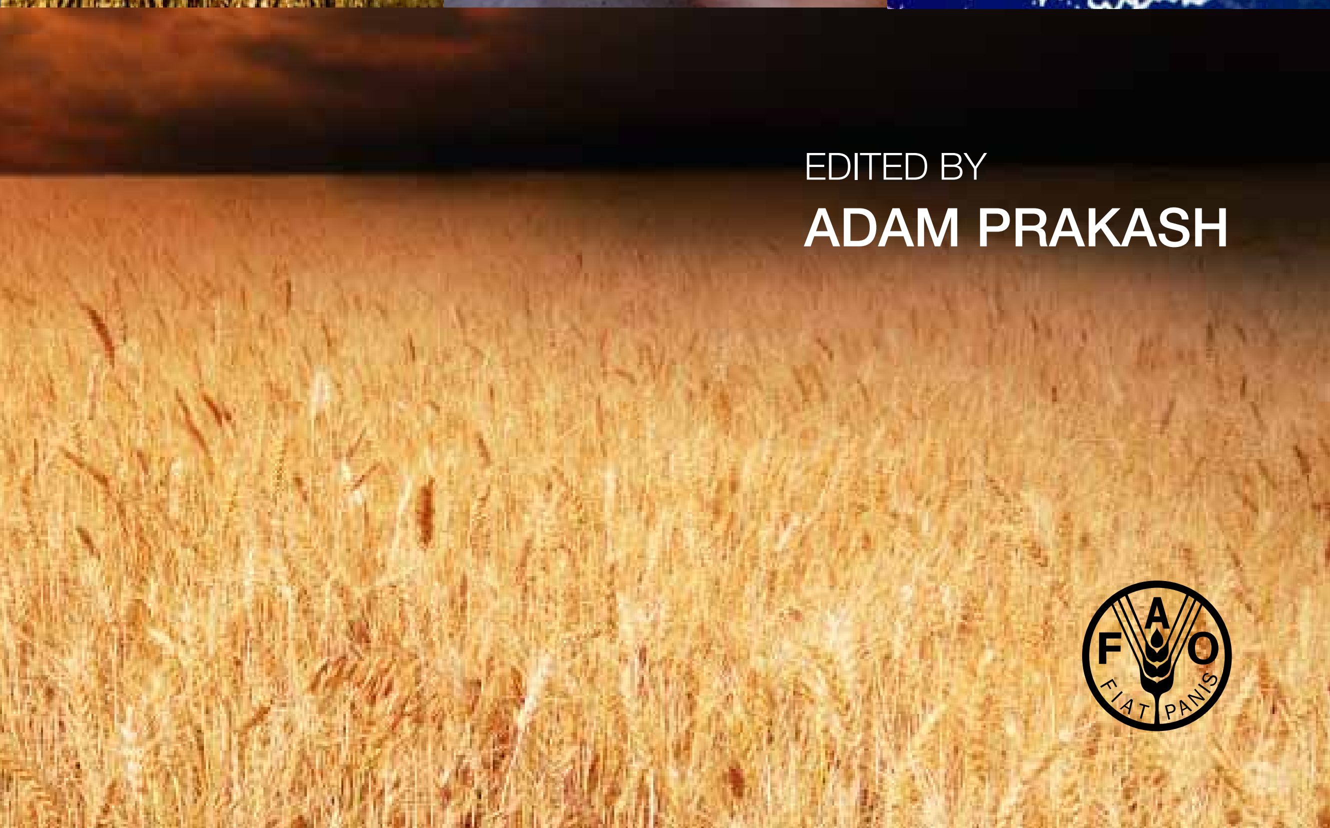
SAFEGUARDING FOOD SECURITY IN VOLATILE GLOBAL MARKETS



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Safeguarding food security in volatile global markets

Edited by Adam Prakash

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Chapter 7

Grains price pass-through, 2005-09[1]

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The world's press and television are replete with statistics on the level of volatility of "world prices" for grains, and political discussion, particularly at the multilateral level, focuses on these prices. In most cases, the so-called world prices are prices associated with the main grains futures markets. However, these are not the prices that consumers pay in any national market and neither are they prices that farmers receive unless they are sufficiently large and well-placed to be able to deliver onto these markets. Instead, consumers pay and farmers receive local prices denominated in their own local currencies and which, to some extent, reflect local market conditions. These local prices will follow the movements in world prices to a greater or lesser extent and with a shorter or longer lag. This defines the topic of price pass-through – to what extent and how rapidly are movements in world grains prices passed through into local prices.

Pass-through is critical in the evaluation of the impact of the large movements in grains prices over the volatile five year period 2006-10, particularly in relation to the plight of consumers in Low-Income Countries (LICs). Movements in world prices are relevant to LICs only to the extent that there is pass-through. Volatility is attenuated if this pass-through is slow. Policies must reflect the extent and speed of pass-through – the focus of multilateral attention should be on those countries where pass-through is high and rapid. In those countries where pass-through is low, local prices will not have been so much affected by rises in world prices; in those countries in which it is slow, there will have been a degree of smoothing.

These issues have already been analysed in the current policy context for maize, rice, soybeans and wheat in Guatemala by de Janvry & Sadoulet (2009). Daviron (2008) looked at transmission of the high 2007-08 wheat and rice prices in six African countries.[3] In both cases, the analysis finishes in mid-2008 so the authors were unable to examine the effects of the September 2008 financial crisis on falling prices.

In this chapter, I look at the pass-through of movements in world prices of the three major grains - maize (corn), wheat and rice - in six developing countries - Benin, Kenya, Malawi, Nepal, Peru and Viet Nam (the first four of which are on the World Bank's list of LICs). The results show that pass-through varies in interesting ways both across the three grains

[1] This chapter derives from section 3 of Gilbert (2010) which was prepared under contract to the FAO. I am grateful to the following collaborators in the FAO project: Rose Edwige Fiamoh (Benin), Harriet Mugera (Kenya), Sridhar Thapa (Nepal), Hien Minh Vu (Viet Nam), Santos Maza Ysillupu (Peru) and Wouter Zant (Malawi). The views expressed are those of the author and not of those of his collaborators or the FAO.

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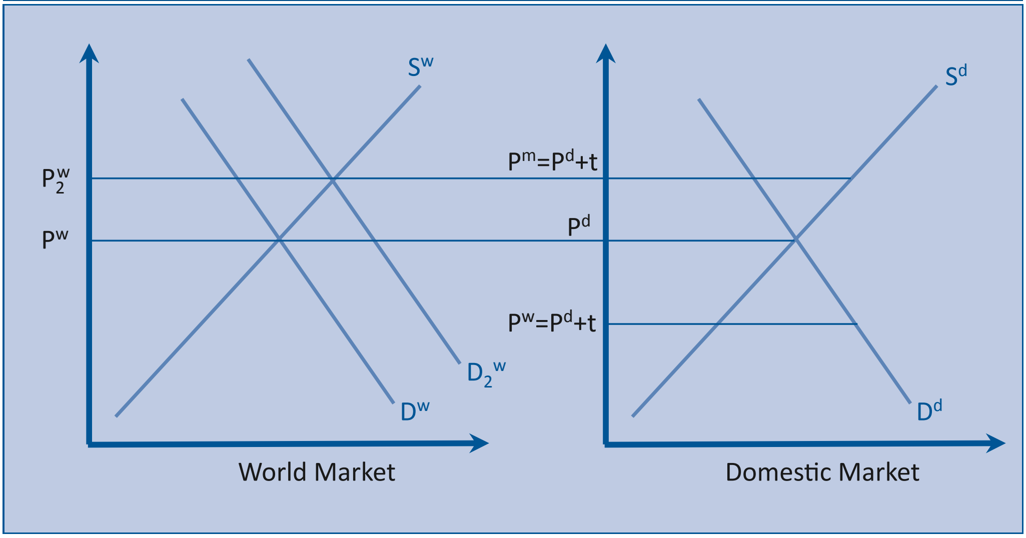
[3] Cameroon, Guinea, Madagascar, Mali, the Niger and Senegal.

Figure 7.1: Import and export parity prices: theory



and across the six countries. The following section looks at the methodologies available to analyse pass-through. These are then applied to maize, wheat and rice, respectively. The last section concludes.

Pass-through analysis

Economists often take full pass-through of world prices to local markets as an ideal as any limitation of pass-through will reduce the sensitivity of both production and consumption to price signals. Maximum pass-through will act to reduce the price response to production and consumption shocks and hence will limit price volatility.

