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ARE GLOBAL FOOD PRICES BECOMING MORE VOLATILE AND MORE PERSISTENT?*

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Abstract

Global food prices have recently exhibited highly volatile and very persistent dynamics. Greater fluctuations in food commodities than manufacturing goods often put more serious hardship to poor countries that tend to specialize in raw commodity industries. The present paper attempts to identify the factors that help explain recent phenomena in world commodity markets. We first document strong dynamic correlations (Engle, 2002) between food commodity prices and the US dollar exchange rate. Employing the PANIC method (Bai and Ng, 2004), we then estimate a latent common factor from 27 food and beverage commodity prices, which seem to be closely related with the exchange rate. Once controlled for the effect of the common latent factor, idiosyncratic components of food commodity prices show substantially lower volatilities and persistence. Recent trends in global food commodity prices, therefore, seem to be well explained by a fairly simple but influential factor, that is, highly volatile and persistent movements of the US exchange rate since the recent financial crisis. Other than that, we do not see any compelling evidence of higher volatility or greater persistence in price dynamics. Our findings also call for special attention on the importance of financial markets in addition to factors that influence economic fundamentals.

Keywords: Global Food Prices; Volatility; Persistence; US Dollar Exchange Rate; DCC; PANIC

JEL Classification: C51; F31; Q02

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I Introduction

Global commodity prices have recently exhibited strong and persistent movements. For example, Corn (maize) price increased from $158.16 per metric ton in December 2008 to $308.72 in December 2012, which is roughly a 70% price hike in 4 years. During the same period, Beef, Sugar, and Crude Oil (Dubai) prices increased by 56.4%, 52.8%, and 97.0%, respectively. It is also well-documented that highly volatile price dynamics were observed in most commodity markets.

Volatile and persistent commodity prices may generate terms of trade shocks that can result in macroeconomic instability, which eventually hinders long-run economic growth. Naturally, countries that have a high concentration in producing these raw commodities/materials would become more vulnerable to such adverse effects. See, among others, Aghion et al. (2009, 2005), Bleaney and Greenway (2001), Kose and Reizman (2001), Deaton (1999), and Mendoza (1997) for discussions.

For instance, greater price fluctuations in food commodity markets can put a serious challenge to poor countries if they specialize in agricultural products, which is often the case. As Jacks *et al*. (2009) documented using long-horizon data since 1700, primary commodities always show a greater price volatility than manufacturing goods. Based on the same logic, Koren and Tenreyro (2007) point out that poor countries tend to experience more frequent and severe fluctuations in the GDP growth rate. Such fluctuations in national incomes add further hardship on households in developing countries because they might find it difficult to smooth their expenditures due to insufficient social safety net. Capital markets in those countries are less likely to be accessible by consumers.

This paper empirically investigates underlying economic factors that have generated recent price dynamics in world food commodity markets, using 27 IMF

Primary Commodity Prices data from January 1980 to August 2013. We conjecture that these prices might be heavily influenced by changes in the US dollar exchange rate, because they are denominated in dollars. Given national prices, a depreciation of US dollars results in price hikes in world commodity markets.[1] This conjecture is plausible if we recall the extremely volatile and persistent movements of the US dollar exchange rate after the financial market melt-down in 2008 triggered by the bankruptcy of Lehman Brothers on September 15th.

In what follows, we provide some empirical evidence that support this view. For instance, we report strong and persistently negative DCC (Dynamic Conditional Correlation; Engle, 2002) estimates between the foreign exchange rate and food commodity prices.

Employing the PANIC (Panel Analysis of Non-stationary in the Idiosyncratic and Common Components; Bai and Ng, 2004) method, we then extract a common component (from 27 food and beverage commodity prices) that governs these prices as well as the factor loading of each commodity, which provides useful information on the dependency of these prices on the common factor.

Our econometric findings are roughly two-fold. First, our estimated common factor has a mirror image of the US dollar exchange rate, while all factor loading estimates are positive, which shows an inverse relation between food prices and the value of US dollars. Second, controlling for the effect of the common factor (or the exchange rate), the remaining idiosyncratic components of commodity prices exhibit a lot less volatile and less persistent dynamics.

[1] This argument critically hinges upon the validity of the law of one price (LOP) in commodity markets. Since seminal work of Isard (1977), some (among others, Ardeni, 1989; Engel and Rogers, 2001; Parsley and Wei, 2001; Goldberg and Verboven, 2005) find evidence against the LOP, while others (for instance, Goodwin, 1992; Michael *et al*., 1994; Obsfeld and Taylor, 1997; Lo and Zivot, 2001; Sarno *et al*., 2004), find evidence in favor of the LOP. Since we focus only on highly tradable world commodities, this assumption is more likely to be valid.

For example, the ADF (augmented Dickey-Fuller) test rejects the null of non-stationarity only for 5 commodity prices at the 5% significance level, while it rejects the null for 18 prices once the common factor is removed from the series. That is, adjustments of *de-factored* idiosyncratic components might be consistent with dynamic stability of equilibria by interactions between the demand and supply of commodities.[2] This implies that policy makers need to pay extra attention to changes in the foreign exchange (financial) market in addition to factors that influence economic fundamentals.

The present paper proceeds as follows. In Section II, we present data descriptions and some preliminary analyses for the persistence and the volatility observed recently in food commodity markets. Section III explains two sets of our main empirical models, the DCC and the PANIC methods, then present our major findings. We also provide a brief discussion on our findings. The last section offers our conclusions.

II Data and Preliminary Analysis

We use monthly observations of 27 food and beverage commodity prices and a nominal US dollar exchange rate. We obtain the commodity prices from the IMF Primary Commodity Prices website, while the nominal exchange rate is the trade-weighted US dollar index against a subset of major currencies (ID: TWEXMANL) from the Federal Reserve Bank of St. Louis Economic Research Database (FRED).[3] The sample period is January 1980 to August 2013. Table 1 provides detailed explanations for commodities.

