.12 | -0.50 | 1.70 |
| Sugar (PRC) | 50.64 | 0.11 | -0.22 | 2.09 |
| Soybean Oil (PRC) | 74.87 | 0.16 | 1.93 | 7.01 |
| Rice (Indonesia) | | | | |
| World Prices | | | | |
| Rice | 3.45 | 0.58 | 1.33 | 6.05 |
| Wheat | 1.55 | 0.43 | 1.13 | 5.40 |
| Corn | 1.22 | 0.44 | 0.47 | 2.55 |
| Sugar | 2.27 | 0.95 | 3.83 | 23.88 |
| Soybean Oil | 5.24 | 0.47 | 1.31 | 5.55 |
| Palm Oil | 4.47 | 0.50 | 0.75 | 3.44 |
| Coconut Oil | 6.97 | 0.57 | 1.22 | 5.70 |
| Coffee | 19.03 | 0.70 | 2.11 | 11.04 |
| Beef | 20.13 | 0.39 | 0.50 | 2.74 |
| Rubber | 10.27 | 0.49 | 0.89 | 3.49 |

Source: Author's calculations.

Positive skewness dominates the results for both domestic and world prices, although some slight negative skewness is found for the domestic prices of rice and sugar in the PRC. Skewness appears to be generally of the same magnitude for domestic prices and world prices. Regarding domestic prices, the highest skewness is found for the PRC retail price of soybean oil (1.93) and for Philippine prices in general. Turning to world prices, prices series for which upward spikes are more important than downward spikes are those of sugar (3.83) and coffee (2.11) once again. The prices of rice wheat, soybean oil, and coconut oil exhibit as well significant positive skewness. In general, positive

skewness means that these series are dominated by episodes of high prices, which are not offset by episodes of low prices of the same magnitude.

Kurtosis varies from 1.39 to 7.01 for domestic prices and from 2.55 to 23.88 for world prices and appears therefore significantly higher in the world prices. High kurtosis is usually associated with high skewness and characterizes the PRC soybean oil prices (7.01), the prices of maize and coffee in the Philippines, of wheat in India, and of rubber in Thailand. In the case of world prices, sugar and coffee are still characterized by high kurtosis (23.88 and 11.04 respectively), followed by rice, wheat, soybean oil, and coconut oil. A possible interpretation of high kurtosis is that more of the variability is explained by a few extreme events. It could then also be construed as a measure of uncertainty. This seems to characterize well both domestic and world prices.

To summarize, commodity prices seem to be well described by positive skewness and kurtosis; they are well described by extreme events and tend to increase more than they fall. World prices appear to be more volatile and exhibit more kurtosis. Sugar and coffee price are characterized by high coefficients of variation, high skewness, and high kurtosis. These characteristics are found in the domestic prices of coffee in the Philippines, but not in the domestic prices of sugar in the PRC and India. However, the domestic price of rubber in Thailand exhibits the same pattern. In general, the domestic prices in the PRC and India are characterized by less monthly volatility, less skewness, and less kurtosis. This is also the case of the domestic and world prices of beef.

B. Trends in International Commodity Prices

This report analyzes the historical trends by using the prevalence of trends approach due to KW and the approach of allowing for breaks to appear in the null and alternative hypothesis in commodity prices based on a study by Ghoshray (2010). Table 2 describes the minimum two and one break LM test results. First, all the data are tested for a unit root under the alternative of stationarity allowing for a break in the null and the alternative hypothesis. The three columns under the heading “LM test with two breaks” tabulate the LM test statistic, and the first and second structural breaks denoted by TB1 and TB2, respectively. Six out of the 10 international commodity prices (soybean oil, palm oil, coconut oil, beef, wheat, and tea) are found to reject the null hypothesis of a unit root. The implication is that these commodities are trend-stationary processes, such that any shocks that occur to these prices tend to be transitory in nature. The break points are found to be significant in all cases except rice, sugar, corn, and coffee. For these four commodities where the null hypothesis of a unit root is not rejected, a further test is made to test for the presence of a single structural break. The two columns under the heading “LM test with one break” tabulate the LM test statistic and the single structural break denoted by TB. The results show that though the null cannot be rejected for all the four commodity prices, the break points for two of the commodities (corn and coffee) are significant. For the other two commodities (rice and sugar) we conclude that the prices

are characterized as unit root processes with two structural breaks, since the breaks were significant under the LM test with two breaks. In this case the conclusion is that two commodities (rice and sugar) are characterized as difference-stationary processes with two structural breaks and the other two commodities (corn and coffee) are difference-stationary processes with a single structural breaks. This implies that any shock to these prices would tend to be permanent in nature. The final column of Table 2 summarizes the results from both the tests. For all commodities we find evidence of at least a single structural break in the trend. Four of the 10 commodities (rice, sugar, corn, and coffee) are difference-stationary processes and the remaining six commodities (soybean oil, palm oil, coconut oil, beef, wheat, and tea) are trend-stationary processes.

Table 2: Structural Breaks in International Commodity Prices

| | LM test with 2 breaks | | | LM test with 1 break | | Conclusion DS/TS(B) |
|-------------|-----------------------|--------|--------|----------------------|--------|------------------------|
| | LM-Stat | TB1 | TB2 | LM-Stat | TB | |
| Rice | -4.26 (6) | Jun-73 | Feb-03 | -3.29 (4) | Jul-81 | DS(B) |
| Soybean Oil | -5.93 (11)** | Mar-73 | Nov-00 | | | TS (B) |
| Sugar | -5.05 (9) | Nov-71 | Jan-82 | -4.29 (9) | Jan-82 | DS(B) |
| Palm Oil | -5.58(11)* | Feb-73 | Nov-85 | | | TS (B) |
| Coconut Oil | -6.93 (10)** | Jun-77 | Dec-86 | | | TS (B) |
| Beef | -6.06 (11)** | Oct-71 | Mar-95 | | | TS (B) |
| Corn | -4.97 (1) | Apr-73 | Jun-99 | -4.14 (1) | Apr-85 | DS(B) |
| Coffee | -5.04 (10) | Oct-75 | Dec-01 | -3.75 (10) | Oct-79 | DS(B) |
| Wheat | -5.34 (11)* | Sep-72 | Mar-00 | | | TS (B) |
| Tea | -5.34 (7)* | Jan-97 | Jun-01 | | | TS (B) |

** and * denote significance at the 5% and 10% levels, respectively.

DS = difference-stationary, TS = trend-stationary.

Note: TB1, TB2 denote the first and second structural breaks, respectively.

Source: Author's calculations.