In practice, there are three groups of factors which may limit pass-through:

1. Transport and other costs drive a wedge between prices on world markets and domestic prices. These costs can be particularly high for landlocked countries. Given a world price p^w, we can think of an export parity price $p^x = p^w - c^x$, where c^x is the cost of exporting and an import parity price $p^m = p + c^m$, where c^m is the cost of importing. So long as the domestic supply and demand curves intersect at a price p^d in the range $p^x < p^d < p^m$ domestic prices will be unaffected by world prices (Timmer et al., 1983; Baulch, 1997). This is illustrated in Figure 7.1 where the domestic price (right panel) p^d is initially equal to world price p^w (left panel). For simplicity, export and import costs are both set equal to the transport cost t resulting in a non-variation band of $2t$ around the domestic price. Shifts in the domestic supply and demand curves which keep domestic prices within this band will result in changes in domestic prices which are uncorrelated with world prices. The world price has to rise at least to $p_2^w = p^m = p^d + t$ before it affects the domestic price. High transport costs can lead to a wide band and correspondingly high levels of volatility of food prices in landlocked countries - see Dana et al. (2006) on Malawi and Zambia and Daviron (2008) on francophone Africa.
2. Governments may run successful stabilization policies serving to insulate domestic consumers from changes in world prices. Governments of food-exporting countries can do this through export

controls or variable export taxes. Many important Asian rice producing-consuming countries have adopted policies of this type - see Dawe (2007) and Timmer (2010). Such actions have the potential to reduce volatility in the protected markets at the cost of increasing volatility on the residual world market. Governments of food-importing countries can use variable import tariffs, possibly in conjunction with purchases for or sales from a food security stock - see Jayne & Tschirley (2010) and Gilbert & Tabova (2011). However, if poorly managed, such policies can aggravate volatility.

3. The prices which are regarded as measuring world prices may be unrepresentative of actual transactions prices in world trade. I suggest below that this is the case in the rice market. For other food commodities, such as cassava, there is no recognized international price. The second way is that a discrepancy of this sort can emerge is if there are substantial regional or grade differences. White maize, which is the principal staple food in most of eastern and southern Africa, is only partially substitutable by yellow maize, grown in North America and Europe. Changes in the widely quoted Chicago corn (yellow maize) price will therefore only be partially reflected in domestic African prices for white maize.

Supposing one or more world prices is relevant in a particular country, the second question is the rapidity of adjustment of local prices to the world price. This literature originated with Timmer et al. (1983) and Mundlak & Larson (1992) who regressed local on world prices. This procedure is problematic when prices trend, or are more generally non-stationary. Estimation in first differences, as in de Janvry & Sadoulet (2010), results in estimates which are robust with respect to non-stationarity but which may only measure short term responses. This may result in under-estimation of responses if adjustment is slow. On the other hand it does give precise and robust estimates of the impact responses.

Baffes & Gardner (2003) used the error correction specification to estimate pass-through. This gives rise to estimates of both impact and equilibrium responses. Error correction can be justified by the Granger Representation Theorem (Engle & Granger, 1987) if the two prices considered are cointegrated but that hypothesis is not directly tested. When three or more prices are considered, as in this chapter, the supposition that there is a single cointegrating vector is problematic.

Following Ardeni (1989), it has become standard to adopt cointegration-based methods (Johansen, 1988). This approach has four advantages over more traditional techniques:

1. The number of cointegrating vectors is determined by the data. This is important if there is more than a single contender for "world price".
2. Short run adjustment responses are distinguished from equilibrium outcomes (if present).
3. The equilibrium pass-through is not restricted to be unity. This allows for the possibility that local prices are either more or less volatile than world prices.
4. Adjustment of national and world prices is considered symmetrically allowing the possibility of reverse pass-through from national prices to the world price as well as the forward pass-through from world to national prices.

Baulch (1997) and criticized the cointegration approach as failing to allow for the wedge between export and import parity prices, illustrated above in Figure 7.1. These authors prefer to adopt a switching regime model. By contrast, Rapsomanikis & Karfakis (2008) use the Balke & Fomby (1997) threshold cointegration model to accomplish the same objective.

The analysis in this chapter relies on relatively short time series which makes these more sophisticated methods unattractive. My objective is that of modelling the local impact of recent extreme movements in world prices. In these cases, the resulting movements in the parity prices will almost certainly force an adjustment of local prices. Balke & Fomby (1997) report that standard cointegration test procedures, such as the Johansen (1988) VAR-based test, work reasonably well if the true process has the threshold structure with reversion taking place only beyond the (in this case export and import parity) thresholds. I therefore combine VAR estimates of impulse response functions with cointegration analysis to obtain estimates both of immediate impacts and, where cointegration is established, of long run responses.

Maize

Maize is the main staple food in most of Southern and Eastern Africa. It is less important in West Africa but also forms an important component of the more varied diet in Latin America. In the developed world, maize is predominantly used as a livestock feed and in North America, increasingly, as a biofuel feedstock. Different maize varieties have different colours - in Europe and the Americas, yellow maize (corn) dominates while in Southern and Eastern Africa, white maize is preferred for human consumption while yellow maize is predominantly used as a feedstock.

There are two candidates for regional or world prices in maize. The standard reference price is the Chicago Board of Trade (CBOT)[4] corn (i.e. yellow maize) futures price. However, for the African countries (Kenya and Malawi), it seems possible that the South African Futures Exchange (SAFEX) white maize futures price may be relevant. Figure 7.2 charts maize prices from January 1999 (Benin: January 2005) to December 2009 at the national level for the three African countries under consideration and Peru.[5] In addition, the figure shows the South African (SAFEX) and U.S. (Chicago) free market prices. All prices have been converted to United States Dollars per tonne at the prevailing exchange rate. The two international prices move closely together but the other four prices only track these approximately. The Malawian price series shows three “hungry season” peaks corresponding to the poor harvests in 2001 and 2005 and to the spike in world prices in 2008. The Kenyan series shows much less variation prior to the 2007-08 spike but prices appear to have been generally higher than those in Malawi.

Table 7.1 lists the nominal and real price ranges in each of the four countries considered (columns 3 and 4) and also the 2005-09 price change (columns 1 and 2). The range measures the maximum extent of the price spike while the change shows the long run impact, if any. The final two columns give the standard deviations of price changes over the period. Care must be taken in the interpretation of these real prices as, in countries in which maize forms a substantial component of the household budget, deflation makes little sense for poor households.

Over the 2005-09 period, price rises are comparable to those in world markets except in Benin. Maize prices in Benin were very high in 2005 for local reasons and hence the 2005-09 price rise appears misleadingly modest - see Figure 7.2. A 2007 base would have given very different results. Effective price stabilization has resulted in relatively stable prices in Kenya and, to the extent that this interpretation can be sustained, the maize price has declined in real terms. It is evident that that maize price variability has been acute in Malawi, whatever basis is used for making the judgement.

Statistical analysis of the nominal dollar maize price series over the five years 2005-09 confirms that the series are non-stationary, although this result in quite marginal for the SAFEX series (DF = -2.74 against a 5 percent critical value of -2.91). This near stationarity precludes the use of cointegration analysis to analyse the inter-relationship with the series using the Johansen procedure on a pair-wise basis for each of the Beninois, Kenyan, Malawian and Peruvian (logarithmic) prices and each of the log exchange prices gives within a VAR(2),

4 CBOT is now part of the Chicago Mercantile Exchange (CME) group.

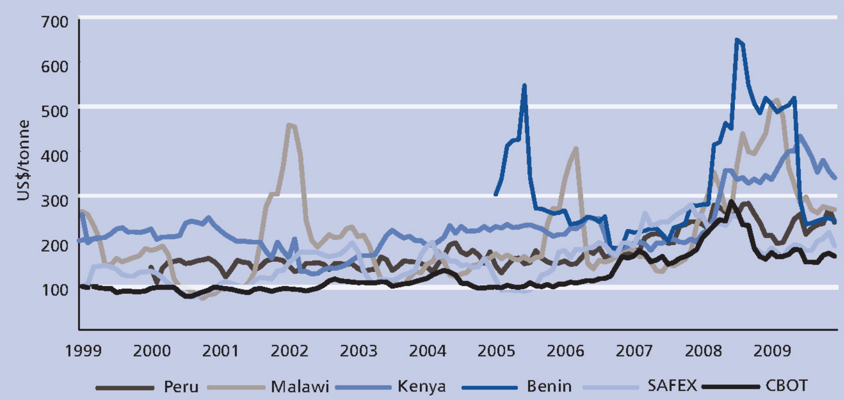
5 National prices are medians of prices across a range of locations - see Gilbert (2010) for details.

