Discussion Paper: The impact of the US dollar on international agricultural prices

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The views expressed in this paper are those of the author and may not in any circumstances be regarded as stating an official position of Defra or the UK government.

Abstract

A strong US dollar and high international commodity prices have characterised much of 2022. To some, this juxtaposition is strange given the observed historical relationship between the two. This paper uses some simple regression models to provide insights and then utilises non-linear auto regressive distributed lag (NARDL) models to investigate the possibility of an asymmetric relationship between the US dollar and international food prices. A sizeable proportion of the historical changes in international food prices seem to be explained by changes in the value of the US dollar. Evidence for short-run asymmetry in the relationship is found. In the long-run, such asymmetry is rejected by two of the three regression models utilised and the "unit-elastic" association between the US dollar and international prices cannot be rejected at the aggregate 'food price' level. Across individual agricultural products there are however notable differences. The provisional econometric modelling suggests that (i) a stronger US dollar in 2022 put some significant downward pressure on international food prices and (ii) allowing for asymmetry can be important in capturing economic relationships.

Keywords: agricultural prices, international markets, econometrics

JEL code: Agriculture Q11: Prices; Agriculture Q17: Agriculture in International Trade

Introduction

A strong dollar and high commodity prices have characterised much of 2022. A strong USD in combination with high prices of commodities has put pressure on import prices in many countries, and caused particular issues in emerging economies and developing countries recovering from the pandemic. High commodity prices, for energy and food, have been put forward as causal factors in rising inflation at both producer and consumer level in developed countries.

To some the juxtaposition of a strong dollar and high commodity prices is strange as it has long been 'known' that there exists an inverse relationship between the US dollar and the prices of internationally traded commodities. Simple correlations between the US dollar and the prices of some commodities, including agricultural goods, have seemingly broken down in the recent past and this 'stylized fact' has been highlighted by some international organisations (e.g., IMF). Whilst the dollar may have been strong and international commodity prices high, this does not preclude the fact that had the dollar not strengthened then international commodity prices may have been higher.

There is relatively limited recent quantitative work on the impact of the US dollar on the international prices of a broad spectrum of individual agricultural products, and few studies which examine potential for asymmetry in this relationship. This paper attempts to address some of that gap and uses a non-linear auto regressive distributed lag (NARDL) model to investigate the relationship between the US dollar and international agricultural prices, both at aggregate index level and individual product level. As a discussion paper, the analysis is mostly descriptive and the econometric findings are both provisional and grounds for discussion. It is hoped the work provides some additional insights into the drivers of international price changes which have been the subject of much attention and policy interest. The does not provide a full overview of all the various issues to do with the topic.

Nothing in the analysis in this paper assumes the US dollar is necessarily the sole causal factor in the relationship between the dollar and international agricultural commodity prices, rather it is more concerned with *association* between these variables. It is naturally the case that the US dollar exchange is partly collinear with sentiment and global macroeconomic conditions and that both exchange rates and (storable) commodity prices reflect *expectations* about future market conditions. These factors may be driving the econometric relationships identified in the paper but such conditions are difficult if not impossible to capture accurately in real time and include in regression models.

International Commodity Prices and the US dollar

Most internationally traded commodities, including agricultural and food products, are traded in US dollars and international reference prices are quoted in dollars. As a result, there is a relationship between the value of the US dollar and international prices. For instance, an appreciation of the dollar increases international prices faced by holders of foreign currency, reducing import demand and raising export supplies, acting to lower prices. In a stylized world with no frictions and floating exchange rates, then one can sometimes think of this as simply a 'measuring rod' effect. If the US dollar is worth more and there are no changes in the underlying market fundamentals, prices measured in now stronger dollars should fall and the relationship should be close to "unit elastic" i.e., a 1% appreciation (depreciation) in the US dollar should lead to a fall (rise) of 1% in the international dollar price. Of course, there are frictions in the real world and departures from this stylized world. Many countries operate on fixed or at least weakly pegged exchange rates and there are agreements such as contracts which might stymie adjustment in reality.