As we can see in Figure 1, food and beverage commodity prices have recently exhibited sharp hikes and quite volatile and persistent movements. For example, the price of Soybeans increased from $318.81 in December 2008 to $534.79 in December 2012,

[2] See Chen et al. (forthcoming) for related discussions.

[3] Major currency index includes the Euro Area, Canada, Japan, United Kingdom, Switzerland, Australia, and Sweden. For details, see the Board of Governors website.

which is about a 51.7% increase in 4 years. During the same period, Barley and Corn prices also have increased by 77.7% and 66.9%, respectively. We also note that the US dollar exchange rate show similarly persistent and highly volatile dynamics. It is interesting to see that the US dollar has *depreciated* by 10.1% (80.86 to 73.10) during that period right after a sharp 10.6% *appreciation* from September 2008 to March 2009.

Figure 2 provides log differences of these commodity prices over a year to see how volatile those price changes are. These graphs clearly show a lot more volatile movements during the recent financial crisis triggered by the failure of Lehman Brothers on September 15, 2008. We also note higher volatilities in late 1980s from some commodity prices and the dollar exchange rate.

Insert Table 1 around here

Insert Figures 1 and 2 around here

To examine such volatility clustering effects in a univariate framework, we use a simple GARCH (1,1) model to estimate the time-varying volatility of these prices and the exchange rate. Let p_t be the log transformed price of a good i, then $x_{i,t} = \Delta p_{i,t} = p_{i,t} - p_{i,t-1}$ is the log return (growth rate) of the variable. With such notations, we have the following conventional GARCH (1,1) model.

$$h_{t+1} = \omega + \alpha(x_t - m_t)^2 + \beta h_t, \quad (1)$$

where h_t is the conditional variance of the residual of a regression $x_t = m_t + \sqrt{h_t}\varepsilon_t$. ε_t is a unit variance i.i.d. shock.

We report the GARCH volatility estimates for some commodity prices and the exchange rate in Figure 3. Estimated volatilities clearly exhibit clustering effects. Furthermore, these commodity prices and the exchange rate seem to share high volatility regimes, for instance, in late 1980s and in late 2000s. In what follows, we will examine such phenomena using more rigorous empirical methods.

Insert Figure 3 around here

III Empirical Analysis

1 Dynamic Conditional Correlation

We extend the previous univariate GARCH model to a multivariate framework to investigate the dynamic conditional correlation (DCC; Engle, 2002) between commodity price returns and the exchange rate appreciation rate.

Let $\boldsymbol{y}_t = [\Delta p_{i,t}, \Delta s_t]'$, where s_t is the log US dollar exchange rate as the price of one dollar in terms of a basket of foreign currencies. We first filter out the *expected* component of $\boldsymbol{y}_t$ by the following vector autoregressive process.

$$\boldsymbol{y}_t = \boldsymbol{\Phi}(L)\boldsymbol{y}_{t-1} + \boldsymbol{e}_t, \tag{2}$$

where $\boldsymbol{\Phi}(L)$ is a lag polynomial matrix. $\boldsymbol{e}_t$ obeys the following bivariate normal distribution.

$$\boldsymbol{e}_t|\Omega_{t-1} \sim N(\boldsymbol{0}, \boldsymbol{H}_t), \tag{3}$$

where Ω_{t-1} denotes the adaptive information set at time t. The conditional covariance matrix $\boldsymbol{H}_t$ is defined as follows.

$$\boldsymbol{H}_t = \boldsymbol{D}_t\boldsymbol{R}_t\boldsymbol{D}_t, \tag{4}$$

where $\boldsymbol{D}_t = diag(\sqrt{h_{i,i,t}})$ is the diagonal matrix with the conditional variances along the diagonal, and $\boldsymbol{R}_t$ is the time-varying correlation matrix.[4] The equation (4) can be re-parameterized as follows.

$$E_{t-1}\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t' = \boldsymbol{D}_t^{-1}\boldsymbol{H}_t\boldsymbol{D}_t^{-1} = \boldsymbol{R}_t = [\rho_{i,j,t}], \tag{5}$$

where $\boldsymbol{\varepsilon}_t = \boldsymbol{D}_t^{-1}\boldsymbol{e}_t$ is the standardized innovation. Engle (2002) proposes the following mean-reverting GARCH (1,1) type conditional correlations.

$$\rho_{i,j,t} = \frac{q_{i,j,t}}{\sqrt{q_{i,i,t}}\sqrt{q_{j,j,t}}}, \tag{6}$$
$$q_{i,j,t} = \bar{\rho}_{i,j}(1-\alpha-\beta) + \alpha\varepsilon_{1,t-1}\varepsilon_{2,t-1} + \beta q_{i,j,t-1},$$

where $\bar{\rho}_{i,j}$ is the unconditional correlation. We report the estimated conditional correlations ($q_{i,j,t}$) for the exchange rate return with each of six food commodity price returns in Figure 4.

There are two major findings. First, the correlations are overall negative with an exception of Corn, which makes sense because given national commodity prices, appreciations of the US dollar will decrease global commodity prices. Second, the estimated correlations seem to be time-varying, even though we do not have strong statistical evidence against the constant conditional correlation.[5] Especially, the dynamic conditional correlation became greater in absolute values recently since around the bankruptcy of Lehman Brothers when financial markets entered a long turbulent period.

These findings may imply that recent movements of commodity prices, for example, high volatility and persistence might have been largely influenced by

[4] Note that the constant conditional correlation (CCC; Bollerslev, 1990) is a special case of the DCC when $\boldsymbol{R}_t = \boldsymbol{R}$ for all t.
[5] All test results are available from author upon request.

turbulence and turmoil in financial markets during the sub-prime mortgage market crisis and the Great Recession.

Insert Figure 4 around here

2 The PANIC Method

(a) The Baseline Model

In previous sections, we noticed a close relationship between commodity prices and the US dollar exchange rate, which makes some sense because these world commodities are priced in the US dollar. That is, given national prices, a depreciation of US dollars tends to result in an increase in commodity prices and vice versa.