Table 7.1: Maize price changes: 2005-09

	January 2005 - December 2009		Price range over the same period		Standard deviation of monthly changes	
	Nominal	Real	Nominal	Real	Nominal	Real
	....%					
CBOT	71.5	50.8	204.3	140.7	26.1	24.3
SAFEX	67.2	47.1	223.9	178.3	31.3	29.3
Benin	5.0	-18.4	223.3	217.5	38.7	37.0
Kenya	62.9	-14.2	141.8	96.3	27.7	27.9
Malawi	59.2	36.3	380.6	288.2	52.3	49.4
Peru	65.3	27.7	91.1	51.5	16.5	15.1
Average price	41.7	24.6	130.4	126.6	20.0	20.3

Note: The first two columns of the table gives the percentage change in the free market (Chicago and SAFEX) maize prices and local rice prices respectively converted to United States Dollars and local prices deflated by the local commodity price index (CPI) at national prices or the Advanced Countries Export Unit Values (exchange prices and average price) over the period January 2005 – December 2009. The second two columns give the percentage range between the maximum and minimum prices over the same period and the final two columns the standard deviation of logarithmic price changes on an annual basis (i.e. the standard deviation of monthly price changes annualized by multiplying by $\sqrt{12}$). Source for exchange rates, CPI and export unit value indices: IMF, International Financial Statistics.

Figure 7.2: US Dollar maize prices: 1999-2009



fails to reject $rank(\alpha\beta')=0$ implying each of the series is stationary. This suggests moving to longer series, I therefore analyse data from 1999-2009 for the three exchange prices and for Kenya, Malawi and Peru (price starts in 2000), for which long price series are available.[6]

[6] These prices are the national prices as stated by government departments or international agencies: Kenya - Ministry of Agriculture; Malawi - World Bank; Peru - Ministry of Agriculture. Prices converted to United

Table 7.2: Statistical properties of maize price series: 1999-2009					
	Stationarity	Trace cointegration tests			Implied cointegrating vectors
		rank = 0	rank ≤ 1	rank ≤ 2	
CBOT	ADF (1) - 1.36	18.7 [1.4%]	2.23 [13.6%]	-	1
SAFEX	ADF (1) - 2.16				
Kenya	DF - 1.22	37.8 [0.4%]	10.7 [23.6%]	2.51 [11.3%]	1
Malawi	ADF (2) - 2.58	34.9 [1.1%]	16.7 [3.1%]	2.46 [11.7%]	2
Peru	ADF (2) - 2.37	44.3 [< 0.1%]	15.5 [4.8%]	1.94 [16.4%]	2
Kenya Malawi Peru	-	34.7 [1.2%]	8.88 [38.4%]	0.83 [36.4%]	1

Note: Sample period is April 1999 (Peru and final row: April 2000) – December 2009. The ADF lag was selected over the range 0-3 on the basis of the Akaike Information Criterion, the 5 percent critical value is – 2.91. Cointegration is examined using the Johansen trace tests within a set of bivariate VAR(2) models. The combined first two rows consider bivariate cointegration between the Chicago and SAFEX price. The middle three rows consider trivariate cointegration between Chicago, SAFEX and each national price in turn. The reported statistics test $rank(\alpha\beta') \leq 1$ $(r = 0,1,2)$. The final row considers cointegration among the three national prices. Tail probabilities in "[.]" parentheses.

ADF tests confirm that the four (logarithmic) series are all non-stationary over this longer sample, although this result is marginal for the Malawian series - see the initial column of Table 7.2. This allows application of the Johansen (1988) procedure to analyse cointegration. I first consider the two exchange prices. A preliminary check establishes that I can consider this within a VAR(2) framework. I fail to reject the hypothesis that $rank(\alpha\beta') \leq 1$ but reject the hypothesis that $rank(\alpha\beta') = 0$ confirming that there is a single cointegrating vector - see the second and third column of Table 7.2. The estimated α coefficients are $\left(\hat{\alpha}_{Chicago}, \hat{\alpha}_{SAFEX}\right)$ =(−0.047, 0.087) with standard errors (0.023, 0.034). Both coefficients differ significantly from zero implying each market reacts to the other, but the coefficient for SAFEX is approximately double that for the Chicago market consistent with the leading role played by Chicago. I also fail to reject the hypothesis that this is a unit cointegrating vector (χ_1^2 =0.46 with p-value 50.0 percent) implying that over the long term the Chicago and SAFEX prices may be seen as moving in line with each other.

I now add the Kenyan, Malawian and Peruvian prices in turn to the VAR. Given that we already know that the Chicago and SAFEX prices are cointegrated, cointegration of the African prices requires $rank(\alpha\beta') = 2$. This result is established for Malawi and Peru but not for Kenya - see Table 7.2. Despite its proneness to weather-related shortages, in the long term, the Malawian maize market appears integrated with world markets. In line with the visual impression obtained from Figure 7.2, this is not true of Kenya.

States Dollars per kilogram at prevailing exchange rates; source: IMF, International Financial Statistics.

It is interesting to look at the Malawian and Peruvian cases in greater detail. As $rank(\alpha\beta') = 2$, we have only identified a two dimensional basis for the space in which the cointegrating vectors lie. It is therefore open to us to rotate the estimated cointegrating vectors within that space. As we have already established that the Chicago and SAFEX prices are cointegrated with unit cointegrating vector, we can reasonably impose $\beta_1' = (1,-1,0)$ with $\alpha_{13} = 0$ (i.e. the two local prices do not react to temporary discrepancies between SAFEX and Chicago). Thereafter, the procedure differs between the two countries.

Malawi: Normalizing the second cointegrating vector, I hypothesize that this depends equally on the two exchange prices so $\beta_2' = (1/2,\ 1/2,\ -1\ )$.

As it is reasonable to suppose that the Chicago price is unaffected by the maize situation in Malawi, we can impose $\alpha_{21} = 0$ in the Malawian case. More tendentiously, we can impose the same condition on the SAFEX price, i.e. $\alpha_{22} = 0$. This implies a total of 5 restrictions on the $\alpha\beta'$ matrix. The likelihood ratio fails to reject these restrictions ($\chi_5^2 = 2.39$ with p-value 79.2 percent) implying an acceptable identification. The estimated α matrix is:

$$\begin{pmatrix} \hat{\alpha}'_{Chicago} \\ \hat{\alpha}'_{SAFEX} \\ \hat{\alpha}'_{Malawi} \end{pmatrix} = \begin{pmatrix} -0.056 & 0 \\ 0.087 & 0 \\ 0 & 0.014 \end{pmatrix} \quad \text{with standard errors:} \begin{pmatrix} 0.026 & - \\ 0.038 & - \\ - & 0.028 \end{pmatrix}$$

Conditional on the acceptability of these restrictions, the estimates show that we can reject both the null hypotheses $\alpha_{Chicago} = 0$ and $\alpha_{SAFEX} = 0$ implying that the markets are interdependent. However, in line with the earlier result, SAFEX reacts more than Chicago to deviations from parity.