Based on this information a test is carried out to determine the size and direction of the trends. The results of this test are contained in Table 3. Over the entire sample period we can identify eight of the 10 commodities to have two breaks in the trends and therefore three regimes that characterize the broken trends. For instance, in the case of rice, in the first regime (January 1967 to June 1973), the trend is negative, with the declining rate estimated as 0.87%. This decline continues in the second regime (July 1973 to February 2003) but at a slower rate of 38%. In the final regime (March 2003 to February 2010) the direction of the trend changes, and we can observe a growth of rice prices of 1.05%. Soybean oil shows a trendless behavior in the first regime (January 1967 to March 1973) followed by a negative trend in the second regime (April 1973 to November 2000), and finally a positive trend in the third regime (December 2000 to February 2010). Palm oil and coconut oil show very similar trends. Though the break dates are different, both show no trend in the first regime, a negative trend in the second regime, and finally a trendless behavior in the last regime. Sugar displays a negative trend in all three regimes; however, the rate of decline in sugar prices varies over the three regimes. In the first

regime, the decline is 0.25%, followed by a relative rapid decline of 0.39% in the second regime. Finally, in the last regime, the decline is reduced to 0.11%. Tea on the other hand displays a trendless behavior in all three regimes. While the trend estimates show differences in signs and magnitude, none of these estimates are found to be significant. Wheat displays a trendless behavior in the first regime followed by a negative trend in the second. The final regime shows a positive trend since March 2000. Corn and coffee are found to possess a single structural break and therefore two regimes. Corn displays a negative trend in both regimes and the difference in the trend estimate appears to be marginal. Coffee displays a trendless path in the first regime followed by a negative trend in the second. Both commodities experience a single structural break, and display a decline in prices in the last regime.

Table 3: Trend Estimates from Trend/Difference-Stationary Models

| | TB1 | TB2 | Regime 1 | Regime 2 | Regime 3 | ARMA |
|-------------|------------|------------|-----------------|-----------------|-----------------|-------------|
| Rice | Jun-73 | Feb-03 | -0.87 (6.40) | -0.38 (26.9) | 1.05 (8.64) | (0,1) |
| Soybean Oil | Mar-73 | Nov-00 | 0.13 (0.93) | -0.35 (5.88) | 0.48 (2.16) | (1,1) |
| Sugar | Nov-71 | Jan-82 | -0.25 (2.90) | -0.39 (3.70) | -0.11 (4.46) | (0,1) |
| Palm Oil | Feb-73 | Nov-85 | 0.12 (0.50) | -0.60 (2.72) | -0.07 (0.63) | (1,1) |
| Coconut Oil | Jun-77 | Dec-86 | 0.05 (0.48) | -0.74 (4.07) | -0.08 (1.02) | (3,3) |
| Beef | Oct-71 | Mar-95 | 0.39 (3.95) | -0.34 (8.45) | -0.06 (0.78) | (1,2) |
| Corn | Apr-85 | NA | -0.18 (1.87) | -0.19 (1.87) | NA | (1,0) |
| Coffee | Oct-79 | NA | 0.20 (1.00) | -0.49 (3.60) | NA | (1,0) |
| Wheat | Sep-72 | Mar-00 | -0.07 (0.51) | -0.34 (6.51) | 0.30 (1.69) | (2,0) |
| Tea | Jan-9 | Jun-01 | -0.16 (0.45) | -0.41 (1.36) | 0.20 (1.41) | (1,2) |

Note: TB1, TB2 denote the first and second structural breaks, respectively.

Source: Author's calculations.

Export restrictions may explain the trendless behavior, in that when export restrictions are put in place by countries, other countries can experience price increases that may lead to differing price trends across countries. Significant trends are observed for commodities such as rice and palm oil, which are highly tradable. These results should be interpreted with some caution. The sample size is quite small given the lack of availability of reliable data. Also over the year it is possible that large variations in prices may have been caused by adverse weather conditions. Alternatively it is possible that world prices are transmitted at a faster rate when prices are unusually higher or depressed than normal; or alternatively, when prices are increasing or decreasing at different rates. The analysis will now focus on long-term relations between international and domestic prices based on asymmetric adjustment discussed earlier.

C. Econometric Analysis of Price Transmission

The postulated long-run relationship between international commodity prices and domestic prices can be determined by conducting a cointegration test for commodity

prices that exhibit stochastic trends, in this case, are integrated of order one, that is $I(1)$.⁴ Three tests for cointegration are employed. The first is the standard linear cointegration test due to Engle and Granger (1987) that has been tested by Mundlak and Larson (1992), Quiroz and Soto (1995), and more recently by Sharma (2002) and Minot (2010). The second and third approaches are cointegration with threshold adjustment due to Enders and Siklos (2001) and are quite novel applications in the issue of international price transmission in agriculture. The approach has been employed to study market integration and co-movement of prices (see Abdulai 2001, Ghoshray 2002 and 2008).

The second column of Table 4 produces the results of the standard linear cointegration test due to Engle and Granger by reporting the ADF t-statistic based on the autoregressive coefficient γ that tests the null hypothesis, that is, $H_0[\gamma = 0]$, of no cointegration. Out of the 13 possible long-run relationships that may exist between international and domestic prices, we find seven commodities—wheat (India), tea (India), rice (Philippines), coffee (Philippines), beef (Thailand), rubber (Malaysia), and rice (Indonesia)—reject the null hypothesis of no cointegration at the 10% significance level, thereby implying a long-run relationship between the domestic and international prices of those commodities. However, as argued earlier, the test for cointegration and its extensions are misspecified if adjustment is asymmetric (Enders and Siklos 2001). They consider an alternative specification, the TAR and the M-TAR, which we employ in this study. Column 3 of Table 4 provides estimates of autoregressive decay in the TAR model. The autoregressive coefficient of decay γ in this case is split into two separate components γ_1 and γ_2 to pick up any asymmetries in the adjustment to long-run equilibrium. The estimates record the correct signs as expected for stationarity of the error term. The Φ -statistic tests the null hypothesis, that is, $H_0[\gamma_1 = \gamma_2 = 0]$, of no cointegration. In the case of the TAR model specification, we find that the null can be rejected for six commodities, five of which match those of the Engle–Granger test. However, two commodities—rubber (Malaysia) and rice (Indonesia)—do not show any asymmetry as we cannot reject the null hypothesis of symmetric adjustment that is $H_0[\gamma_1 = \gamma_2]$. The commodity that shows cointegration with asymmetry and is not picked up by the Engle–Granger test is rice (the PRC), which produces a Φ -statistic of 7.29 that rejects the null hypothesis at the 5% significance level. The probability value in square brackets clearly indicates that asymmetry exists. In this case, the estimated coefficients γ_1 and γ_2 are negative, thereby displaying the correct signs and differing in magnitude, $\gamma_1 > \gamma_2$. The coefficient γ_1 is insignificant, which suggests that when positive deviations from the long-run regression appear, there is no significant adjustment taking place. However, when deviations are negative, then a proportion of the deviation is corrected in the next month.