Obstfeld (2022) observes that there appears to be a larger than unity correlation between the dollar and international commodity prices (energy, metals, food and raw materials) in aggregate. In other words, when the dollar appreciates then commodity prices in dollars fall by a greater percentage than the percentage change in the dollar (and vice-versa for depreciation of the US dollar). In the literature this is referred to an association which is greater than 'unit elastic' and is sometimes attributed to the collinearity of the dollar with global macroeconomic conditions and/or *expectations* of those conditions. Importantly, observations such as these are made during times of large swings in the value of the dollar which are likely representative of broader macroeconomic shocks, which likely impact on commodity prices too.

Figure 1a shows the evolution of the monthly broad US nominal dollar index and the World Bank's international Food Price Index over the period from 1994 to 2022. It should be noted that the y-axis are different and the inverse US dollar is shown i.e. a rise in the series corresponds to deprecation of the dollar. The reason for doing this is to make it easier to see the correlation. Figure 1b shows the monthly percentage in the World Bank's international Food Price Index versus the monthly percentage in the broad US nominal dollar over the period 1994-2022.

Figure 1a. World Bank Food Price Index & the USD , nominal terms.

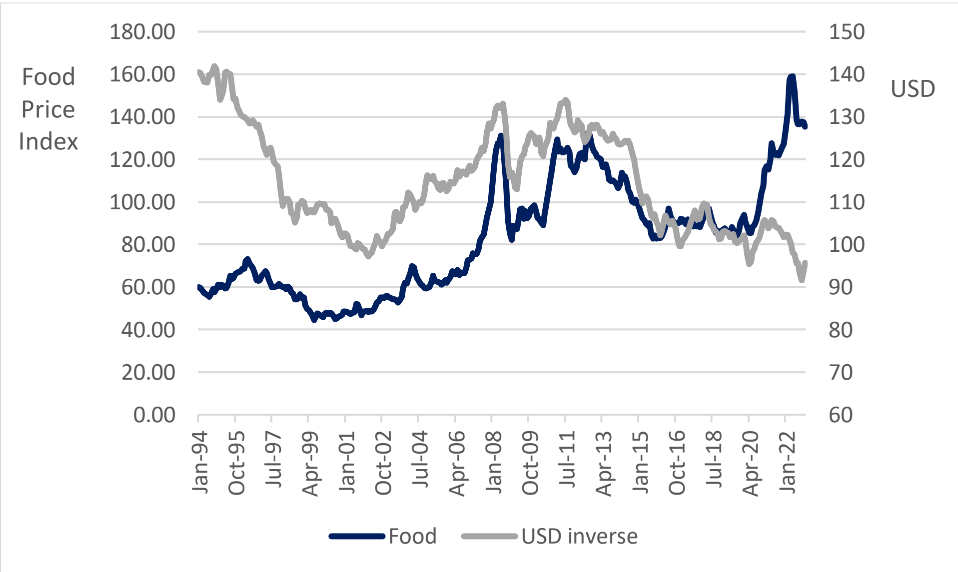
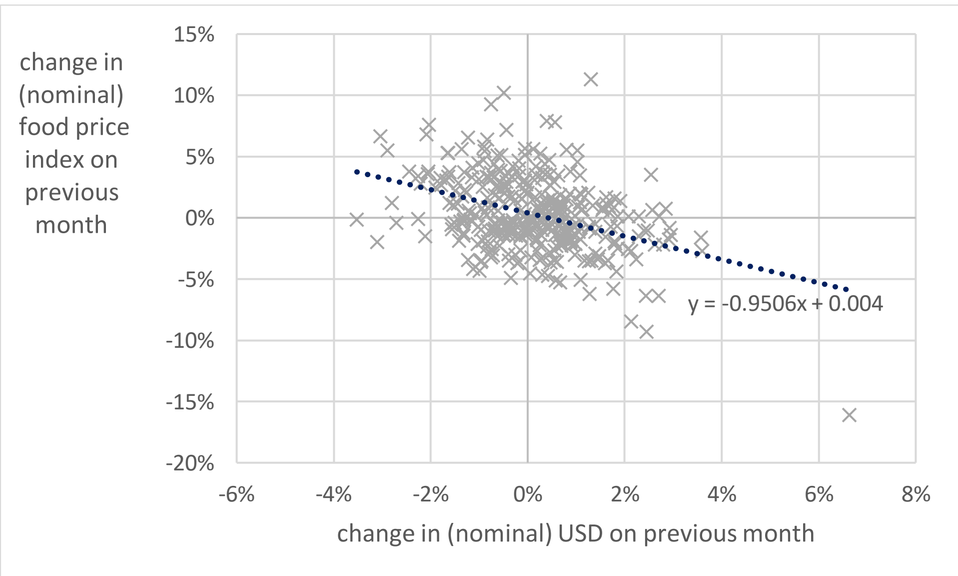