Peru: It seems unlikely that the South African white maize price can influence the yellow maize price in Peru so I set $\beta_2 = (1, 0, -1)$. As in the case of Malawi, we can use a "small country" restriction to suppose that Peruvian prices have no influence on the exchange prices allowing us to set $\alpha_{21} = \alpha_{22} = 0$. As again in Malawi, the likelihood ratio fails to reject these restrictions ($\chi_4^2 = 1.02$ with p-value 90.7 percent) implying an acceptable identification. The estimated α matrix is now

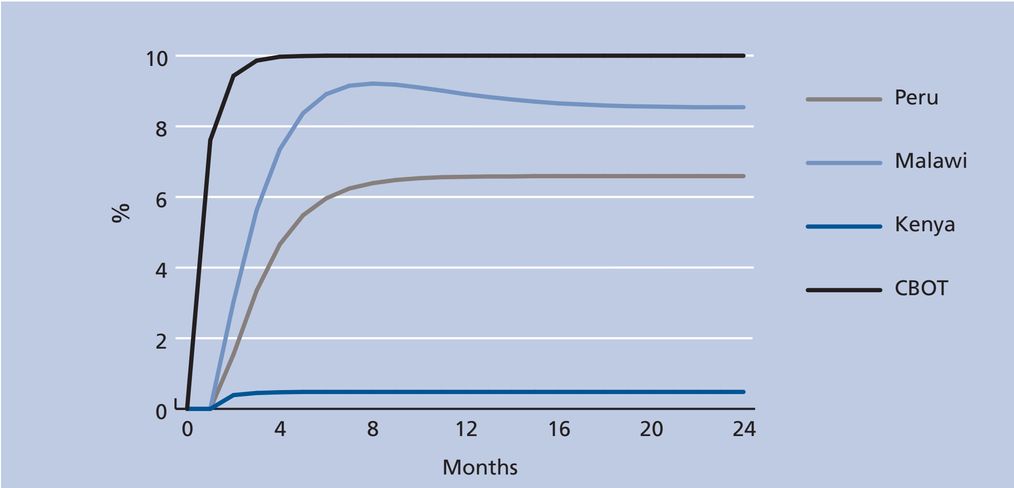
$$\begin{pmatrix} \hat{\alpha}'_{Chicago} \\ \hat{\alpha}'_{SAFEX} \\ \hat{\alpha}'_{Peru} \end{pmatrix} = \begin{pmatrix} -0.062 & 0 \\ 0.074 & 0 \\ 0 & 0.409 \end{pmatrix} \quad \text{with standard errors:} \begin{pmatrix} 0.026 & - \\ 0.039 & - \\ - & 0.070 \end{pmatrix}$$

It is reassuring to see that the estimated α coefficients for the two exchange prices are very similar to those in the Malawian model.

The final row of Table 7.2 looks at cointegration among the three national prices ignoring the exchange prices. The test establishes a unit cointegrating rank implying two common trends consistently with cointegration of the Malawian and Peruvian prices with the exchange prices, which are themselves cointegrated, but lack of cointegration of the Kenyan prices with the exchange prices.

Impulse response functions are computed using a cointegrated VAR(1), i.e. a CVAR(1), relating the change in the maize prices in Kenya, Malawi and Peru to changes in the Chicago maize price (Figure 7.3) . (To maintain simplicity, only a single exchange price was considered in the calculation of impulse responses). The VAR specification imposes block exogeneity such that the changes in the three national prices only enter their respective national equations. The Malawian and Peruvian price equations include an unrestricted error correction term defined in terms of the two month lags of the Chicago price and the national price. Consistently with

Figure 7.3: Estimated impulse response functions: maize



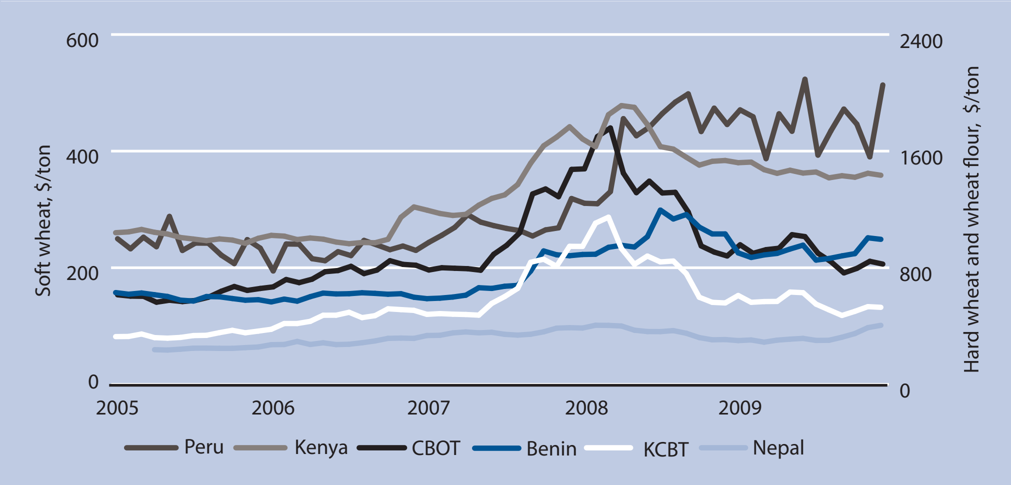
the results reported in Table 7.2, no error correction term is included in the Kenyan equation. Estimation is by Full Information maximum Likelihood (FIML). Results are reported in Tables 7.7 to 7.9.

The impulse response consists of a 7.61 percent shock to the Chicago price which, once the first order autoregressive term is taken into account, generates a long run 10 percent rise in that price. Figure 7.4 charts the resulting rises in the three national prices. The long run effects are similar in Malawi or Peru where local prices rise by 7-9 percent with an approximate six month lag. Malawi is seen as having both a greater and a faster pass-through than Peru. However, consistently with the cointegration analysis, there is almost no impact on Kenyan maize prices where estimated pass-through is effectively zero.

Both the Kenyan and Malawian governments actively intervene in their respective maize markets. Jayne et al. (2006) date that between 1995 and 2004, the Kenyan National Cereals and Produce Board (NCPB) purchased between 15 percent and 57 percent of the domestically marketed maize output while the Malawian Agricultural Development and Marketing Corporation (ADMARC) purchased between 12 percent and 70 percent of output over the same period. The boards have the ability to set maximum purchase prices for the private sector. They also control imports and exports through licensing and duties. There is considerable uncertainty as to the likely availability of import licenses and import duties are varied according to the economic and political circumstances (Jayne & Tschirley, 2010). In principal, therefore, it would not be surprising to find very limited pass-through in either country.

Despite this substantial government presence, the results reported above and the statistics in Table 7.1 indicate that while Kenya has been relatively successful in stabilizing its domestic maize prices, Malawi has been much less so. Jayne & Tschirley (2010) provide a detailed discussion of both Kenyan and Malawian policy in 2007-08. An important factor working against cointegration in Kenya was the January 2009 decision to abolish the 50 percent tariff on maize inputs. Despite this, Kenyan prices remained well above import parity for much of 2009 because of transport shortages from the port of Mombasa. Kenya therefore

Figure 7.4: US Dollar wheat prices: 1999-2010



remained autarchic in practical terms throughout the period under study. This contrasts with the situation in Malawi. There, an over-optimistic official forecast of a 2007/08 maize surplus led the government to authorize exports, probably destined for neighbouring Zimbabwe. The surplus failed to materialize but government, unwilling to lose its reputation as having moved Malawi from being a food deficit to a food surplus country remained unwilling to license imports resulting in prices rising well above import parity. Kenya therefore stands out as having effectively insulated local maize prices against movements in world markets albeit at the cost of a higher average price.

Wheat

Wheat is the major grain entering human consumption in Europe and North America. It is also consumed throughout the developing world where it nevertheless generally forms a smaller component of diets. In tropical countries, wheat is almost entirely an imported crop, either directly or in the form of wheat flour.

Wheat is traded on a number of major futures markets. The two major U.S. markets are the Chicago (CBOT) market for soft wheat, mainly used in confectionary, and the Kansas City Board of Trade (KCBT) market for hard wheat, used for bread. The Paris (Marché à Terme International de France - MATIF) market for unmilled wheat has become increasingly important as a pricing basis for Russian and Ukrainian deliveries. I focus on the two principal U.S. prices. The CBOT[7] price for soft wheat, which is predominantly used for confectionary, is the most widely used reference price. The reference price for hard (durum) red winter wheat, used in making bread, is the KCBT price.

Figure 7.4 charts wheat prices from January 2005 to December 2009 at the national level

7 Originally Chicago Board of Trade, now Chicago Mercantile Exchange. I also investigated using the MATIF price, which has become more important as a reference over recent years, but this did not add to the explanation.

Table 7.3: Wheat price changes: 2005-09						
	January 2005 - December 2009		Price range over the same period		Standard deviation of monthly changes	
	Nominal	Real	Nominal	Real	Nominal	Real
	........................%					
CBOT	34.3	16.7	212.1	155.2	26.2	25.4
KCBT	60.2	39.3	256.7	188.1	28.5	27.5
Benin (wheat flour)	57.5	22.3	110.2	62.9	18.2	16.7
Kenya (wheat flour)	37.7	-27.5	97.6	47.2	13.9	12.1
Nepal	68.4	15.9	70.2	39.6	13.8	12.0
Peru	105.6	58.8	169.1	112.6	39.8	38.6
Average price	54.6	33.0	82.3	55.1	25.9	24.1

Note: The first two columns of the table gives the percentage change in the free market wheat prices and local wheat or wheat flour prices respectively converted to United States Dollars and local prices deflated by the local CPI (national prices) or the Advanced Countries Export Unit Values (exchange prices and average price) over the period January 2005 – December 2009. The second two columns give the percentage range between the maximum and minimum prices over the same period and the final two columns the standard deviation of logarithmic price changes on an annual basis. Source for exchange rates, CPI and export unit value indices: IMF, International Financial Statistics.

for Nepal and Peru (left axis) and the wheat flour prices for Benin and Kenya (right axis).[8] In addition, the figure shows the CBOT price for soft wheat (left axis) and the KCBT price for hard wheat (right axis). All prices have been converted to United States Dollars per tonne at the prevailing exchange rate. The prices move closely together over 2006-07 but diverge in 2008-09 when the Peruvian wheat price and the Kenyan and Nepalese flour prices fail to follow the fall in world prices.