⁴ The results are of the unit root tests are based on standard ADF and ERS tests. Only the price pairs (that is, domestic and world) that exhibited the same order of integration $I(1)$ are included for cointegration analysis. The other price pairs that exhibited differing orders of integration are excluded from this study.

Table 4: Cointegration Results of Engle–Granger and TAR Model

| | EG | TAR | | | |
|----------------------|----------|--------------|--------------|---------|------------------------------|
| | γ | γ_1 | γ_2 | Φ | $H_0: [\gamma_1 = \gamma_2]$ |
| Wheat (India) | -3.18* | -0.13 (3.06) | -0.06 (1.38) | 5.65 | 1.16 [0.28] |
| Rice (India) | -2.66 | -0.04 (1.29) | -0.12 (2.73) | 4.57 | 1.98[0.16] |
| Sugar (India) | -2.36 | -0.08 (2.94) | -0.02 (0.95) | 4.75 | 2.31[0.13] |
| Tea (India) | -3.61** | -0.21 (3.76) | -0.07 (1.63) | 8.43** | 3.67 [0.05] |
| Rice (Philippines) | -3.34* | -0.06(2.23) | -0.08 (2.57) | 5.72 | 0.27 [0.60] |
| Coffee (Philippines) | -5.37*** | -0.08 (1.28) | -0.46 (6.79) | 23.5*** | 16.22[0.00] |
| Beef (Thailand) | -3.45** | -0.13 (3.69) | -0.04 (1.48) | 7.84** | 3.59 [0.06] |
| Rubber (Malaysia) | -6.10*** | -0.23 (5.74) | -0.13 (2.56) | 19.8*** | 2.24 [0.13] |
| Sugar (Indonesia) | -0.59 | -0.05 (0.53) | -0.10 (1.27) | 0.95 | 1.56 [0.22] |
| Rice (Indonesia) | -3.51** | -0.56 (4.38) | -0.30 (1.67) | 10.8*** | 1.47[0.23] |
| Rice (PRC) | -2.80 | -0.01 (0.26) | -0.25 (3.79) | 7.21** | 6.83 [0.01] |
| Sugar (PRC) | -1.46 | -0.09 (1.85) | -0.02 (0.41) | 1.79 | 1.35 [0.25] |
| Soyabean Oil (PRC) | -2.79 | -0.18 (2.89) | -0.06 (1.11) | 4.82 | 1.76 [0.18] |

EG = Engle–Granger test, TAR = threshold auto-regression.

***, ** and * denote significance at the 1%, 5%, and 10% levels, respectively.

Source: Author's calculations.

Table 5 reports the results of the M-TAR model. The second column contains the Engle–Granger tests results again to facilitate comparison. Columns 2 and 3 of Table 5 provide the estimates of the autoregressive decay in the M-TAR model. The parameter estimates are of the correct signs showing that the error term is stationary and therefore the model is well specified. Column 4 reports the Φ^* - statistic, which tests the null hypothesis, that is, $H_0[\gamma_1 = \gamma_2 = 0]$ of no cointegration. The null hypothesis of no cointegration under the M-TAR model specification can be rejected for nine out of the 13 commodities. Out of the nine commodities, the no cointegration hypothesis can be rejected for the same seven commodities under the Engle–Granger specification and an extra commodity under the TAR specification. A further commodity (soybean oil, PRC) is found to be cointegrated at the 10% significance level. Overall the M-TAR specification shows further support for cointegration. Given that all nine commodities reject the null hypothesis of asymmetry, that is, $H_0[\gamma_1 = \gamma_2]$, at the 10% significance level, we can conclude the adjustment is asymmetric for these commodities and the M-TAR specification as a result will have superior power properties than the standard linear cointegration case (Enders 2001). For four commodities (tea, India; coffee, Philippines; beef, Thailand; and rice, the PRC) the results show that both forms (TAR and M-TAR) of asymmetry may exist. Conducting a model selection test, that is the Schwartz Bayesian criterion, we conclude that for three commodities (tea, India; coffee, Philippines; and beef, Thailand), the M-TAR model is selected. For rice (the PRC), the TAR is found to have a better fit. Under the M-TAR specification, we find that for five commodities (wheat, India; rice, Philippines; coffee, Philippines; rubber, Malaysia; and rice, Indonesia), $|\gamma_1| < |\gamma_2|$, implying that when both international and domestic prices are increasing, the disequilibrium between the prices are corrected at a slower rate relative to when both prices are decreasing. The opposite

form of adjustment behavior takes place when we consider tea (India), beef (Thailand), and soybean oil (the PRC).

Table 5: Cointegration Results of Engle–Granger and Momentum-TAR Model

| | EG | M-TAR | | | |
|----------------------|----------|--------------|--------------|---------|------------------------------|
| | γ | γ_1 | γ_2 | Φ | $H_0: [\gamma_1 = \gamma_2]$ |
| Wheat (India) | -3.18* | -0.04 (0.74) | -0.15 (4.26) | 9.39*** | 8.25 [0.00] |
| Rice (India) | -2.66 | -0.12 (2.99) | -0.03 (1.00) | 4.97 | 2.77[0.09] |
| Sugar (India) | -2.36 | -0.03 (1.30) | -0.13 (2.80) | 4.78 | 2.77[0.09] |
| Tea (India) | -3.61** | -0.22 (3.78) | -0.07 (1.71) | 8.63** | 4.04 [0.04] |
| Rice (Philippines) | -3.34* | -0.05 (2.23) | -0.14 (3.16) | 7.42** | 3.52 [0.06] |
| Coffee (Philippines) | -5.37*** | -0.10 (1.62) | -0.48 (3.16) | 23.1*** | 15.63[0.00] |
| Beef (Thailand) | -3.45** | -0.13 (3.35) | -0.05 (1.92) | 7.36** | 2.68 [0.10] |
| Rubber (Malaysia) | -6.10*** | -0.16 (4.75) | -0.36 (4.48) | 21.3*** | 5.02 [0.02] |
| Sugar (Indonesia) | -0.59 | -0.08 (0.92) | -0.14 (1.74) | 1.95 | 3.53 [0.06] |
| Rice (Indonesia) | -3.51** | -0.28 (1.83) | -0.62 (4.59) | 11.8*** | 2.76 [0.10] |
| Rice (PRC) | -2.80 | -0.06 (1.12) | -0.27 (3.24) | 5.95* | 4.48 [0.03] |
| Sugar (PRC) | -1.46 | -0.22 (2.66) | -0.02 (0.55) | 3.70 | 5.18 [0.02] |
| Soyabean Oil (PRC) | -2.79 | -0.23 (3.35) | -0.05 (1.01) | 6.13* | 4.21 [0.04] |

EG = Engle–Granger test, M-TAR = momentum-threshold autoregression.