In order to investigate such conjectures for overall movements in global food prices, we attempt to estimate latent common factors that may govern the movement of these commodity prices. For this purpose, we assume that the log price $p_{i,t}$ has the following factor structure.

$$p_{i,t} = \boldsymbol{\lambda}_i' \boldsymbol{f}_t + e_{i,t}, \tag{7}$$

where $\boldsymbol{f}_t$ is a vector of the *common* factor components for $x_{i,t}$ across all commodities $i \in [1, N]$ at time t. $\boldsymbol{\lambda}_i$ denotes a vector of the commodity-specific factor loadings to its associated common factors, $\boldsymbol{f}_t$, while $e_{i,t}$ is a commodity i's *idiosyncratic* component. Since the log price is highly likely nonstationary, we transform (7) to the following equation with differenced variables.

$$\Delta p_{i,t} = \boldsymbol{\lambda}_i' \Delta \boldsymbol{f}_t + \Delta e_{i,t} \tag{8}$$

Then, we implement the principal component analysis (PCA) to estimate the latent factors and associated loadings from (8) after proper normalization.[6]

Note that we implement PCA for log-differenced series due to possible non-stationarity problems in the level variables. That is, we employ the PANIC method proposed by Bai and Ng (2004). After obtaining the loadings ($\hat{\boldsymbol{\lambda}}_i$) and differenced factors ($\Delta\hat{\boldsymbol{f}}_s$), we recover the common component factor for the level variable via cumulative summation of differenced series. That is,

$$\hat{\boldsymbol{f}}_t = \sum_{s=2}^{t} \Delta\hat{\boldsymbol{f}}_s \qquad (9)$$

Before estimating the factors, we implement an array of preliminary analyses. We first run unit-root tests for the level commodity prices and report the results in Table 2. Among 27 food and beverage commodity prices, the augmented Dickey-Fuller (ADF) test rejected the null of non-stationarity only for 5 time series at the 5% significance level, which implies that these series are highly persistent or even nonstationary.

Insert Table 2 around here

(b) Cross-Section Dependence

Next, we check whether there is sufficiently high degree cross-section dependence due to existence of common factors by employing a test by Pesaran (2004),

$$CD = \left(\frac{2T}{N(N-1)}\right)^{1/2} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{i,j}\right) \xrightarrow{d} N(0,1) \qquad (10)$$

where $\hat{\rho}_{i,j}$ is the pair-wise correlation coefficients from the residuals of the previous ADF regressions for commodity prices. The test rejects the null of no cross-section

[6] Normalization is required because the principal component analysis is not scale-invariant.

dependence at any conventional significance level (see Table 3). We also report overall high correlations between the food and beverage commodity prices. Naturally, in-group correlations are higher than those from all food prices. We also estimate optimal number of common factors by the method proposed by Bai and Ng (2002). All three information criteria propose a single common factor as can be seen in Figure 5.

Insert Table 3 around here

Insert Figure 5 around here

(c) Common Factor and Idiosyncratic Components

With a single common factor specification, we estimate a scalar common factor ($\hat{f}_t$) and factor loadings $\{\hat{\lambda}_i\}_{i=1}^{N}$, then obtain idiosyncratic components by taking residuals. That is,

$$\hat{e}_{i,t} = \sum_{s=2}^{t} \Delta\hat{e}_{i,s}, \qquad (11)$$

where $\Delta\hat{e}_{i,t} = \Delta p_{i,t} - \hat{\lambda}_i \Delta\hat{f}_t$.

The estimated common factor is reported in panel (a) of Figure 6, while factor loading estimates are provided in panel (b). Further, we evaluate the importance of the common factor in explaining the variations of the commodity prices relative to the idiosyncratic components by the following relative variation statistics.

$$rv_i = \frac{\sigma\left(\hat{\lambda}_i \hat{f}_t\right)}{\sigma\left(\hat{e}_{i,t}\right)}, \qquad (12)$$

where $\sigma(x)$ denotes the standard deviation of x.

As can be seen in Figure 6, the common factor seems to be a highly persistent stochastic process, which plays a critically important role in determining the commodity price dynamics. Especially for a majority of cereals and vegetable oils commodities, the common factor explains more than each idiosyncratic component, because $rv_i > 1$. For other commodities, the common factor often explains non-negligible portion of variations.

It should be also noted that factor loading estimates are all positive, which implies that the common factor has a qualitatively homogeneous effect on all 27 commodities. To see if the exchange rate has close relationship with this common factor, we displayed an inverted common factor with the exchange rate in Figure 7. We noticed a strong resemblance between the shapes of these two time series, especially for the latter half of the full sample since around the mid-1990s, while overall similar time trends are also observed in an earlier sample period.

Insert Figure 6 and 7 around here

(d) PANIC Unit-Root Test Results

We report panel unit-root test results by the PANIC in Table 4. We first report the univariate unit-root test for idiosyncratic components, which is the ADF test for the residual in (11). Note that the test is with no deterministic terms as explained in Bai and Ng (2004).

The ADF test rejects the null of non-stationarity for 18 time series at the 5% significance level. When we use the 10% level, the test rejects the null for 4 more series. For only 5 variables, Beef, Fishmeal, Fish, Cocoa beans, and Coffee, the ADF test fails to reject the null. Such results provide sharp contrast with the one in Table 2, where we obtained results in favor of stationarity for only 5 prices at the 5% level (3 more at the 10%). We also implement a panel unit-root test for these de-factored series, which rejects the null of non-stationarity at any significance level.

Following Bai and Ng (2004), we then implemented the ADF test with an intercept for the estimated latent common factor. The test fails to reject the null of non-stationarity even at the 20% significance level, which is not surprising because the estimated common factor exhibits very persistent movements as can be seen Figures 6 and 7.

Insert Table 4 around here

(e) Some Discussions

Our empirical analyses so far make it possible to conjecture the following. Consider a set of commodity prices that are governed by a single nonstationary common components plus stationary idiosyncratic components in each of those commodity markets. If that is the case, persistent movements of individual prices are generated largely by the single common factor. Therefore, controlling the effect of the common factor, the remaining dynamics must exhibit a much lower persistence, which may be consistent with a mean-reverting process.