Table 7.3, which has the same format as Table 7.1 for maize, lists the nominal and real price ranges in each of the four countries considered (columns 3 and 4) and also the 2005-09 price change (columns 1 and 2). The range measures the maximum extent of the price spike while the change shows the long run impact, if any. The final two columns give the standard deviations of price changes over the period.

The KCBT hard wheat price has tended to rise relative to the CBOT soft wheat price over the period concerned. Wheat and wheat flour prices in Benin, Kenya and Nepal were less variable than the exchange prices although, except in Kenya, the overall rise in prices was comparable. Peruvian wheat prices were the most variable on a month-to-month basis even though there was no pronounced spike in prices in 2008.

Statistical analysis of the nominal dollar maize price series over the five years 2005-09 confirms that the series are non-stationary - see Table 7.4. I therefore turn to cointegration analysis following the sequential procedure previously adopted for maize. Accordingly, I first consider the two exchange prices. A preliminary check establishes that I can consider

[8] The Kenyan price is the official national price while the remaining national prices are medians of prices across a range of locations - see Gilbert (2010) for details.

Table 7.4: Statistical properties of wheat price series: 2005-09

	Stationarity	Trace cointegration tests			Implied cointegrating vectors
		rank = 0	rank ≤ 1	rank ≤ 2	
CBOT	ADF (1) - 1.52	12.2 [15.1%]	3.45 [6.3%]	-	0
KCBT	ADF (1) - 1.61				
Benin	DF - 0.67	21.4 [31.4%]	7.82 [49.2%]	1.73 [18.9%]	0
Kenya	ADF (1) - 1.25	20.5 [40.0%]	8.68 [40.3%]	2.60 [10.7%]	0
Nepal	ADF (3) - 1.38	18.9 [50.9%]	6.70 [61.8%]	0.97 [32.5%]	0
Peru	ADF (2) - 0.28	26.6 [11.5%]	11.2 [20.4%]	3.33 [6.8%]	0
Benin Kenya Peru	-	31.1 [3.5%]	13.8 [8.7%]	1.95 [16.2%]	1 or 2

Note: ADF sample period is January 2005 (Nepal: August 2005) – December 2009. The ADF lag was selected over the range 0-3 on the basis of the Akaike Information Criterion. 5 percent critical value is – 2.92. Cointegration sample January 2005 (Benin and final row: March 2005; Nepal: June 2005) – December 2009. Cointegration is examined using the Johansen trace tests within a set of bivariate VAR(2) models. The combined first two rows consider bivariate cointegration between the Chicago and Kansas City prices. The next four rows consider trivariate cointegration between Chicago, Kansas City and each national price in turn. The final row considers cointegration among the three listed national prices. The reported statistics test $rank(\alpha\beta') \le 1$ $(r = 0,1,2)$. The final row considers cointegration among the three national prices. Tail probabilities in "[.]" parentheses.

this within a VAR(2) framework. I fail to reject both the hypothesis that $rank(\alpha\beta') \le 1$ and the hypothesis that $rank(\alpha\beta') = 0$ indicating lack of cointegration - see the second and third column of Table 7.4. Extension of the sample back to March 1999 gives the same result.[9] A test fails to reject block exogeneity of the VAR - $\chi^2(4) = 5.55$ with tail probability 23.5 percent implying the comovement in the hard and soft wheat prices is entirely owing to common shocks. This is consistent with the view that hard and soft wheat are different grains and not different grades of the same grain. Adding each of the four national prices in turn fails to produce any departure from non-stationarity indicating that none of the national prices is cointegrated with either of the world prices.

Despite this negative result, it is nevertheless possible to establish cointegration between three of the four national wheat prices (those for Benin, Kenya and Peru) - see Table 7.4, final row.[10] The Johansen trace test fails to reject $rank(\alpha\beta') \le 2$ but does reject $rank(\alpha\beta') = 0$ implying at least one cointegrating vector. The intermediate hypothesis $rank(\alpha\beta') \le 1$ is rejected at the

9 Test statistics for *rank* 0 and *rank* ≤ 1 respectively 7.39 and 2.30 with *p*-values 53.9 percent and 13.0 percent respectively.

10 Nepal was excluded because it limits the sample size and because it does not appear to enter any of the cointegrating relationships.

10 percent but not the 5 percent level. Analysis of the estimated α and β matrices for the case of $rank(\alpha\beta') = 2$ allows imposition of the restrictions

$$\beta = \begin{pmatrix} -1 & 0 \\ 1 & 0 \\ 0 & -1 \end{pmatrix},$$

implying a unit cointegrating vector linking each pair of prices, and

$$\alpha = \begin{pmatrix} 0.258 & 0 \\ 0 & 0 \\ 0 & 0.206 \end{pmatrix},$$

with estimated standard errors

$$\alpha = \begin{pmatrix} 0.065 & 0 \\ 0 & 0 \\ 0 & 0.080 \end{pmatrix}.$$

The zero restriction on the second row of the α matrix implies that the Beninois and Peruvian prices react to discrepancies relative to the Kenyan price but not vice versa. The likelihood ratio test on the combined set of restrictions is $\chi^2(6) = 8.21$ with tail probability 22.3 percent.

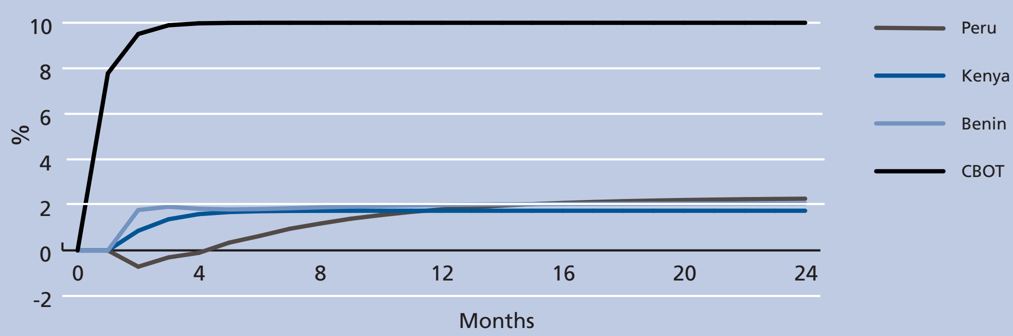
This surprising result suggests that, relative to the group of countries we have considered here, the Kenyan wheat flour price is playing the role that we would have expected the world price to play. Kenya is not an important player in the world wheat market. The result should therefore be interpreted as indicating that movements in the Kenyan wheat flour price were representative of the prices other countries were paying and not indicative of a causal relationship.

As in the case of maize, we compute impulse response functions using a cointegrated CVAR(1), relating the change in the wheat or what flour prices in Benin, Kenya and Peru to changes in the Chicago wheat price. However, differently from the maize case and reflecting the lack of cointegration of the national prices with the Chicago price and the apparent representativeness of the Kenyan price, the VAR specification imposes block exogeneity such that the changes in the Beninois and Peruvian prices only enter their respective national equations while changes in the Kenyan price do affect the other two prices. Similarly, the Beninois and Peruvian price equations included as an unrestricted error correction term the two month lag of both the Kenyan price and the respective national price. No error correction term is included in the Kenyan equation. Estimation was by Full Information maximum Likelihood (FIML). Results are reported in Table 7.8.

The impulse response consists of a 7.78 percent shock to the Chicago price which, once the first order autoregressive term is taken into account, generates a long run 10 percent rise in that price.[11] Figure 7.5 charts the resulting rises in the three national prices. The long run effects are similar in the three countries with a pass through of 2 percent, i.e. 20 percent of the shock to the Chicago price. Adjustment is seen as much slower in Peru than either Malawi or Peru.