Figure 1b. Monthly % changes in World Bank Food Price Index & nominal USD



A visual inspection of the data suggests a negative correlation between the strength of the dollar and international food prices over the period, though the raw correlation is lower in recent years than in the past. Nothing is new here. There is a significant literature on the relationship between the US dollar and international commodity prices, dating back to Shuh (1974). It is not the purpose of this discussion to provide an overview of that literature. Nonetheless, much of the recent literature in the last few decades has focussed on the relationships between the US dollar exchange rate, other macroeconomic variables (e.g. oil) and commodity prices. Many of these studies utilise Vector Auto-Regression (VAR) and Vector Error Correction Models (VECM). These are very useful models which can capture the endogeneity of exchange rates, commodity prices and macroeconomic variables to one another. However, both VAR and VECM techniques used in applied work typically focus on (a) broad commodity price indices; (b) impose linear or symmetric relationships between the variables considered in the models and/or (c) tend to not include contemporaneous changes in the regressors. In this paper an alternative approach is used which allows us to examine the issue of potential asymmetry in the relationship between the US dollar and international food prices and gives us flexibility on the lag structure so that we can include contemporaneous as well as lagged effects easily.

Data & Econometric specification

International prices of agricultural commodities are taken from the World Bank's Pink Sheet, US broad dollar indices and producer price indices come from the Federal Reserve Board. The estimation period covers monthly data from 1994 to 2021, when we end the sample to allow for 2022 out-of-sample forecasts.

Price data are often characterised by non-stationarity i.e., their mean and variance changes over time. This feature poses problem for regression analysis and can lead to spurious regression. Augmented Dickey-Fuller (ADF) tests are performed on the price and exchange rate series are shown in Figure 2 to help establish the properties of the data. For all series, the logarithmic price is non-stationary, but is stationary in first differences.

Figure 2. ADF Tests[1] (null hypothesis : unit root)

ADF Tests	**Nominal**		**Real**	
	constant	constant + trend	constant	constant + trend
ln (food)	p-value: 0.7537	p-value: 0.5425	p-value 0.1492	p-value 0.2855
ln (USD)	p-value 0.3484	p-value 0.6952	p-value 0.5818	p-value 0.849
ln (PPI)	p-value 0.9037	p-value 0.4789	NA	NA
ADF Tests	**Nominal**		**Real**	
	constant	constant + trend	constant	constant + trend

[1] The lag selection is tested down from 12 lags using the AIC criterion.

Δ ln (food)	p-value <0.001	p-value <0.001	p-value <0.001	p-value <0.001
Δ ln (USD)	p-value <0.001	p-value <0.001	p-value <0.001	p-value <0.001
Δ(PPI)	p-value <0.001	p-value <0.001	NA	NA

Before proceeding to the NARDL modelling, a simple regression model is run in first differences. These regressions are conceptually similar to fitting a line through a scatter plot of the % change in nominal international food prices and the % change in the nominal US dollar. The first simple model (S1) estimates the change in the international price ("Price") on time (represented by the intercept) and the change in the US dollar index. The second simple model (S2) does the same thing but includes an auto-regressive term, modelling the change in the international price as a function of time, the change in US exchange rate and the previous month's change in the international food price index[2]. Regressions are run for the World Bank Food Price Index and selected individual agricultural products from the World Bank's pink sheet. Formally, the regression models are as follows:

Model S1: $\Delta \ln Price_t = \beta_0 + \beta_1 \Delta \ln USD_t + \varepsilon_t$

Model S2: $\Delta \ln Price_t = \beta_0 + \beta_1 \Delta \ln USD_t + \beta_2 \Delta \ln Price_{t-1} + \varepsilon_t$

These parsimonious regressions are provided as context for the results later in the paper and whilst useful exercises in themselves, they suffer from the omission of any long-run dynamics. The results are therefore shown in Figure 3 for the short-run coefficient β_1 for both models S1 and S2 . Both models are run firstly with nominal international prices and the nominal dollar exchange rate, and secondly using real prices and the real dollar exchange rate. Qualitatively, the short-run effects are similar between the two reflecting the fact that over most of the sample period inflation has been rather low. Clearly this has not been the case in 2022.