***, ** and * denote significance at the 1%, 5%, and 10% levels, respectively.

Source: Author's calculations.

Finally, diagnostics tests were carried out to check for serial correlation. The Ljung Box Q statistic was calculated and the null hypothesis of no serial correlation could not be rejected.

Having found that cointegration exists for nine out of the 13 commodities considered; and that each cointegrating regression is characterized by asymmetric adjustment, we proceed to estimate a threshold ECM. The results of the ECM are given in Table 6.

In the case of India, wheat and tea are the two commodities that display a long-run relationship between international and domestic prices. The price transmission elasticity of this estimated long-run relationship is found to be 0.11 for wheat and 0.54 for tea, both being statistically significant at the 1% level. This implies that for wheat, 11% of a change in world prices will be transmitted to domestic wheat prices in India. For tea, the price transmission is higher, given that we find 54% of any change in world prices will be transmitted to domestic tea prices in India. For wheat, there seems to be no adjustment to possible deviations when prices are increasing, with correction only taking place when possible deviations arise when prices are falling. In the case of tea, when prices are rising, deviations from equilibrium are corrected at a faster rate as opposed to when prices are falling. When prices are rising, 13% of the deviation is corrected every month, whereas when prices are falling, only 7% of the deviation is corrected every month. This implies that the rate of adjustment shows directional asymmetry. For example, in the case

of tea, using the method of calculating half-life shocks, one can estimate that the amount of time it takes to correct 50% of any deviation that arises when prices are increasing is approximately 5 months. When prices are decreasing, the amount of time taken to correct 50% of any deviation would be approximately 10 months. In the case of wheat, deviations are corrected only when prices are decreasing. The estimated time to correct half of any deviation is calculated to be approximately 4.26 months. For both commodities (tea and wheat), the short-run elasticity is found to be insignificant. Therefore, any changes to world prices will have no immediate impact on possible changes to domestic prices of tea and wheat in India.

Table 6: Results of the Threshold Error Correction Model

| | CI | ECT(+) | ECT(-) | Short-Run Elasticity | Long-Run Elasticity |
|----------------------|-----|--------------|--------------|----------------------|---------------------|
| Wheat (India) | Yes | 0.002 (0.03) | -0.15 (4.43) | 0.004 | 0.11*** |
| Rice (India) | No | | | | |
| Sugar (India) | No | | | | |
| Tea (India) | Yes | -0.13 (3.00) | -0.07 (2.13) | 0.04 | 0.54*** |
| Rice (Philippines) | Yes | -0.07 (1.74) | -0.05 (2.88) | 0.02 | 0.23*** |
| Coffee (Philippines) | Yes | -0.16 (2.81) | -0.64 (9.50) | -0.18* | 0.78*** |
| Beef (Thailand) | Yes | -0.02 (1.91) | -0.01 (0.49) | 0.01 | 0.42*** |
| Rubber (Malaysia) | Yes | -0.13 (3.98) | -0.21 (2.86) | 0.22*** | 0.89*** |
| Sugar (Indonesia) | No | | | | |
| Rice (Indonesia) | Yes | -0.36 (2.14) | -0.53 (3.39) | -0.04 | -0.07*** |
| Rice (PRC) | Yes | -0.07 (2.15) | -0.01 (0.16) | -0.01 | 0.46*** |
| Sugar (PRC) | No | | | | |
| Soybean Oil (PRC) | Yes | -0.24 (4.44) | -0.01 (0.35) | 0.11*** | 0.51*** |

***, ** and * denote significance at the 1%, 5%, and 10% levels, respectively.

CI = cointegrated, ECT(+) = error correction coefficient for a positive deviation, ECT(-) = error correction term for a negative deviation.

Source: Author's calculations.

Moving on to the Philippines, rice and coffee display a long-run relationship between international and domestic prices. The price transmission elasticity of the estimated long-run relationship is found to be 0.23 for rice and 0.78 for coffee, both being statistically significant at the 1% level. This implies that for rice, 23% of a change in world prices will be transmitted to domestic rice prices in the Philippines, whereas for coffee, 54% of any change in world prices will be transmitted to domestic coffee prices in the Philippines. In the case of rice, when prices are rising, deviations from equilibrium are corrected at a faster rate, albeit very small, as opposed to when prices are falling and when deviations appear between the prices. In the case of rice when prices are increasing and a deviation is created in the long-run equilibrium between the two prices, half of the deviation is corrected in approximately 10 months. However, in the opposite scenario, when prices are falling, half of the deviation is corrected in 13.5 months. The short-run elasticity is found to be insignificant. Therefore, any changes to world prices will have no immediate

impact on possible changes to domestic prices. For coffee, the results are different in terms of asymmetric adjustment. In contrast to rice, when prices are rising, deviations from equilibrium are corrected at a relatively slower rate, as opposed to when prices are falling. However, the magnitude of the speed of adjustment coefficients are quite high. For instance, when prices are increasing and a deviation appears between domestic and international prices, half of the deviation is corrected in less than 3 months. On the other hand when prices are increasing and a disequilibrium is created, half the deviation is estimated to be corrected in less than a month; approximately 3 weeks. The short-run elasticity is found to be significant; however, surprisingly negative. This implies that an increase in world prices of coffee will have an immediate impact on domestic prices of coffee, causing prices to fall.

In the case of beef from Thailand, the price transmission elasticity is estimated to be 0.42, implying that 42% of any variation in world prices is transmitted to domestic beef prices in Thailand. The short-run elasticity is insignificant suggesting that there is no immediate impact of a change in world price on domestic beef prices. The adjustment to any possible deviation takes place only when prices are increasing. This adjustment is found to be very slow; it takes over 34 months for only 50% of the deviation to be corrected.

When considering rubber from Malaysia, we find that both long-run and short-run price transmission elasticity is positive and significant. In this case, the long-run price transmission is quite high as 89% of any deviation in world prices is transmitted to domestic prices. The short-run elasticity demonstrates a significant immediate impact of a change in world prices on domestic rubber prices. Approximately 5 months are required for any deviation to be corrected when prices are rising, whereas when prices are falling 21% of the deviation is corrected every month, implying that it takes less than 3 months to correct half of any deviation.

Moving on to Indonesia, in the case of rice we find that a long-run relation exists between domestic and international prices. However, a surprising result is that the long-run price transmission elasticity is negative, albeit the magnitude of the elasticity coefficient is very small. The adjustment mechanism shows that when prices are rising, deviations from equilibrium are corrected at a relatively slower rate, that is, 36% of the deviation every month. In contrast, when prices are falling, deviations are corrected at a relatively faster rate, that is 53% of the deviation every month. The short-run price elasticity is found to be insignificant.