This point is closely related with recent research work on commodity price dynamics. Global commodity prices are determined in world markets that equate the supply with demand for the product. Dynamic stability of equilibria in these markets may suggest, therefore, that these prices should exhibit mean reverting behavior, which is at odds with the data as we've seen in Table 2. However, since these commodities are priced in the US dollar and because the nominal exchange rate is nonstationary, it seems obvious that one would not find the evidence of stationarity for these prices unless one controls for the influence of the exchange rate. In other words, recent price hikes and highly volatile and persistent movements of food commodity prices are generated mostly by extreme changes of the nominal exchange rate caused by the recent financial system melt-down triggered by the US sub-prime mortgage market crisis.

We provide the common factor adjusted by factor loadings along with estimated idiosyncratic components for 6 food commodity prices in Figure 8. In comparison with highly persistent dynamics of the common factor, each idiosyncratic component exhibit quite stable and less persistent movements, which might be consistent with dynamic adjustment toward the equilibrium in each of these markets.

Insert Figure 8 around here

IV Concluding Remarks

The present paper empirically investigates factors that produced highly volatile and more persistent food price dynamics observed recently in world commodity markets. Greater fluctuations in food commodity markets than manufacturing goods market tend to make poor countries become more vulnerable to adverse effects of income shocks because they tend to specialize in agricultural industries. Lack of well-functioning credit and insurance markets may result in more suffering and hardship to households in poor countries who are not able to smooth their consumption in response to income shocks.

Since world commodities are traded in US dollars, depreciations of the dollar tend to increase commodity prices given national prices. We implement an array of econometric analyses that confirm close dynamic correlations between food prices and the value of the US dollar. We also estimate a latent common factor from 27 IMF Commodity Prices data, which shows a mirror image of the dollar exchange rate.

Isolating the effect of the common factor from individual food prices, we found that remaining idiosyncratic prices exhibit very low volatilities and significantly less persistent dynamics. That is, high fluctuations of prices recently observed in commodity markets might be largely explained by financial market turmoil triggered by the bankruptcy of Lehman Brothers on September 15, 2008. Policy makers, therefore, should put more attention to changes in financial market conditions on top of factors that affect economic fundamentals.

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Table 1. Data Description

Category	ID	Commodity
Cereal	1	Barley, US$ per metric ton
	2	Groundnuts (peanuts), US$ per metric ton
	3	Maize (corn), US$ per metric ton
	4	Rice, US$ per metric ton
	5	Soybean Meal, US$ per metric ton
	6	Soybeans, US$ per metric ton
	7	Wheat, US$ per metric ton
Vegetable Oil	8	Rapeseed Oil, US$ per metric ton
	9	Olive Oil, US$ per metric ton
	10	Palm oil, US$ per metric ton
	11	Soybean Oil, US$ per metric ton
	12	Sunflower Oil, US$ per metric ton
Meat	13	Beef, US cents per pound
	14	Lamb, US cents per pound
	15	Swine (pork), US cents per pound.
	16	Poultry (chicken), US cents per pound
Seafood	17	Fishmeal, US$ per metric ton
	18	Fish (salmon), US$ per kilogram
	19	Shrimp, US$ per kilogram
Other Food	20	Bananas, US$ per metric ton
	21	Oranges, US$ per metric ton
	22	Sugar, Free Market, US cents per pound
	23	Sugar, US Import Price, US cents per pound
Beverage	24	Cocoa Beans, US$ per metric ton
	25	Coffee, Other Mild Arabicas, US cents per pound
	26	Coffee, Robusta, US cents per pound
	27	Tea, US cents per kilogram

Note: We obtain all data from IMF Primary Commodity Prices website. Monthly observations span from January 1980 to August 2013.

Table 2. Univariate Augmented Dickey-Fuller Test Statistics

ID	*ADF*	*p*-value
1	-1.548	0.504
2	-3.009*	0.030
3	-1.949	0.302
4	-1.980	0.286
5	-2.299	0.165
6	-2.083	0.237
7	-1.970	0.286
8	-1.595	0.480
9	-1.676	0.440
10	-2.622	0.082
11	-2.158	0.213
12	-2.283	0.165
13	-0.604	0.868
14	-2.602	0.086
15	-3.601*	0.005
16	-0.316	0.925
17	-1.407	0.585
18	-1.989	0.278
19	-3.196*	0.018
20	-2.216	0.189
21	-1.874	0.334
22	-2.756	0.060
23	-4.072*	0.001
24	-2.218	0.189
25	-2.454	0.116
26	-2.042	0.254
27	-3.555*	0.006

Note: *ADF* and *p*-value denote the augmented Dickey-Fuller *t*-statistics and its associated *p*-value, respectively, when an intercept is present. Asterisk (*) refers the case when the null hypothesis of a unit-root is rejected at the 5% significance level.

Table 3. Cross-Section Dependence

ID	Mean Corr	Group Corr
1	0.188	0.293
2	0.103	0.090
3	0.196	0.360
4	0.071	0.062
5	0.206	0.374
6	0.248	0.418
7	0.172	0.273
8	0.143	0.251
9	0.096	0.063
10	0.179	0.331
11	0.239	0.374
12	0.159	0.291
13	0.076	0.029
14	0.092	-0.007
15	0.051	0.069
16	0.048	-0.017
17	0.094	0.047
18	0.102	0.118
19	0.040	0.022
20	0.078	0.037
21	0.052	0.034
22	0.115	0.140
23	0.110	0.134
24	0.105	0.117
25	0.114	0.295
26	0.116	0.316
27	0.085	0.047
CD = 32.710, p-value = 0.000		