Figure 3. Simple estimates of short-run relationship between USD and international food prices & selected individual products (1994:01 to 2021:12).

Commodity	Estimate of β_1			
	Nominal prices & Nominal USD		Real prices & Real USD	
	Model S1	Model S2		
Food Price Index	**-1.01***** (0.23)	**-0.86***** (0.17)	**-0.70***** (0.163)	**-0.61***** (0.13)
wheat	**-1.19***** (0.31)	**-1.05***** (0.30)	**-0.98***** (0.29)	**-0.87***** (0.28)
maize	**-1.05**** (0.37)	**-0.89***** (0.34)	**-0.70**** (0.30)	**-0.62**** (0.27)
rice	**-0.80***** (0.36)	**-0.64***** (0.34)	**-0.58*** (0.34)	**-0.52*** (0.26)
sugar	**-1.12*****	**-1.04*****	**-0.83****	**-0.79****

[2] One technical reason to do this is to capture inherent serial correlation in the regression residuals, and the number of lags of the dependent variable is increased in some regressions to ensure no residual serial correlation is present.

	(0.37)	(0.26)	(0.36)	(0.32)
soymeal	**-0.99***** (0.29)	**-0.85***** (0.23)	**-0.71***** (0.26)	**-0.64***** (0.21)
soy oil	**-1.69***** (0.34)	**-1.47***** (0.27)	**-1.34***** (0.28)	**-1.17***** (0.24)
beef	-0.38 (0.28)	-0.24 (0.24)	0.003 (0.20)	0.08 (0.18)
lamb	**-0.82***** (0.17)	**-0.72***** (0.14)	**-0.57***** (0.17)	-0.49*** (0.14)
chicken	-0.29 (0.27)	-0.19 (0.28)	0.17 (0.23)	0.22 (0.18)

The R-squared metrics for these types of models are in the range of 0.05 to 0.2, and they are not very informative as we do not expect changes in the US dollar to explain large elements of the change in international food prices from month to month over this length of time. Weather and growing conditions as well as changes in expectations for supplies have strong impacts in these markets.

Despite their problems the simple regressions in first differences suggest a strong negative association between changes in the US dollar and international food prices, which remains even when a lagged dependent variable is included in the regression specification. The standard errors of the estimates are rather wide so whilst the results are highly statistically significant – in the sense of being different to zero – they are imprecisely estimated. It should also be noted that whilst many of these coefficient estimates are close to a unit elastic association in magnitude, for some commodities like beef or chicken they are much lower and not statistically significant. Clearly, there are likely to be commodity-specific differences in the relationship which in part will depend on the homogeneity of the product in question and how well integrated different markets are with one another. In the case of beef and chicken, the international price itself is a looser concept than in, say, sugar or maize. The timeframe of the sample is also relevant and going back to 1994 in order to gain observations may mean not accounting for commodity-specific structural breaks.

From the S2 model, a "long-run multiplier" can be calculated which gives the long-term association between the change in the dollar and international prices. From the regression results using real prices, this long-run multiplier cannot be rejected to equal unity in all the regressions aside from soy oil (where it is greater than unity) and beef/chicken where it is not significant.

The results from the simple models will not be discussed further in this paper but provide some context for the next section.

NARDL modelling

The Non-Linear Auto-Regressive Distributed Lag Model (NARDL) is chosen as the appropriate regression specification. The NARDL model (Shin, Yu, and Greenwood-Nimmo, 2014) is a representation of an error correction model and the 'Bounds Testing' procedure (Pesaran, Shin & Smith 2001) is used to assess the evidence for co-integration between the USD and prices. A NARDL approach is used here to test for the possibility of an asymmetric relationship between the USD and international food prices and does so by incorporating the partial sums of the positive and negative changes in the regressors, here the US dollar exchange rate. This allows us to test for non-linear associations between international food prices and the dollar, depending on whether the dollar is appreciating or depreciating.