In the case of the PRC, a long-run relationship is found to exist between the domestic and international prices of two commodities, rice and soybean oil. For both commodities, there seems to be no adjustment to possible deviations when prices are decreasing with correction only taking place when prices are increasing. In the case of rice the deviations are corrected at the rate of 7% every month when prices are increasing, whereas for

soybean oil, the deviations are corrected at the rate of 24% every month. For both commodities the speed of adjustment is insignificant when prices are falling, implying that no adjustment takes place when deviations appear in this scenario. The long-run price transmission elasticity for both commodities is significant showing that approximately half of the variations in world prices are transmitted to domestic prices. For rice there is no immediate impact of world price changes on domestic prices, however, for soybean oil we find that the short-run elasticity is positive and significant. Therefore, any changes to world prices of soybean oil will have an immediate impact on possible changes to domestic prices in the PRC.

VI. Policy Implications

This section describes the policy implications of the empirical findings of trends of international commodity prices and the price transmission of international to domestic commodities with reference to Asian countries.

A. Policy Implications of Trends in Commodity Prices

When considering the underlying trends of primary commodity prices, we find that all prices are characterized by changing trends that can be classified into different regimes over time. What do these results mean in terms of policy? The underlying price movements, whether trend-stationary or difference-stationary, may affect the income levels of developing countries reliant on primary commodities. For example, stabilization policies can aid in smoothing income flows. However, conducting appropriate econometric analysis on the price series to determine whether the price series is trend- or difference-stationary would have an impact on these policies. For example, when the prices series is trend-stationary, it has been argued that stabilization policies are effective. However, Reinhart and Wickham (1994) warn that the stabilization policies may be difficult to implement if the price series have a varying trend. The evidence from this study indicates that such policies would be either ineffective or difficult to implement. This conclusion is broadly similar to the findings of Kellard and Wohar (2006) and Ghoshray (2010) where the analysis was conducted on lower frequency data over a substantially longer period of time.

The results of this study show that for two key commodities such as coffee and sugar we find evidence of difference-stationary behavior. These two commodities have abandoned price stabilization policies. Regulated exports were abandoned in 1989 for coffee, and price stabilization measures were removed in 1992 for sugar. When countries experience a period of negative trend in primary commodity prices, the setback can be partially offset by securing funds from the International Monetary Fund under the Compensatory

Financing Facility scheme, or from the European Union under the Stabilization System for Export Earnings scheme. These policies should be based on the underlying prevalence of the trend, and evidence from this study suggests that the design and form of assistance would be difficult to implement given the mixed and varying trend results (Ghoshray 2010).

We find no evidence of a secular long-run trend over the entire sample considered in this study. Rather, the results show that some commodities experience segments of a negative trend or positive trend interspersed by periods of approximate stability. The structural breaks that are determined by the test occur on different dates. Since the structural breaks are unpredictable, forecasting of commodity prices can prove to be difficult (Kellard and Wohar 2006).

It has been argued that countries that face a declining trend in the prices of their primary commodities may lead to a current account deficit. This deficit can be compensated by allowing for a capital account surplus driven by increased public and private borrowing. In the event where a country experiences both current and capital account deficits, one may observe a decline in international monetary reserves and currency instability. Given that for several commodities (palm oil, coconut oil, tea, and beef) we find no evidence of a trend in recent years, one may infer from these results that foreign exchange constraints facing developing countries can be relaxed (Newbold et al. 2005). However, Newbold et al. point out that the risk of revenue shortfall during the period of time when the repayments are anticipated can be quite high.

The experience of some of the countries in Asia, such as Hong Kong, China; the Republic of Korea; Singapore; and Taipei, China, shows that they moved toward successful export diversification. However, other Asian countries have not been as successful. Various policy recommendations have been made regarding the composition of exports. While Singer (1999) recommended that developing countries need to diversify their exports into manufacturing, institutions such as the World Bank encouraged countries to diversify into new exports of primary commodities. This policy of export diversification to other primary commodities would depend to a large extent on the availability of resources and potential export destinations. Pindyck and Rotemberg (1990) opposed this view by providing econometric evidence that many commodity prices tend to move together since they are generally considered as substitutes. This evidence was subsequently challenged by Leybourne et al. (1994) and Deb et al. (1996) who found weak evidence for such relationship. However, Page and Hewitt (2001) argue that for small countries, this policy is an unlikely substitute for industrialization.

B. Policy Implications of International Price Transmission

What do the results relating to world price transmission imply in terms of policy? The recent boom and subsequent decrease in agricultural commodity prices have thrown up

serious questions about the impact of such changes on the domestic prices of developing countries. Governments of developing countries need to have a good knowledge of the functioning of their domestic markets. This knowledge includes how changes in world prices are transmitted to domestic prices.

The above analysis is based on how international commodity prices are transmitted to domestic prices through means of a cointegration and ECM. In addition, to examine whether there are significant differences in price transmission as a result of an increase or decrease in world prices, a TAR/M-TAR model is adopted to facilitate this comparison. The data consists of 13 possible commodity/country pairs. Based on the Enders and Siklos (2001) test, nine out of the 13 commodity/country pairs show evidence of a long-run relationship with asymmetric adjustment. This result proves a significant case in point; that if asymmetry were to exist, the power of these TAR/M-TAR tests are greater than the standard linear test. Accordingly we find only seven out of the 13 pairs cointegrated when using the linear model. However, a further two pairs are found to cointegrate when adopting the nonlinear threshold model. When observing the long run price transmission elasticity, we find that of the nine pairs that show a long-run relationship, all have a long-run price transmission elasticity that is statistically significant, one of which surprisingly is found to be negative. This result however, needs to be treated with caution. For this particular case, the data span was too small to be rendered as forming a “long-run relationship.” Considering this negative elasticity as an outlier, for the remaining eight elasticities, we find that the values range from 0.11 to 0.89, with a median value of 0.51. The median value implies that 51% of the variation in world prices is transmitted to domestic prices.

The high values of the price transmission elasticity denote closer connection between world and domestic prices, indicating an integrated market for the commodity in question. The short-run price transmission elasticity for coffee in Philippines and rice in the PRC and India is found to be negative. An explanation for this result may be that applied tariffs were raised to limit falling world prices. When these policies (tariffs) are intense then it can lead to the short-run price transmission elasticity being negative.