[11] Slightly larger than in the case of maize reflecting differences in the two samples.

Figure 7.5: Estimated impulse response functions: wheat



These results contrast strongly with those obtained earlier for maize. The econometric results indicate only a low level of pass-through from world to local prices with no clear-cut long-run relationships between the two groups. Nevertheless, the prices in Benin, Kenya and Peru, widely separated countries with no important trade links, appear to move together. Furthermore, local wheat and wheat flour prices did rise in 2008 indicating pass-through of the price rises. But, as world prices fell back in 2009, the local prices failed to follow through. The countries we have examined appear to have suffered from the 2007-08 price rises but failed to benefit from the price falls.

There are several possible explanations for the absence of comovement between the national wheat prices and the corresponding exchange prices. One reaction is that the five year period considered is too short to look at cointegration which, if it exists, is a long run relationship. Nevertheless, if one is interested in pass-through, this is the relevant horizon. This analysis has pointed to the failure of national wheat and wheat flour prices to decline from mid-2008 to the same extent as exchange prices. Possible explanations might relate to forward pricing or long term contracting arrangements which locked importing countries into high prices, or perhaps the exercise of market power either by importers or parastatal grains agencies.

Rice

Rice is the main staple grain throughout most of Asia. It also forms an important component of the diet in the remainder of the world, particularly in West Africa. The standard reference price is the Bangkok spot price reflecting the fact that Thailand is the major world rice exporter. This spot price is related to the white rice futures contract on the Agricultural Futures Exchange of Thailand (AFET) although trading in this contract remains very thin. Rough (i.e. unmilled) rice is also traded in Chicago on CBOT but volumes are low relative to those on other Chicago grains markets and the contract is regarded as being primarily of domestic interest.

Figure 7.6: US Dollar rice prices: 2005-09

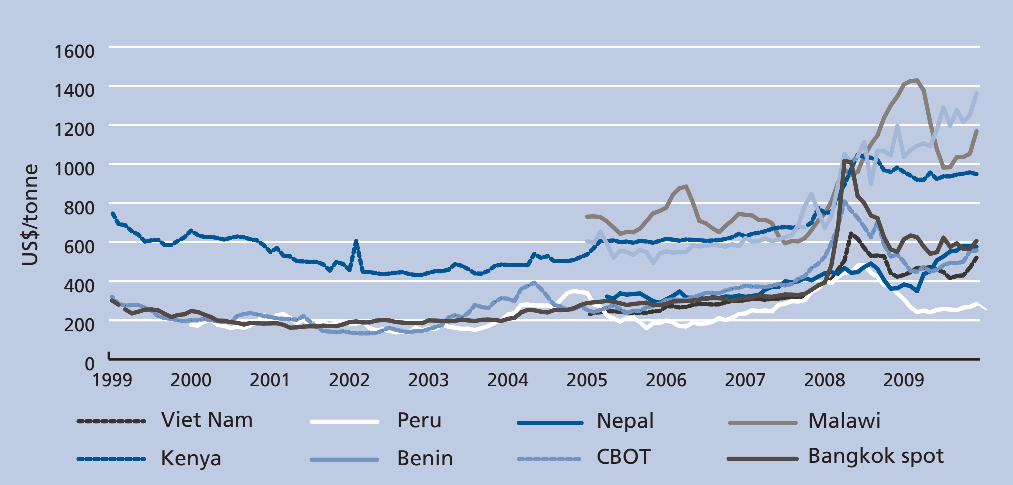


Figure 7.6 charts local rice prices from January 2005 to September 2010 for the six countries under consideration, and in addition includes the Bangkok and CBOT prices.[12] All prices have been converted to United States Dollars per tonne at the prevailing exchange rate. The figure shows the prices falling into two groups: three low price countries (Nepal, Peru and Viet Nam) and the three high price African countries (Benin, Kenya and Malawi). The world free market price is typically closer to prices in the non-African group but rose to African levels in 2008. It is apparent from Figure 7.6 that the Bangkok price has been much more variable than any of the national prices.

Table 7.5 shows considerable variation across countries in the evolution of deflated rice prices over 2005-09. Although the change in the Bangkok price over the five year period is comparable to that in the six countries considered, the range between the maximum and minimum prices was approximately double for the former. As was the case of maize, care must be taken in the interpretation of these real prices in countries in which rice forms a substantial component of the household budget. With this qualification, the statistics show a decline in real rice prices in Peru over the five years 2005-09 and almost no change in Kenya, while real prices have risen by 20 percent-40 percent in Malawi and Nepal, and 60 percent-75 percent on Benin and Viet Nam. Prices in each country clearly reflect specific local conditions. This motivates consideration of the average world price, averaging over the six countries considered. The final row of Table 7.5 confirms this greater variability of the free market prices relative to the average price.

Table 7.6 lists the statistical properties of the nine nominal United States Dollar rice price series considered in Table 7.5. ADF tests show that all nine logarithmic series are

[12] The Chicago price relates to rough rice whereas the remaining prices are for white rice. I have used a conversion factor of 1.67 to convert the Chicago price onto a milled basis. Rice prices for Malawi, Nepal, Peru and Viet Nam are calculated as the medians of prices across localities - see Gilbert (2010). Local prices are not available for Benin and Kenya: Benin - “regular rice”, Cotonou, urban market, F.CFA/kg, Ministère du commerce, Direction de la Promotion du Commerce Intérieur; Kenya - Average wholesale price, rice grade 2, loose, Ksh/kg, Kenyan National Bureau of Statistics. Because of different specifications, prices are not perfectly comparable across countries.

Table 7.5: Rice price changes: 2005-09						
	January 2005 - December 2009		Price range over the same period		Standard deviation of monthly changes	
	Nominal	Real	Nominal	Real	Nominal	Real
	....................%					
Free market (Bangkok)	110.1	82.7	266.1	181.8	30.8	29.0
Free market (Chicago)	119.9	91.2	240.9	165.4	21.9	19.3
Benin	124.9	74.6	175.7	99.9	27.0	25.1
Kenya	76.5	-7.1	95.2	33.3	10.7	9.6
Malawi	59.9	36.8	140.3	91.6	19.3	16.9
Nepal	79.3	23.5	100.6	42.7	19.8	15.8
Peru	-16.7	-35.6	202.4	135.1	27.4	26.2
Vietnam	129.7	58.1	184.1	92.6	19.4	18.2
Average price	89.0	62.5	89.2	71.5	10.1	9.3

Note: The first two columns of the table gives the percentage change in the free market rice prices and local rice prices respectively converted to United States Dollars and local prices deflated by the local CPI over the period January 2005 (Nepal and Average: April 2005) – December 2009. The second two columns give the percentage range between the maximum and minimum prices over the same period and the final two columns the standard deviation of logarithmic price changes (at an annual rate). The deflator for the free market price and the average price is Advanced Countries Export Unit Values. Source for exchange rates, CPI and export unit value indices: IMF, International Financial Statistics.

non-stationary. Table 7.6 also reports Johansen cointegration tests. In the first two rows, I ask whether the Bangkok and Chicago free market prices are cointegrated. The results are inconclusive but consistent with a lack of cointegration - the Johansen test marginally fails rejects $rank(\alpha\beta') \leq 1$ which would imply both prices are stationary, contradicting the clear results from the ADF tests. Ignoring that result, the second test also fails to reject $rank(\alpha\beta') = 0$. I conclude that the rank is zero and there is no cointegration. The ADF test on the logarithmic difference of the two prices gives a statistic of ADF(1) = -2.26.

To check on this result, the test was rerun using the longer sample March 1999 - December 2010. ADF tests again confirm non-stationarity (Bangkok, ADF(2) = -2.58; Chicago, ADF(1) = - 0.76). The Johansen test now gives a clear rank zero result. The tests fail to reject both $rank(\alpha\beta') \leq 1$ (trace statistic 0.50 with p-value 48.1 percent) and $rank(\alpha\beta') = 0$ (trace statistic 9.74 with p-value 30.7 percent). I conclude that the two rice exchange prices are indeed not cointegrated.

I now return to the original 2005-09 sample and add the six national rice prices, one at a time, to give a trivariate VAR(2). Whereas in the case of maize, where the two exchange prices are cointegrated, I tested for $rank(\alpha\beta') = 2$, in rice, lacking this cointegration, I test for $rank(\alpha\beta') = 1$.