To test the sensitivity of the results to alternative specification, 3 different regression models are used; two using nominal international prices and the nominal USD dollar index and one using real international prices and the real USD exchange rate index. With respect to the latter, the choice of the appropriate deflator for international commodity prices is somewhat debatable. For monthly data, the options are usually the US CPI and the US PPI. Here, the US PPI is used to deflate the price indices since we are dealing with food commodity prices which are at wholesale rather than consumer level.

Other control variables were considered including global monthly trade volumes, crude oil prices and stock prices but none of these added very much to the regression analysis and they were statistically insignificant when included alongside the auto-regressive terms, lags and existing variables in these equations. The same holds for seasonal dummies which were omitted by the Wald test. More investigation of these variables and lag structure is often required but these results provide some initial grounds for discussion.

The 3 regression models are as follows. Each model essentially regresses the World Bank International Food Price index on the partial sums of the increases in the US dollar exchange rate (i.e. appreciations) and the partial sums of the decreases in the US dollar exchange rate (i.e. depreciations). The short-run relationships are captured by the first differenced variables and the long-run relationships by the variables included in levels, which makes the NARDL model look like a transformed version of an unrestricted error correction model (Shin, Yu, and Greenwood-Nimmo, 2014).

<u>Regression for nominal prices (Models “N1” and “N2”)</u>

Model N1:

$$\Delta \ln Price_t = \beta_0 + \sum_{t=0}^{t=k} \beta_{1k}\, \Delta \ln USD_{t-k}^{+ve} + \sum_{t=0}^{t=k} \beta_{2k}\, \Delta \ln USD_{t-k}^{-ve} + \\ + \sum_{t=1}^{t=k} \beta_{3k}\, \Delta \ln Price_{t-k} + \theta_1 \ln Price_{t-1} + \theta_2 \ln USD_{t-1}^{+ve} + \theta_3 \ln USD_{t-1}^{-ve} + \varepsilon_t$$

Model N2:

$$\Delta \ln Price_t = \beta_0 + \sum_{t=0}^{t=k} \beta_{1k}\, \Delta \ln USD_{t-k}^{+ve} + \sum_{t=0}^{t=k} \beta_{2k}\, \Delta \ln USD_{t-k}^{-ve} + \\ \sum_{t=0}^{t=k} \beta_{3k}\, \Delta \ln PPI_{t-k} + \sum_{t=1}^{t=k} \beta_{4k}\, \Delta \ln Price_{t-k} + \theta_1 \ln Price_{t-1} + \theta_2 \ln USD_{t-1}^{+ve} + \\ \theta_3 \ln USD_{t-1}^{-ve} + \theta_4 \ln PPI_{t-1} + \varepsilon_t$$

<u>Regression for real prices (Model “R3”)</u>

Model R3:

$$\Delta \ln rPrice_t = \beta_0 + \sum_{t=0}^{t=k} \beta_{1k}\, \Delta \ln rUSD_{t-k}^{+ve} + \sum_{t=0}^{t=k} \beta_{2k}\, \Delta \ln rUSD_{t-k}^{-ve} + \\ \sum_{t=1}^{t=k} \beta_{3k}\, \Delta \ln rPrice_{t-k} + \theta_1 \ln Price_{t-1} + \theta_2 \ln USD_{t-1}^{+ve} + \theta_3 \ln USD_{t-1}^{-ve} + \varepsilon_t$$

Where :

Price : World Bank International Food Price Index in nominal terms

USD^{+ve} comprises the partial sum of positive changes in the broad US dollar exchange rate

USD^{-ve} comprises the partial sum of negative changes in the dollar exchange rate

rPrice : World Bank International Food Price Index in real terms

rUSD^{+ve} comprises the partial sum of positive changes in broad US dollar real exchange rate

rUSD^{-ve} comprises the partial sum of negative changes in the broad US real dollar exchange rate

These three different regression models which are run for the World Bank's Food Price Index and the following statistical tests are performed on the regressions to assess the evidence for short-run symmetry, long-run symmetry and a unit-elastic response.