When considering an important commodity for Asian countries such as rice, we find that except for India, all the other domestic prices (the PRC, Indonesia, the Philippines) cointegrate with world prices. For the PRC and the Philippines, we find a similar mechanism of price adjustment. For both countries, when prices are increasing, the gap between the international and domestic prices is corrected at a faster rate than when prices are decreasing. In the case of the PRC, however, the correction is statistically insignificant suggesting that when world prices are decreasing they are not transmitted to domestic prices. The long-run price transmission elasticity for the PRC is significantly higher than that of the Philippines and the short-run elasticity is insignificant for both countries.

Existence of a cointegrating relation implies a high degree of similarity between world and domestic prices. The absence of cointegration will imply that domestic producers are completely insulated from long-term price developments in the world market. This interpretation does not mean that prices are unrelated, but that the “tracking” of prices implied by cointegration is unobserved (Lloyd et al. 1999). As reported in this study, a significant degree of transmission is observed at the short-run level.

What does the existence of price transmission convey? Price transmission relays unbiased information on prices to agricultural producers leading to efficient allocation of resources. Incomplete price transmission on the other hand creates biased incentives to agricultural producers, which in turn leads to suboptimal or reduced productivity. Also no price transmission implies reforms do not reach agricultural producers. For example, some Asian countries, such as India and the Philippines, used to insulate their domestic economy from world price increases by implementing various commodity-based policies. These policies include government storage, procurement, and distribution as well as restrictions on international trade (Dawe 2008b).

Where price transmission is found to exist, we conclude that world price variation is transmitted to domestic markets. For example, world price changes in rice are found to be transmitted to the domestic market of the PRC. Though the PRC does not allow the private sector to trade at all, in recent years it had allowed changes in international prices to be reflected in domestic prices.

Indonesia historically has stabilized domestic prices (Timmer 1986). But domestic prices have been more volatile than the international prices in recent years. Domestic prices have surged in the recent past as rice imports were restricted in an attempt to boost farm incomes (Dawe 2008b). This happened at times even when world prices were relatively stable. Though Indonesia is classified as a country that employs price stabilization, the results show that rice prices have been insulated from the world market. Indonesia has a history of market intervention aimed at stabilizing domestic market prices with the help of the state trading agency BULOG. The parastatal BULOG continued until 1998 after which its operations were scaled down. These included abolishing import monopoly of rice and wheat and ending the domestic pricing and distribution of wheat and flour. According to Sharma (2002), given the scale of operations of BULOG throughout the 1990s, the weak relation observed between world and domestic prices may not be surprising after all.

India and the Philippines have more or less successfully stabilized their prices. However, world price variation is found to be transmitted to domestic prices. In recent years, especially during the food price spike, the changes in world prices may have led to price increases to levels higher than expected (Dawe 2008b). In the case of India we find transmission of world prices to domestic prices taking place for wheat but not rice. This could be explained by the fact that in India it was only in 1994 that the economic liberalization program that began in 1991 was extended to agriculture (Sharma 2002).

Prior to the reforms taking place, quantitative controls in trade were in place for a long time, which happened to be a prominent feature of trade policy at the time. However, after 1994 exports of cereals took place at regular intervals but were often regulated with quotas and minimum export prices. One of the main reasons for implementing these measures was the domestic price situation (Sharma 2002). The Indian government also initiated in 1993 open market operations, that is, sales of public stocks aimed at stabilizing domestic prices. Such interventions were carried out from time to time since then, which also weakened the relation between world and domestic prices.

What does asymmetric price transmission convey? Asymmetry can worsen income inequalities inside the country. This is because part of the benefit of an increase in world prices may be siphoned off by agents higher up the marketing chain, and a small residue of the benefits may accrue to the consumers/ producers, leading to inequality.

What are the causes of asymmetric responses of domestic prices to international commodity prices? Morisset (1998) provides possible explanations for this type of dynamic behavior between prices. In the presence of constraints on the quantity of the commodity being exported, the decline in world commodity prices will not be transmitted to domestic prices because there is no incentive for exporters to make their exports more attractive by lowering their export prices. An increase in the price spread between world and domestic prices can be observed, which can be caused by import barriers that exporters face. Examples for this could be levies and variable tariffs.

Foster and Baldwin (1986) provide an explanation that can help explain the asymmetric transmission. When world prices are falling they will be transmitted imperfectly to domestic prices, because if existing sales are constrained by marketing capacity, exporters will compensate for rising marketing costs by increasing their selling prices. This increase will partially offset the initial impact of declining world prices on domestic prices. Because there are no similar constraints when world prices are rising, domestic prices are expected to respond more to rising rather than falling world prices. However, statistics suggest that over time transportation and insurance costs have been decreasing. While the international evidence on marketing costs is limited, the costs have shown a declining trend in the case of the US. These observations seem to suggest that transactions costs may have caused asymmetric adjustment in domestic prices.

The asymmetry in price transmission is evident from the results. In the case of tea from India, rice from the PRC and the Philippines, beef from Thailand, and soybean oil from the PRC, decreases in world prices are incompletely or slowly passed through to the domestic market as compared to increases. This asymmetry may be due to the floor price level policy resulting in a smoothing of the downward changes in world prices. Another reason may be the market power exerted by the distribution levels of the supply chain. The implication is that if a fall in world prices is not fully transmitted to domestic prices, then the expected increase in world demand and decrease in world supply, which would

have otherwise occurred, will not take place. On the other hand, in the case of coffee from the Philippines, rubber from Malaysia, and rice from Indonesia, domestic prices adjusted more rapidly to world prices when world prices were declining. This may suggest that exporters chose to increase their market shares rather than their mark-ups.

Agricultural policies may or may not hinder market integration, depending on the nature of the policy instruments employed. For example, the rice and sugar market of India was found not to be integrated with the international market. A possible explanation may be that for a long period, tariffs for agricultural commodities were characterized by a high degree of dispersion that translated to domestic prices for many commodities being lower than the duty-inclusive import prices. This may have caused the lack of integration between world and domestic rice and sugar prices. On the other hand, minimum price support policies implemented in the Indian wheat market were found not to impede market integration, but to result in relatively slow and asymmetric adjustment to international price changes. In general, for markets that are subject to policies, the speed of adjustment, as reflected by the error correction coefficients, was estimated to be relatively low. Although several authors stress that policies impede the extent of price transmission (see for example, Mundlak and Larson 1992, Quiroz and Soto 1993, Baffes and Ajwad 2001, Abdulai 2000, Sharma 2002), it should be noted that other reasons such as high transaction costs and other distortions may also be the cause of slow adjustment.