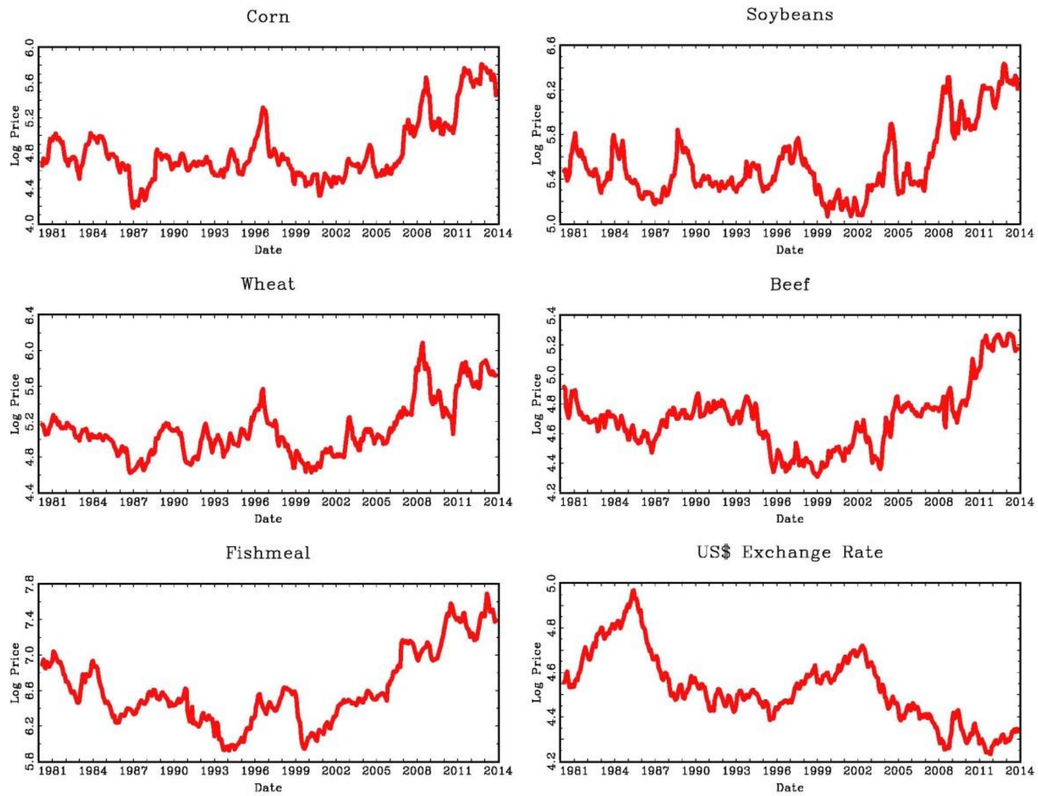
Note: Mean Corr denotes the average correlation of the *ADF* regression residuals. Group Corr is the average correlation of the residuals among the same group commodities.

Table 4. Panic Test Results

	Idiosyncratic Component	
ID	*ADF*	*p*-value
1	-1.894	0.053
2	-2.695*	0.006
3	-3.009*	0.003
4	-1.665	0.090
5	-3.118*	0.002
6	-2.299*	0.019
7	-3.811*	0.000
8	-2.498*	0.011
9	-1.854	0.058
10	-3.814*	0.000
11	-1.690	0.086
12	-3.308*	0.001
13	-1.023	0.278
14	-1.983*	0.042
15	-3.229*	0.002
16	-3.990*	0.000
17	-1.214	0.205
18	-1.315	0.173
19	-2.101*	0.031
20	-4.518*	0.000
21	-3.079*	0.002
22	-2.159*	0.027
23	-2.933*	0.003
24	-1.232	0.197
25	-2.472*	0.012
26	-1.368	0.165
27	-4.182*	0.000
Panel Test Stat = 19.725, *p*-value = 0.000		
Common Factor *ADF* = -2.176, *p*-value = 0.205		

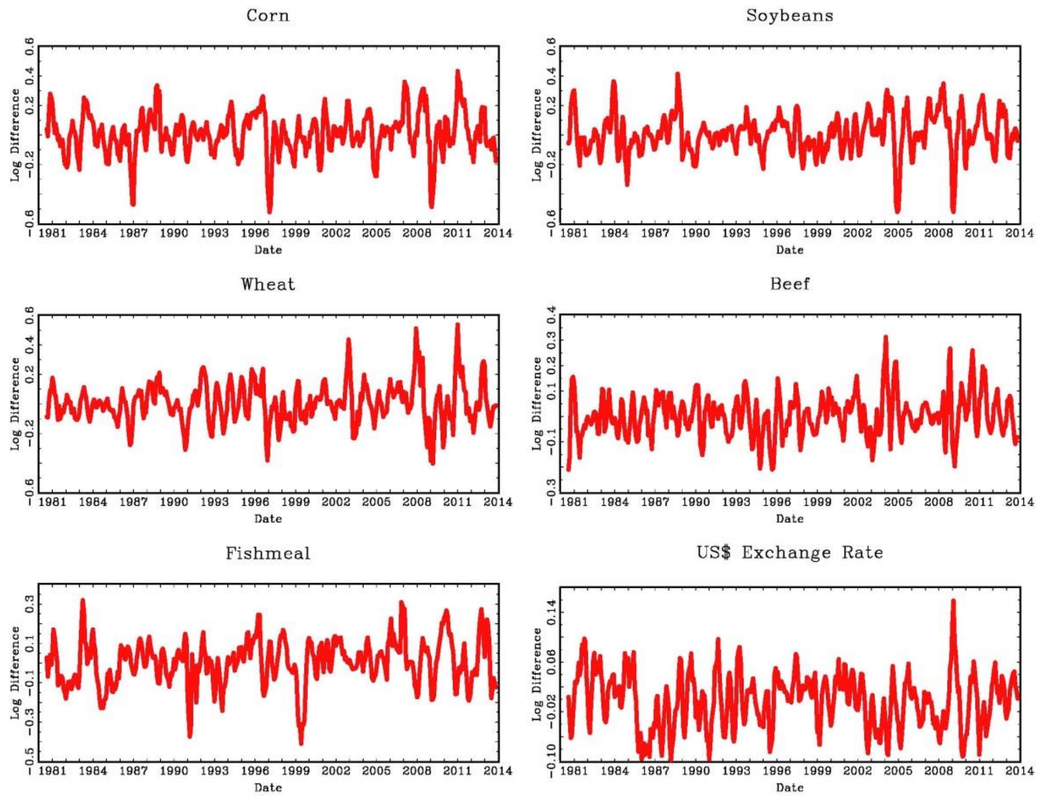
Note: *ADF* for the idiosyncratic component is the ADF *t*-test statistic with no deterministic term for the de-factored commodity prices, while *ADF* for the common factor is the ADF *t*-test statistic with an intercept for the latent common factor estimate. See Bai and Ng (2004) for detailed explanations. Panel test statistic is from the idiosyncratic component *ADF* test statistics.