I. "Short-run" symmetry test: the lags on the partial sums should be the same & coefficients on them not statistically different.
II. Short-run unit elastic test: $\beta_{1k} = -1$ & $\beta_{2k} = -1$
III. Bounds co-integration test: $H_0 : \theta_1 = \theta_2 = \theta_3 = 0$
IV. Long run symmetry test: : $H_0 : \theta_2 = \theta_3$
V. Unit-elastic response: $H_0 : \frac{\theta_2}{\theta_1} = 1$ & $\frac{\theta_3}{\theta_1} = 1$

Results

The econometric estimates from the three models for the Food Price Index are shown in Figure 4. Details of the regressions are provided in the Annex to this discussion paper. Both agricultural commodity prices and exchange rates are noisy series and the main area of focus is on the tests of asymmetry and whether a unit-elastic association can be rejected or not. The short-run coefficients for changes in the dollar are also provided in Figure 4 to aid understanding of the results and are probably the most relevant to policymakers concerned with drivers of changes in international food prices.

All 3 models find evidence of co-integration between the US dollar and international food prices using the 'Bounds Test' and the model's residuals are neither auto-correlated nor heteroskedastic using standard tests. Some details of the regressions are reproduced at Annex A.

Figure 4. International Food Price Index Regression Results

	short-run USD appreciation	short-run USD depreciation	short-run symmetry	short-run unit elastic test	long-run symmetry test	long run unit-elastic test
Model N1	L0: -1.29*** (0.25)	L3: 0.46*** (0.19)	reject	cannot reject-appreciation; reject-depreciation	reject	reject-appreciation; reject-depreciation
Model N2	L0: -1.04*** (0.20)	L3: 0.45*** (0.19)	reject	cannot reject-appreciation; reject-depreciation	cannot reject	cannot reject-appreciation, cannot reject - depreciation

Model R3	LO: -0.98*** (0.21)	L3: 0.45** (0.18)	reject	cannot reject- appreciation; reject- depreciation	cannot reject	cannot reject- appreciation, cannot reject - depreciation

NB: L refers to the lag length i.e. L0 is contemporaneous, L1 is lagged one period etc. ***indicates statistically significant at 99% level. Standard errors are in parentheses.

Model Diagnostics	LM test for autocorrelation (null hypothesis: autocorrelation not present)	White's test for heteroskedasticity (null hypothesis: heteroskedasticity not present)	R-squared
Model N1	p-value : 0.471	p-value: 0.298	0.30
Model N2	p-value: 0.624	p-value: 0.118	0.36
Model R3	p-value: 0.223	p-value: 0.430	0.21

As with all econometric analysis, the results should be interpreted rather cautiously. The main findings from these three regression models suggest that asymmetry in the relationship between the US dollar and international food prices is potentially notable:

- In the short-run, in all three models, dollar appreciation is found to be associated with a more immediate and larger change in international food prices than dollar depreciation. More specifically, dollar appreciation is associated contemporaneously with a fall in international food prices which is close to unity in its relation to the change in the US dollar. Conversely, depreciation takes 3 months to be associated with a statistically significant rise in international food prices and the relationship is below unity in relation to the change in the US dollar (the point estimate suggests around half as large as the impact of appreciation).

- In the long-run, two out of the three regression models are unable to reject (i) symmetry between dollar appreciation and depreciation, and (ii) an association which is unit-elastic. Speed of adjustment to this long-run relationship is slow in these models. The proper interpretation of this finding is consistent with what we would expect in agricultural commodity markets; there are plenty of shocks (e.g. weather, trade shocks, policy shocks etc.) to knock the underlying equilibrium relationship off course and these are rather persistent.

Naturally the results described here may be sensitive to the time considered in the sample and there may be structural breaks in the relationship between the US dollar and international food prices. Initial testing for structural breaks does not seem to suggest that the regression results would be different if a sub-sample were used rather than the whole sample. For example splitting the dataset into either side of the 2007/8 spike in agricultural prices and performing a Chow test does not indicate a structural break in the relationship around that time. It is obvious that further testing should be done to establish whether there are other structural breaks.