VII. Conclusion

Nonlinearities and asymmetric adjustment remain an important issue to be considered especially when the objective of the research is to provide a mechanism that can be incorporated in a structural partial equilibrium model. Although asymmetric adjustment may also be the outcome of market imperfections, it is plausible that price support policies result in positive and negative changes in the international price affecting the domestic market in different ways. More importantly, such policies may imply a “threshold” or minimum price, above which transmission of price signals take place. Such a discrete adjustment process implies that movements toward long-run equilibrium do not take place at all points in time, but only when the divergence from equilibrium exceeds a certain threshold. For example, policies such as price support mechanisms and tariff rate quotas may result in such an adjustment process. In the former case, governments may intervene in the market when market prices fall below a floor level, while in the latter, international price signals pass-through when import volumes are sufficiently within, or out of the quota.

Thus, future research should continue to focus on threshold cointegration, which may be beneficial, as it provides additional information in the form of the threshold, if the objective of the analysis is the development of price transmission mechanisms. Apart

from assessing the effect of food and trade policies on market integration, threshold cointegration applied in commodity markets of developing countries may also provide a rough indication of transfer costs. Transfer costs in developing country markets may give rise to a threshold over which arbitrage possibilities are obliterated, resulting in an absence of market integration. A threshold cointegration framework can encompass the possibility of non stationary transfer costs and provide valuable information that can lead to policy prescriptions.

In light of these findings, an obvious question is how can Asian countries protect themselves from variations in world commodity prices? Minot (2010) argues the simplest answer is food self-sufficiency, which can be achieved by investment in agricultural research, extension, disease control, and reduction in postharvest losses. However, Minot goes on to argue that this would be a long-term strategy that is not so appealing in the political arena. The likelihood of success varies according to whether the country in question is more or less self-sufficient in one or more commodities. Another approach put forward by Minot is to restrict imports through tariffs or quotas. This would result in self-sufficiency with no additional cost to the government. However, the downside is that it would significantly raise the price of staple agricultural commodities. Since staple commodities such as rice and wheat continue to be imported by various Asian countries, the price at which self-sufficiency is achieved still needs to be higher. To avoid the effects of a food price spike such as the one experienced in 2007–2008, this may require that the staple commodity prices remain high, which would have serious effects on food security especially among the urban poor.

Appendix

Trends/Unit Root Tests with Structural Breaks

To briefly describe the Lee and Strazicich (2003) method, consider the following data generating process (DGP):

$$P_t = \psi' X_t + u_t \text{ and } u_t = \phi u_{t-1} + \varepsilon_t; \text{ where } \varepsilon_t \sim \text{iid } N(0, \sigma^2) \quad (\text{A.1})$$

where P_t is the price series and the two changes in level and trend are given by

$$X_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]',$$

where

$$DT_{jt} = \begin{cases} t - TB_j & \text{for } t \geq TB_j + 1 \\ 0 & \text{otherwise} \end{cases} \quad \text{for } j = 1, 2.$$

TB_j denotes the points at which the breaks occur. Note that the DGP contains breaks in the null hypothesis when $H_0 : (\phi = 1)$ and the alternative hypothesis when $H_A : (\phi < 1)$. The break fractions are denoted as $\lambda_j = TB_j / T$ where T denotes the total number of observations.

When employing the Lee and Strazicich (2004) method, which considers a single structural break, the single change in level and trend in (A.1) is now given by $X_t = [1, t, D_t, DT_t]'$,

where

$$DT_t = \begin{cases} t - TB & \text{for } t \geq TB + 1 \\ 0 & \text{otherwise} \end{cases}$$

TB denotes the points at which the breaks occur. The break fraction is denoted as $\lambda = TB/T$.

The LM unit root test statistic can be estimated by the following regression:

$$\Delta P_t = \phi' \Delta X_t + \gamma \bar{Y}_{t-1} + \sum_{i=1}^p \psi_i \Delta \bar{Y}_{t-i} + u_t \quad (\text{A.2})$$

where $\bar{Y}_t = P_t - \mu - X_t \bar{\phi}$, $t = 2, 3, \dots, T$; $\bar{\phi}$ are coefficients on the regression of ΔP_t on ΔX_t ; μ is given by $P_1 - X_1 \bar{\phi}$.⁵ The lagged terms $\Delta \bar{Y}_{t-i}$ are added to correct for serial correlation. The augmentation is determined using the general to specific method. The LM test statistics are given by the τ statistic testing the null hypothesis $H_0 : (\gamma = 0)$. The LM unit root test determines the break points endogenously by utilizing a grid search. To eliminate endpoints trimming of the *infimum* (inf) is made at 10%. The breakpoints are determined where the test statistic is minimized.

The LM test is given as $LM_\tau = \inf \hat{\tau}(\lambda)$.

⁵ where P_1 and X_1 denote the first observations of the P_t and X_t sequences, respectively.

If we find that we can identify up to two structural breaks in any price series, in the next step we attempt to determine whether the sign of the trend is negative or positive and whether it is significant or not.

For the trend-stationary process the estimation process is carried out using the following equations:

$$P_t = \gamma + \delta_1 t + \delta_2 D_{L1,t} + \delta_3 D_{T1,t} + \delta_4 D_{L2,t} + \delta_5 D_{T2,t} + u_t \quad (\text{A.3})$$

$$u_t - \phi_1 u_{t-1} - \dots - \phi_p u_{t-p} = \varepsilon_t - \psi_1 \varepsilon_{t-1} - \dots - \psi_q \varepsilon_{t-q} \quad (\text{A.4})$$

where ε_t is a white noise process. D_{Li} and D_{Ti} ($i = 1, 2, 3$) denote the level and slope dummy, respectively; i refers to the regime defined by the prior identification of the break dates. The regimes are defined as:

regime 1 = start date to $TB1$

regime 2 = $(TB1+1)$ to $TB2$

regime 3 = $(TB2+1)$ to end date

Following Kellard and Wohar (2006), equation (A.3) is reparameterized to facilitate the estimation of the trend coefficient in the three different regimes. Equation (A.3) was reparameterized in the following way:

$$\begin{aligned} P_t^* = & \gamma R_{L1,t} + \delta_1 R_{T1,t} + (\gamma + \delta_2) R_{L2,t} + (\delta_1 + \delta_3) R_{T2,t} + (\gamma + \delta_2 + \delta_4) R_{L3,t} \\ & + (\delta_1 + \delta_2 + \delta_3) R_{T3,t} + u_t \end{aligned} \quad (\text{A.5})$$

where $R_{Li,t}$ denotes the intercept dummy for a level shift in regime i , ($i = 1, 2, 3$) and $R_{Ti,t}$ denotes the slope dummy for a trend shift in regime i , ($i = 1, 2, 3$). For the three regime model, P_t^* is defined as:

$$P_t^* = \begin{cases} P_t & \text{if } 1 \leq t \leq TB1 \\ P_t - \delta_1 (TB1 - 1899) & \text{if } TB1 + 1 \leq t \leq TB2 \\ P_t - \delta_1 (TB1 - 1899) - \delta_3 (TB2 - TB1) & \text{if } t > TB2 \end{cases}$$

The estimation was conducted by exact maximum likelihood and the ARMA order (p, q) was selected through the Schwartz Bayesian Criteria (SBC) allowing all possible models with $p + q \leq 6$.