A cointegrating rank of one is established for Peru. For Viet Nam, the failure to reject $rank(\alpha\beta') = 0$ is marginal. The test outcomes are problematic for Kenya and Malawi where the test rejects $rank(\alpha\beta') \leq 1$ (marginally in the Kenyan case) which would imply all three prices are stationary, again contradicting the results of the ADF tests. However, if we override

Table 7.6: Statistical properties of rice price series: 2005-09

	Stationarity	Cointegration tests			Implied cointegrating vectors
		rank = 0	rank ≤ 1	rank ≤ 2	
Free market (Bangkok)	ADF (2) -1.12	12.8 [12.3%]	3.94 [4.7%]	-	0 (?)
Free market (Chicago)	ADF (1) -1.30				
Benin	ADF (2) 0.43	24.2 [19.9%]	6.59 [63.1%]	0.31 [57.5%]	0
Kenya	ADF (1) - 0.64	43.1 [0.1%]	11.2 [21.5%]	4.00 [4.6%]	1 (?)
Malawi	ADF (3) - 1.36	32.5 [2.3%]	12.7 [12.8%]	5.56 [1.8%]	1 (?)
Nepal	DF - 0.23	25.2 [15.8%]	5.80 [72.1%]	0.27 [60.2%]	0
Peru	ADF (1) - 1.19	31.8 [2.8%]	11.2 [20.3%]	3.70 [5.4%]	1
Viet Nam	ADF (1) - 0.89	28.6 [7.0%]	7.68 [30.6%]	2.28 [13.1%]	0 or 1
Average price	ADF (1) - 0.22	31.9 [2.8%]	9.37 [33.9%]	1.52 [21.8%]	1

Note: Sample period is March 2005 (rank tests; Nepal and Average – June 2005) and May 2005 (ADF tests; Nepal and Average – August 2005) – December 2009. The ADF lag was selected over the range 0-3 on the basis of the Akaike Information Criterion. 5 percent critical value is – 2.91. Cointegration is examined using the Johansen trace tests within a set of bivariate VAR(2) models considering the free market price with each of the other prices. The reported statistics test $rank(\alpha\beta') \leq 1$ $(r = 0,1,2)$. The final row considers cointegration among the three national prices. Tail probabilities in "[.]" parentheses.

this result, we again establish a cointegrating rank of unity. By contrast, cointegration is completely rejected in the case of Benin and Nepal where local prices continued to rise through 2009 while they fell back elsewhere - see Figure 7.6. The price constructed as the average of the six national prices does appear to be cointegrated with the two free market prices.[13]

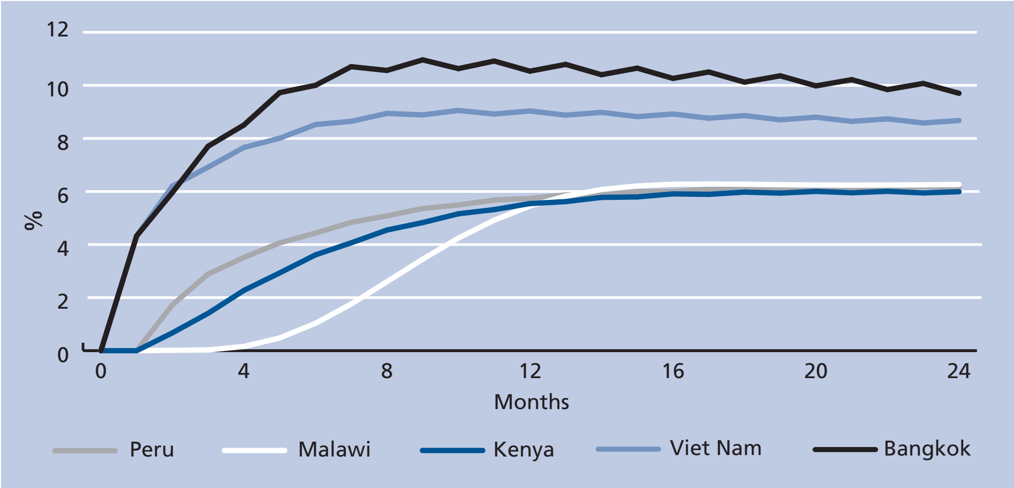
The results reported in Table 7.6 demonstrate possible links between the various national prices, with the exception of those of Benin and Nepal,[14] and the two world prices. To enable examination of these links in greater detail, I report the estimated α and β coefficients in Table 7.7.[15] The first two columns of the table test whether the β coefficients for respectively the Bangkok and Chicago prices can be set to zero. I fail to reject the hypothesis with respect to the Chicago price for Kenya and, marginally, for Viet Nam (column 2). The same hypothesis with respect to the Bangkok price is rejected for all four countries (column 1). The third and

13 This result is also problematic. If any one of the six prices composing the average is not cointegrated with the exchange prices, the average itself cannot be cointegrated. However, using relative short samples, apparently contradictory results of this sort can emerge.

14 Agricultural prices in Nepal relate to Indian prices in the first instance, and not to world prices.

15 Benin and Nepal are excluded because of lack of cointegration.

Figure 7.7: Impulse response functions: rice



fourth columns of the table report the estimated β coefficients, relative to a normalization of minus one on the coefficient of the national price. I set the coefficient β_{Ch} on the Chicago price to zero for Kenya and Viet Nam based on the test outcomes reported in column 2 and also for Malawi where $\beta_{Ch}$is estimated as negative. Similarly, I set $\beta_{Bk} = 0$ for Peru where this coefficient was estimated as negative. In summary, the Kenyan, Malawian and Vietnamese rice prices are seen as being cointegrated with the Bangkok price and the Peruvian price with the Chicago price. The final three columns give the estimated α coefficients. The size of these coefficients α_j shows the speed at which the national prices react to any discrepancy with respect to world prices. Reaction is seen as fastest in Viet Nam and slowest in Malawi.

The final row of Table 7.7 performs the same tests with respect to the average price (which includes the Beninois and Nepalese prices). As in the cases of Kenya and Viet Nam, we fail to reject the hypothesis that $\beta_{Ch} = 0$ so that the average world price is seen as being linked only to the Bangkok price. The estimated β_{Bk} coefficient is 0.75 so prices are seen as varying by 75 percent of those on the free market. The estimated α coefficients show that while the Bangkok and Chicago prices both react to the average world price, there is no evidence of any reaction of the world price to the exchange prices.[16]

The implication of these results is that, in the rice market, the prices taken as representing world prices follow prices in local markets round the world rather than vice versa. The CBOT price for rough rice does not appear to have major relevance to world markets. In line with the generally accepted view, the Bangkok spot price appears more important. Nevertheless, that price appears to be reacting to prices in producing and consuming markets rather than determining those prices. In terms of the literature in financial economics, price discovery appears to take place in the producing and consuming markets more than in centralized spot and futures markets. Instead of asking how fast and to what extent changes in world

[16] We also considered cointegration among the six national rice prices in the same way as for wheat. Tests indicated three cointegrating vectors but no simple identification appears available.

Table 7.7: Estimated maize VAR

	$\Delta\ln Pch_t$	$\Delta\ln Pky_t$	$\Delta\ln Pmw_t$	$\Delta\ln Ppy_t$
$\Delta\ln Pch_{t-1}$	0.239 (0.091)	0.045 (0.114)	0.398 (0.182)	0.202 (0.112)
$\Delta\ln Pky_{t-1}$		- 0.075 (0.091)		
$\Delta\ln Pmw_{t-1}$			0.412 (0.076)	
$\Delta\ln Ppy_{t-1}$				- 0.298 (0.086)
$\Delta\ln Pch_{t-2}$			0.082 (0.043)	0.249 (0.048)
$\Delta\ln Pmw_{t-2}$			- 0.096 (0.027)	
$\Delta\ln Ppy_{t-2}$				- 0.378 (0.069)
R2	0.057	0.007	0.336	0.236
Standard error	6.24%	7.60%	12.44%	7.57%
Autoregression LM (7)	F7, 109 = 1.29 [26.2%]	F7, 108 = 1.11 [36.2%]	F7, 106 = 1.77 [10.1%]	F7, 106 = 1.93 [7.26%]

prices are passed through into national prices, we should therefore ask the reverse question of pass-through from national to world prices. The picture we obtain is one of an imperfectly globalized market in which the prices taken as world prices relate to residual transactions and furthermore, these supposedly representative prices do not even move closely together.