Figure 1. Food and Beverage Commodity Prices: Level Variables



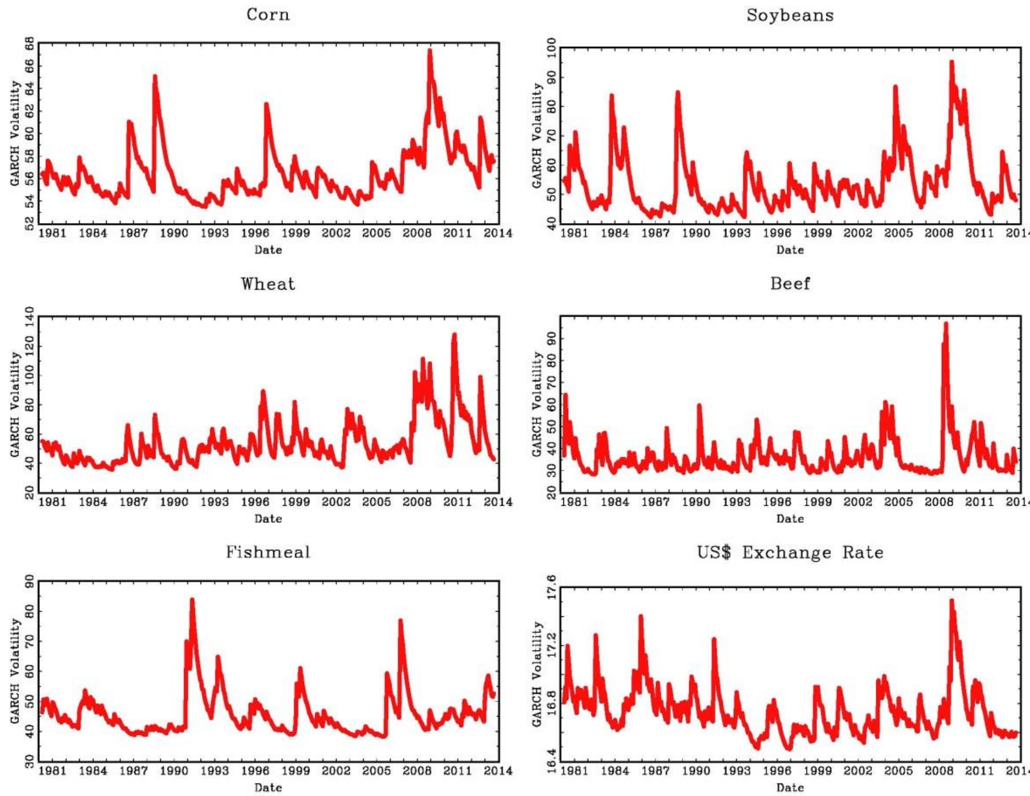
Note: All price data are obtained from IMF website and are log transformed.

Figure 2. Food and Beverage Commodity Prices: Log Differenced Variables



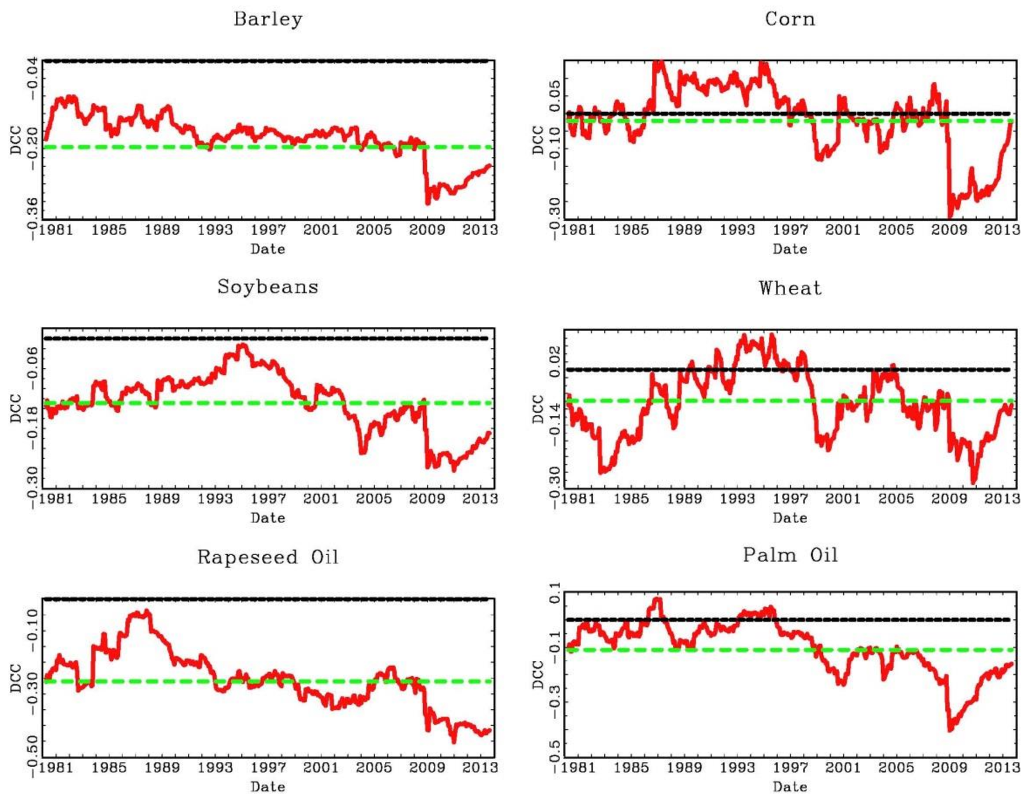
Note: We use log differenced series over one year to get annual growth rates.