TAR and M-TAR Models

Consider the equation below:

$$P_{1t} = \alpha + \beta P_{2t} + \varepsilon_t \quad (\text{A.6})$$

where P_{1t} and P_{2t} are nonstationary I(1) prices, α and β are the estimated parameters. To test the hypothesis (A.6) is estimated using OLS. The second step advocates a Dickey-Fuller test on the estimated residuals of equation (A.6) of the following kind:

$$\Delta \varepsilon_t = \gamma \varepsilon_{t-1} + \omega_t \quad (\text{A.7})$$

where ω_t is a white noise error term.

Enders and Siklos (2001) propose the threshold autoregressive (TAR) model, such that equation (A.7) can be written as:

$$\Delta \varepsilon_t = I_t \gamma_1 \varepsilon_{t-1} + (1 - I_t) \gamma_2 \varepsilon_{t-1} + \omega_t \quad (\text{A.8})$$

where I_t is the Heaviside indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } \varepsilon_{t-1} \geq \tau \\ 0 & \text{if } \varepsilon_{t-1} < \tau \end{cases} \quad (\text{A.9})$$

This specification allows for asymmetric adjustment.

In the above case it is necessary to estimate the value of the threshold that will be equal to the cointegrating vector. A method of searching for a super-consistent estimate of the threshold was undertaken by using a method proposed by Chan (1993).

To utilize Chan's methodology, the estimated residual series was sorted in ascending order, that is, $\varepsilon_1 < \varepsilon_2 < \dots < \varepsilon_T$ where T denotes the number of usable observations. According to the method, the largest and smallest 15% of the ε_i series were eliminated and each of the remaining 70% of the values were considered as possible thresholds. For each of the possible thresholds the equation was estimated using equation (A.8) and equation (A.9). The estimated threshold yielding the lowest residual sum of squares was deemed to be the appropriate estimate of the threshold.

In equation (A.9) the Heaviside Indicator depends on the level of ε_t . Enders and Siklos (2001) suggest an alternative Momentum-Heaviside Indicator as:

$$I_t = \begin{cases} 1 & \text{if } \Delta \varepsilon_{t-1} \geq \tau \\ 0 & \text{if } \Delta \varepsilon_{t-1} < \tau \end{cases} \quad (\text{A.10})$$

In this case the series ε_t exhibits more momentum in one direction than the other.

To estimate the threshold, Chan's methodology is employed. The estimated residual series was sorted in ascending order, that is, $\Delta \varepsilon_1 < \Delta \varepsilon_2 < \dots < \Delta \varepsilon_T$ where T denotes the number of usable observations. As before with the TAR model, the largest and smallest 15% of the ε_i series were eliminated and each of the remaining 70% of the values were considered as possible thresholds. For each of the possible thresholds the equation was estimated using equation (A.8) and equation (A.10). The estimated threshold yielding the lowest residual sum of squares was deemed to be the appropriate estimate of the threshold.

To implement in this test the case of the TAR or M-TAR adjustment, the Heaviside Indicator function is set according to equation (A.9) or equation (A.10), respectively, after which equation (A.8) is estimated accordingly.

The Φ -statistic for the null hypothesis of stationarity of ε_t , i.e., $H_0 : (\gamma_1 = \gamma_2 = 0)$ is recorded. The value of the Φ -statistic is compared to the critical values computed by Enders and Siklos (2001).

If we can reject the null hypothesis, it is possible to test for asymmetric adjustment since γ_1 and γ_2 converge to multivariate normal distributions (Tong 1990).

The F statistic is used to test for the null hypothesis of symmetric adjustment, that is,

$$H_0 : (\gamma_1 = \gamma_2) .$$

Diagnostic checking of the residuals are undertaken to ascertain whether the ω_t series is a white noise process using the Ljung-Box Q tests. If the residuals are correlated, equation (A.8) needs to be re-estimated in the form:

$$\Delta \varepsilon_t = I_t \gamma_1 \varepsilon_{t-1} + (1 - I_t) \gamma_2 \varepsilon_{t-1} + \sum_{i=1}^p \phi_i \Delta \varepsilon_{t-1} + \omega_t \quad (\text{A.11})$$

Equation (A.11) is comparable to equation (A.8) except that it incorporates lagged first differences of the dependent variable to correct for the autocorrelation in the error term ω_t . The Schwartz Bayesian Criteria is used to determine the lag length.

Data Description and Sources

World prices of the agricultural commodities chosen in this study are drawn from the *International Financial Statistics*. The commodity prices are measured in US\$/ton deflated by the US consumer price index (CPI). The span of the data is from January 1960 to February 2010. The subsequent analysis of the data is carried out on prices transformed to natural logarithms.

Domestic prices are collected from various sources. The commodity prices are deflated by the domestic CPI. The subsequent analysis of the data is carried out on prices transformed to natural logarithms (see Appendix Table below).

| Country | Commodity | Data Span | Price Currency | Source |
|----------------------------|--|--------------------------------|----------------------------|--|
| Thailand | Beef, coconut oil, soybean oil, palm oil | January 1993 to February 2010 | Thai baht/kilogram | Ministry of Commerce |
| | Rubber | January 1993 to February 2010 | Thai baht/kilogram | Department of Internal Trade |
| India | Rice, sugar, wheat | January 1994 to February 2010 | Indian rupee/quintal | Department of Consumer Affairs |
| Philippines | Rice | January 2004 to February 2010 | Philippine pesos/kilogram | Bureau of Agricultural Statistics |
| | Coffee, corn, rubber | January 1990 to February 2010 | Philippine pesos/kilogram | CountrySTAT, Bureau of Agricultural Statistics |
| People's Republic of China | Rice, sugar | January 2001 to February 2010 | Renminbi/500 grams | Price Monitoring Centre |
| | Beef | January 2004 to February 2010 | Renminbi/kilogram | Ministry of Commerce |
| | Rubber | January 2001 to February 2010 | Renminbi/ton | Price Monitoring Centre |
| Indonesia | Rice | November 2006 to February 2010 | Indonesian rupiah/kilogram | National Bureau of Statistics |

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About the Paper

Atanu Ghoshray conducts an empirical analysis of commodity prices. The aim of this paper is twofold: to study the underlying trends for selected international agricultural commodity prices, and to test for international price transmission employing the cointegration method in the presence of asymmetric error correction. Using commodity-specific monthly data, this paper examines the extent to which increases in international food prices during the past few years have been transmitted to domestic prices in selected Asian developing countries.

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