The rice VAR model is more problematic than those for maize and wheat. The model takes Malawian rice prices to be cointegrated with those in Kenya and Kenyan and Peruvian prices to be cointegrated with the Vietnamese price which appears to play the role of the average price in the results reported in Table 7.6.[17] The Bangkok spot price is also specified as cointegrated with the Vietnamese price but with the error correction terms confined to the Bangkok equation. The dynamic part of the model relates the change in each price to its own lag and to either the lagged value of the Bangkok price or the Vietnamese price, depending which fits better. The resulting model is only borderline stable, and it was necessary to impose a unit cointegrating vector between the Bangkok and Vietnamese prices to ensure that all roots lie within the unit circle. The estimate model is reported in Table 7.9.

The impulse response consists of a 4.33 percent shock to each of the Bangkok and Vietnamese prices. Combined, these generate a 10 percent cumulative impact on the Bangkok price after six months. Prices both oscillate and overshoot with the maximum at six months for the Bangkok price and ten months for the Vietnamese price. Figure 7.7 charts the resulting rises in the four national prices. The Vietnamese responses closely track the changes in the

17 As each of the prices is seen as cointegrated with one of the other prices, any pair of prices must be cointegrated. However, statistically, one may fail to find cointegration with one choice of pairs while finding it with an alternative choice.

Table 7.8: Estimated wheat VAR				
	$\Delta \ln Pch_t$	$\Delta \ln Pky_t$	$\Delta \ln Pmw_t$	$\Delta \ln Ppy_t$
$\Delta \ln Pch_{t-1}$	0.239 (0.091)	0.045 (0.114)	0.398 (0.182)	0.202 (0.112)
$\Delta \ln Pky_{t-1}$		- 0.075 (0.091)		
$\Delta \ln Pmw_{t-1}$			0.412 (0.076)	
$\Delta \ln Ppy_{t-1}$				- 0.298 (0.086)
$\Delta \ln Pch_{t-2}$			0.082 (0.043)	0.249 (0.048)
$\Delta \ln Pmw_{t-2}$			- 0.096 (0.027)	
$\Delta \ln Ppy_{t-2}$				- 0.378 (0.069)
R2	0.057	0.007	0.336	0.236
Standard error	6.24%	7.60%	12.44%	7.57%
Autoregression LM (7)	F7, 109 = 1.29 [26.2%]	F7, 108 = 1.11 [36.2%]	F7, 106 = 1.77 [10.1%]	F7, 106 = 1.93 [7.26%]

Bangkok spot price but at around 80 percent of that level. The long run effects are similar in Kenya, Malawi and Peru with a pass through of 6 percent, i.e. 60 percent of the combined initial shock. Adjustment is seen as much slower in Malawi than either Kenya or Peru.

These results contrast with those obtained for both maize and wheat. As in wheat, but unlike maize, there is no pass-through from world prices to local prices. However, as in maize, but unlike wheat, price rises in world markets are associated with comparable, although in this case smaller, rises in rice prices in local markets. The Vietnamese price follows the Bangkok price quickly and relatively closely. Prices in Kenya, Malawi and Peru follow to a lesser extent and more slowly. However, the direction of impact appears to be from the national prices to the world price and not the reverse.

Rice is therefore a case of reverse pass-through (or "pass back"). It is acceptable to take movement in the Bangkok price as an indicator of movements in national prices but there is no basis for claiming that shocks to the Bangkok price, including those from possible interventions, would affect national rice prices. Trades on the Bangkok spot market do not determine rice transaction prices away from the market in the way that trades in the CBOT corn price and SAFEX maize price determine world maize transactions prices. Nevertheless, the Bangkok spot price may be seen as a price thermometer, but one which exaggerates changes in temperature.

Conclusions

The three grains considered present a contrasting picture. In all three cases, there are two recognized world prices one of which is generated in the Chicago futures markets. In each

Table 7.9: Estimated rice VAR					
	$\Delta\ln Pbk_t$	$\Delta\ln Pvn_t$	$\Delta\ln Pky_t$	$\Delta\ln Pmw_t$	$\Delta\ln Ppy_t$
$\Delta\ln Pbk_{t-1}$	0.516 (0.117)	0.419 (0.069	0.156 (0.040)		
$\Delta\ln Pvn_{t-1}$	- 0.138 (0.194)	0.016 (0.113)		0.003 (0.094)	0.400 (0.181)
$\Delta\ln Pky_{t-1}$			-0.083 (0.126)		
$\Delta\ln Pmw_{t-1}$				0.485 (0.094)	
$\Delta\ln Ppy_{t-1}$					0.055 (0.123)
$\Delta\ln Pbk_{t-2}$	- 0.268 (0.070)				
$\Delta\ln Pvn_{t-2}$	0.268 (0.070)		0.124 (0.044)		0.073 (0.052)
$\Delta\ln Pky_{t-2}$			- 0.180 (0.061)	0.185 (0.044)	
$\Delta\ln Pmw_{t-2}$				- 0.176 (0.036)	
$\Delta\ln Ppy_{t-2}$					- 0.100 (0.050)
R2	0.410	0.461	0.337	0.560	0.158
Standard error	7.15%	4.24%	2.55%	3.84%	7.15%
Autoregression LM (4)	$F_{4,49}$ = 1.27 [29.5%]	$F_{4,51}$ = 0.80 [53.1%]	$F_{4,49}$ = 2.67 [4.3%]	$F_{4,49}$ = 2.21 [8.1%]	$F_{4,49}$ = 1.78 [14.9%]

case, the two prices relate to different specifications - yellow versus white maize, soft versus hard wheat and (milled) white versus (unmilled) rough rice. The prices therefore differ both in terms of level and monthly changes. However, while the two maize prices are cointegrated and therefore tend back towards a time invariant proportion, there is no clear long term relationship between the two world prices in either the rice or the wheat markets. Furthermore, while maize prices do appear to be set on the major international world maize markets, rice prices appear to be determined in a decentralized manner in rice producing and consuming countries.

Standard pass-through models work well for maize. Kenya is seen as having largely insulated itself from changes in world maize prices, but, despite its efforts, Malawi has failed to do so.[18] The same pass-through models work poorly for wheat and rice. In the case of wheat, the various national prices move together but do not move with the two exchange

18 Daviron (2008) stresses "*une dynamique propre et endogène des marchés des céréales locales*" which results in a "*forte instabilité*". The clear local dynamic is evident in our results for Benin, Kenya and Malawi, but the high volatility only in Malawi.

prices. In the case of rice, transmission is largely in the opposite direction from that supposed in those analyses.

These findings are corroborated by estimated VAR models which reflect the cointegration structure in the variables. Pass through is high and relatively rapid in maize, mixed and complicated in rice and very low in wheat. On the basis of the evidence analysed in this chapter, the world maize market appears integrated while the wheat market does not. The world price is relevant to national prices in maize but does not appear so in wheat. Rice shows evidence of a high degree integration across national markets but in which the supposed world prices play little role. The important prices appear to be the transactions prices of major exporters, such as Viet Nam, and not exchange spot prices.

In policy terms, this analysis indicates caution in interpreting the implications of movements in so-called world prices in relation to the prices paid by consumers and received by farmers in developing countries. We have seen that the extent and speed of pass-through varies both over grains and over countries. The world maize market appears the best integrated of the three markets considered and transmission was both high and rapid for Malawi and Peru, but absent in Kenya where government has succeeded in stabilizing domestic prices, albeit at a high level. The wheat market appears the least well integrated of the three - local wheat and wheat flour prices followed the upward but not the downward movement in world prices (see on this issue Box 3.3 and Stigler & Tortora, 2011). The reasons for this, and the extent to which it was also true in other countries, deserves further examination. Rice is a puzzling intermediate case - the various national prices do appear to move together, both on the way up and the way down, but do not do so in relation to the standard (Bangkok) world reference price which is more volatile than national prices and tends to follow rather than lead these prices.

It follows that policies which directly address world markets, such as a world grain stockpile or actions to curtail supposedly excessive futures market activity, might hope to reduce the volatility of world maize prices, but the effects of these policies on national wheat and rice prices is more difficult to predict. In any case, such direct intervention may be unnecessary and perhaps over costly if countries are able to insulate their domestic prices from movements in world markets, as appears to have been the case with maize in Kenya. This suggests that it may be better to address food security and food safety nets on a country-by-country basis so that policies can be adapted to the severity of the food price problem in each country and to local conditions. Grand schemes attract political attention but, in the view of the author of this chapter, more is to be gained from hard work at the country level.

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