Figure 3. GARCH Volatility Estimates



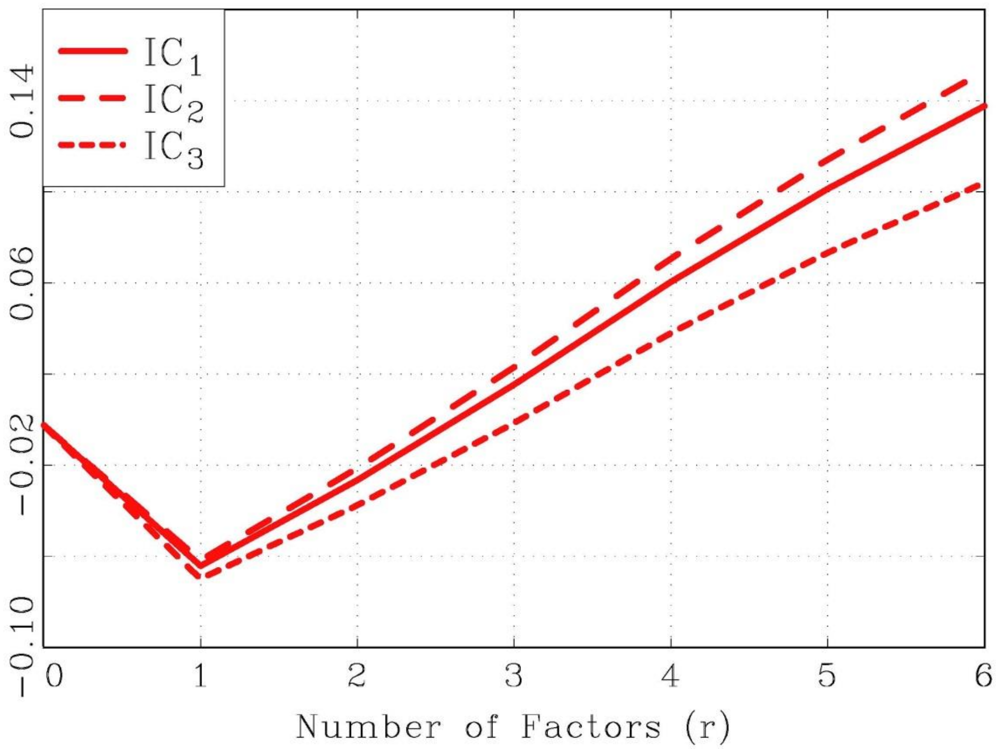
Note: We use the conventional GARCH (1,1) model to obtain GARCH volatility estimates.

Figure 4. Dynamic Conditional Correlation Estimates with the Exchange Rate



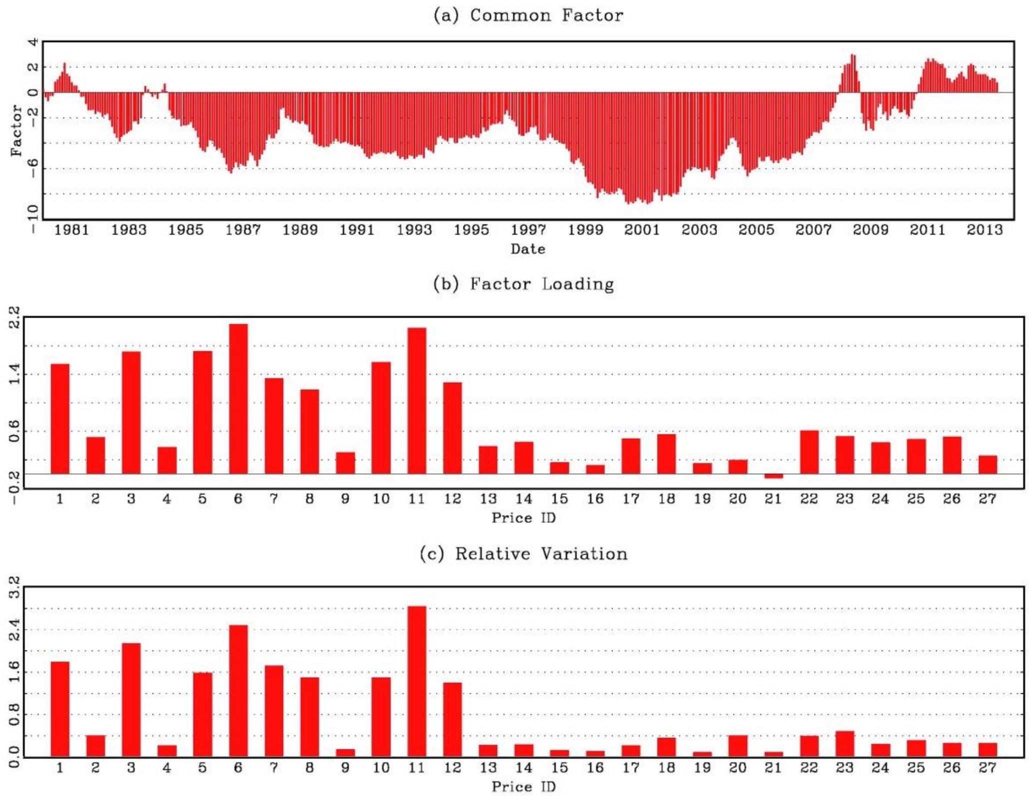
Note: We estimate the dynamic conditional correlation based on the method by Engle (2002). We also report Bollerslev's (1990) constant conditional correlations (dashed lines) and a benchmark zero correlation (dotted lines).