NARDL models for individual agricultural products prices and sub-indices have also been estimated and the results are still being analysed at the time of writing. Significant differences across commodities have been provisionally found and the qualitative findings are summarised in the table below. In general, the findings suggest asymmetries are more prevalent in staple grains (including rice) and symmetry is observed more in processed products.

Asymmetry in association with USD	Symmetry in association with USD
Wheat, maize, sugar, rice, palm oil	Soybeans, soymeal, soybean oil, lamb , beef

Discussion & potential reasons for asymmetry

There are a number of potential reasons for the asymmetry identified by the NARDL approach. These include, but are certainly not limited to, the following:

- **Contracts and the realities of agricultural trade**. Many agricultural products are traded on contract, in advance of delivery dates. In the short-run, it may be the case that is easier to increase supply on global markets through new contracts or spot sales than it is to reduce demand as buyers will to some extent already be 'locked in' in to purchases in the near term. If this is the case, it could help explain why an appreciation has a more immediate impact on international prices than depreciation.

- **Asymmetry in the demand curve.** It is well-known that many demand curves exhibit non-linearity which can be attributed to things like the holding of stocks. This can make demand more elastic to price falls then to price rises and may play a role in the asymmetry we observe in staple grain markets.

On a more technical level, the analysis also highlights that both VAR and VECM models sometimes used to estimate relationships between exchange rates, other macroeconomic variables and agricultural commodity prices might wish to consider the possibility of asymmetry in response to the exchange rate in their specifications. A comparison of these results with VAR and VECM models is one of the future extensions to this work.

The analysis contained in this paper is rudimentary and as such suffers from some notable limitations. There are plenty of potential omitted variables, which auto-regressive models do not necessarily make less of a concern. Establishing causality from the dollar to international food prices is not straightforward given the fact that both (storable) commodity prices and exchange rates can reflect similar information about market expectations of the future economic environment.

Applying the econometric models to 2022 price movements

The econometric models can be used to give a rough feel for the potential role of the stronger dollar during 2022 in international food price changes. As a central estimate these results suggest that the dollar strengthening has played a significant and increasing role over 2022. According to the central estimates, had the dollar not strengthened then international food prices would have been around 10% higher by the end of 2022. These magnitudes are noteworthy. Given that the World Bank's international food price index at the end of 2022 had fallen by around 15%, in nominal terms, since peaking in May then perhaps a sizeable proportion of the fall in prices since the Spring may have been due to the stronger US dollar. This is in line with commentary on market developments from the international organisations who have pointed to the US dollar as one of the reasons for falling international food prices. There is of course much uncertainty in such forecasting exercises and the forecast error range is wide. Individual food commodity prices may have been more or less affected than the food price index in aggregate.

Implications

The analysis demonstrates that global exchange rate fluctuations, when broad, can have implications not only for import prices in a given country but also international agricultural prices. Calculations of changes in import prices due to exchange rate movements are often made by officials and commentators but such calculations do not usually take this endogeneity into account.

Whilst a weaker USD dollar in future and hence stronger foreign currencies may reduce pressure on import prices in some countries, it is possible that some of this 'gain' would be offset by upward pressure on dollar prices of international agricultural products if exchange rate movements are part of a broad decline in the US dollar.

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<u>Annex A:</u>

Model N1: OLS, using observations 1994:06-2021:12 (T = 331)
Dependent variable: d_l_Food
HAC standard errors

	Coefficient	*Std. Error*	*t-ratio*	*p-value*	
const	0.230572	0.0627658	3.674	0.0003	***
d_pos_USD	−1.29279	0.251411	−5.142	<0.0001	***
d_neg_USD_3	−0.463330	0.190749	−2.429	0.0157	**
l_Food_1	−0.0530192	0.0149584	−3.544	0.0005	***
pos_USD_1	−0.0812937	0.0228297	−3.561	0.0004	***
neg_USD_1	−0.115262	0.0314787	−3.662	0.0003	***
d_l_Food_1	0.328338	0.0462328	7.102	<0.0001	***

d_l_Food_2	0.0282416	0.0429499	0.6575	0.5113	

Mean dependent var	0.002448	S.D. dependent var	0.030304
Sum squared resid	0.206958	S.E. of regression	0.025313
R-squared	0.317103	Adjusted R-squared	0.302303
F(7, 323)	16.08045	P-value(F)	3.89e-18
Log-likelihood	751.2840	Akaike criterion	−1486.568
Schwarz criterion	−1456.151	Hannan-Quinn	−1474.436
rho	0.021730	Durbin's h	0.730983

LM test for autocorrelation up to order 12 -
Null hypothesis: no autocorrelation
Test statistic: LMF = 0.976404
with p-value = P(F(12, 311) > 0.976404) = 0.471283

White's test for heteroskedasticity -
Null hypothesis: heteroskedasticity not present
Test statistic: LM = 38.9072
with p-value = P(Chi-square(35) > 38.9072) = 0.298143

Model N2: OLS, using observations 1994:06-2021:12 (T = 331)
Dependent variable: d_l_Food
HAC standard errors, bandwidth

	Coefficient	*Std. Error*	*t-ratio*	*p-value*	
const	0.146468	0.167217	0.8759	0.3817	
d_pos_USD	−1.04995	0.203117	−5.169	<0.0001	***
d_neg_USD_3	−0.452022	0.189917	−2.380	0.0179	**
l_Food_1	−0.0559436	0.0135690	−4.123	<0.0001	***
pos_USD_1	−0.0800477	0.0221423	−3.615	0.0003	***
neg_USD_1	−0.107389	0.0343567	−3.126	0.0019	***
d_l_PPI	0.672136	0.167844	4.005	<0.0001	***
d_l_PPI_1	−0.643579	0.164078	−3.922	0.0001	***
l_PPI_1	0.0195551	0.0350821	0.5574	0.5776	
d_l_Food_1	0.333127	0.0482675	6.902	<0.0001	***
d_l_Food_2	0.0563095	0.0436152	1.291	0.1976	

Mean dependent var	0.002448	S.D. dependent var	0.030304
Sum squared resid	0.188921	S.E. of regression	0.024298
R-squared	0.376620	Adjusted R-squared	0.357140
F(10, 320)	14.75557	P-value(F)	1.36e-21
Log-likelihood	766.3758	Akaike criterion	−1510.752
Schwarz criterion	−1468.928	Hannan-Quinn	−1494.071
rho	0.013409	Durbin's h	0.509966

LM test for autocorrelation up to order 12 -
Null hypothesis: no autocorrelation
Test statistic: LMF = 0.825532
with p-value = P(F(12, 308) > 0.825532) = 0.624059

White's test for heteroskedasticity -
Null hypothesis: heteroskedasticity not present
Test statistic: LM = 79.0622
with p-value = P(Chi-square(65) > 79.0622) = 0.112814

Model R3: OLS, using observations 1994:07-2021:12 (T = 330)

Dependent variable: d_l_rFood
HAC standard errors, bandwidth 5 (Bartlett kernel)

	Coefficient	*Std. Error*	*t-ratio*	*p-value*	
const	0.204703	0.0527695	3.879	0.0001	***
d_pos_USD	−0.983169	0.207106	−4.747	<0.0001	***
d_neg_USD_3	−0.451226	0.178186	−2.532	0.0118	**
l_rFood_1	−0.0528628	0.0137690	−3.839	0.0001	***
pos_USD_1	−0.0636794	0.0236220	−2.696	0.0074	***
neg_USD_1	−0.0699080	0.0248172	−2.817	0.0051	***
d_l_rFood_1	0.328505	0.0470816	6.977	<0.0001	***

Mean dependent var	0.000371	S.D. dependent var	0.027675
Sum squared resid	0.195398	S.E. of regression	0.024596
R-squared	0.224583	Adjusted R-squared	0.210179
F(6, 323)	13.61557	P-value(F)	8.71e-14
Log-likelihood	757.9989	Akaike criterion	−1501.998
Schwarz criterion	−1475.404	Hannan-Quinn	−1491.390
rho	0.020656	Durbin's h	0.724171

LM test for autocorrelation up to order 12 -
Null hypothesis: no autocorrelation
Test statistic: LMF = 1.28852
with p-value = P(F(12, 311) > 1.28852) = 0.223837

White's test for heteroskedasticity -
Null hypothesis: heteroskedasticity not present
Test statistic: LM = 27.6214 with p-value = P(Chi-square(27) > 27.6214) = 0.430688