Figure 5. Information Criteria for the Optimal Number of Factors



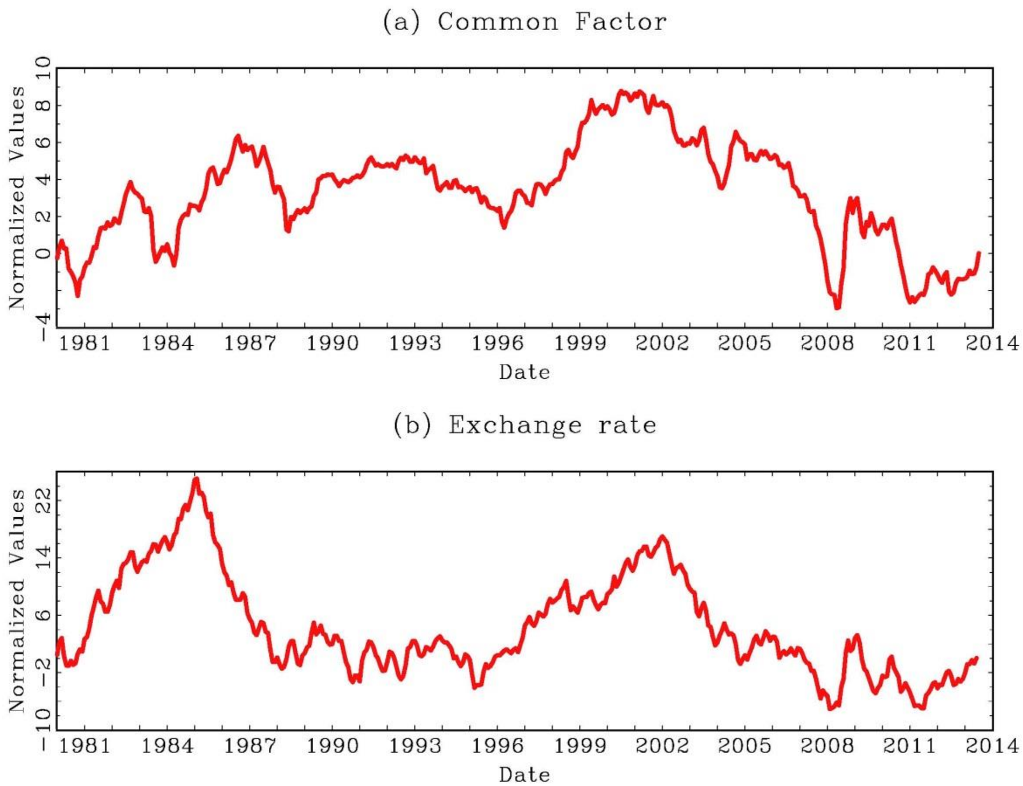
Note: See Bai and Ng (2002) for a detailed explanation on these information criteria.

Figure 6. Latent Factor and Loading Estimates



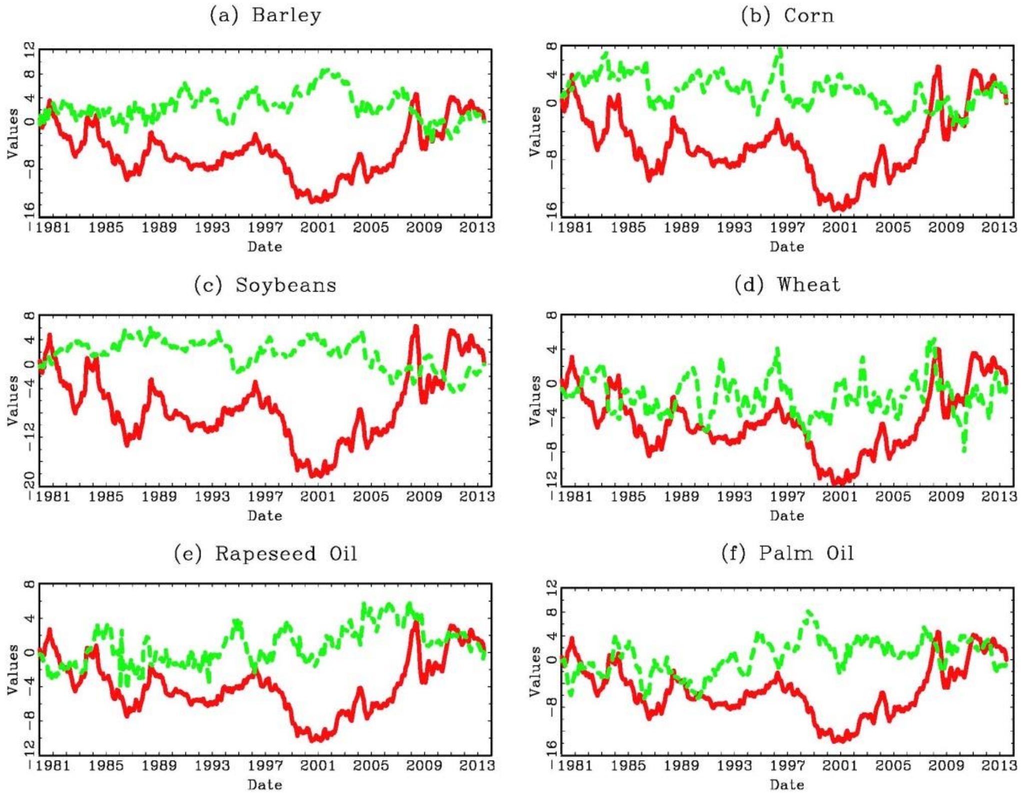
Note: We estimate one latent common factor (f_t) and its associated factor loading (λ_i) for each commodity price series following the PANIC method by Bai and Ng (2004). Relative variations are calculated by taking the standard deviation of $\lambda_i f_t$ relative to the standard variation of the idiosyncratic component ($e_{i,t}$) for each price series.

Figure 7. Common Factor and the Exchange Rate



Note: We report the common factor (f_t) after multiplying by -1. Exchange rate is the price of one US dollar in terms of a basket of major currencies.

Figure 8. Common and Idiosyncratic Components of Food Prices



Note: We report the common component adjusted with the factor loading estimate ($\lambda_i f_t$) and the idiosyncratic component ($e_{i,t}$) for each price series. Dashed lines are estimated idiosyncratic components.