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Abstract

This paper investigates the relationship between global liquidity and commodity and food prices applying a global cointegrated vector-autoregressive model. We use different measures of global liquidity and various indices of commodity and food prices for the period 1980-2011. Our results support the hypothesis that there is a positive long-run relation between global liquidity and the development of food and commodity prices, and that food and commodity prices adjust significantly to this cointegrating relation. Global liquidity, in contrast, does not adjust, it drives the relationship.

JEL-Classification: E52, E58, C32

Keywords: Commodity prices, food prices, global liquidity, cointegration, CVAR analysis,

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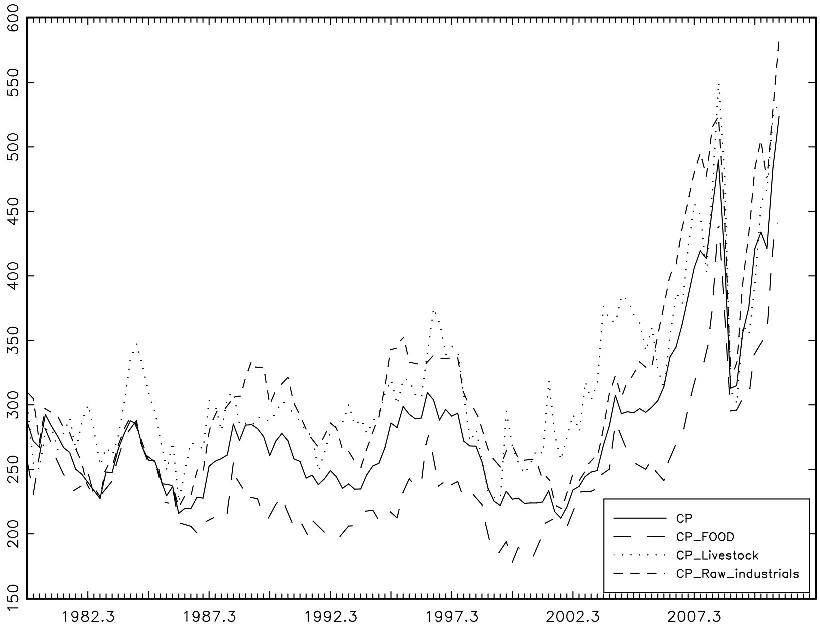
1. Introduction

While prices for most commodities and foodstuff hovered at the same level between 1980 and 2000, they increased dramatically since the early 2000s (Figure 1). Prices peaked in 2008, plummeted during the global financial crisis and started a strong rebound at the beginning of 2009. There have been two major lines of explanation for these developments in food and commodity markets. The first one centres on demand and supply factors. According to Trostle (2008), Krugman (2008), Hamilton (2009), Kilian (2009) and others, the rapid growth of emerging market economies, not least China, has increased world demand for all kinds of food and commodities and led to rapid price increases before the summer of 2008.[5] Prices plunged when demand contracted with the outbreak of the global financial crisis. A second line of explanation argues that these price developments in food and commodity markets have been due to a "financialisation of commodities" (Tang and Xiong 2010, UNCTAD 2011), which has led to a large flow of investment into commodity markets, especially into index investments.[6] According to this view, the rising volumes of financial investments in commodity derivatives markets have led to a synchronised boom and bust of seemingly unrelated commodity prices, driving commodity prices "away from levels justified by market fundamentals, with negative effects both on producers and consumers" (UNCTAD, 2011: vii).

[5] Other studies focusing on supply and demand factors include Thomas, Mühleisen and Pant (2010) and Stürmer (2012).

[6] See also Modena (2011) and Adämmer, Bohl and Stephan (2011).

Figure 1: Development of commodity and food prices, 1980-2011



Source: Compiled by authors with data from CRB and Thomson Reuters.
Note: *CP:* CRB Spot Index (broad index comprising metals, textiles and fibres, fats and oils); *CP_FOOD*: CRB Foodstuffs (hogs, steers, lard, butter, soybean oil, cocoa, corn, Kansas City wheat, Minneapolis wheat, sugar); *CP_Livestock:* CRB Livestock and Products (hides, hogs, lard, steers, tallow); *CP_Raw_Industrials:* CRB Raw Industrials (hides, tallow, copper scrap, lead scrap, steel scrap, zinc, tin, burlap, cotton, print cloth, wool tops, rosin, rubber).

Commodity and food price inflation and volatility has become a major concern for central banks in developing and advanced countries alike. Much of the analysis has focused on the question how monetary policy should respond to such price shocks. For instance, the IMF's World Economic Outlook from September 2011 dedicated a chapter to commodity price swings and monetary policy, finding that commodity prices tend to have stronger and longer-lasting effects on inflation in economies with high food shares in the consumption basket and in economies with less firmly anchored inflation expectations (IMF, 2011).

Instead of investigating monetary authorities' policy responses to commodity and food price shocks, this paper seeks to analyse the effects that monetary policy itself has on commodity and food price movements. In particular, we seek to understand the effects of "global liquidity" – the liquidity created by the world's major central banks – on food and commodity prices. As pointed out in a recent report by the Committee on the Global Financial System (2011: 1), "[g]lobal liquidity has become a key focus of international policy debates", and a potential source of instability. The extremely expansionary monetary policies pursued by the world's major central banks in response to the global financial crisis and the ensuing recession in advanced countries have led to a surge of global liquidity. In this paper we investigate whether such policies, which are certainly warranted from a short-term policy perspective to stabilise financial markets and stimulate output, create unintended negative side effects in terms of long-term inflationary pressures in food and commodity prices.

We use different measures of global liquidity and various indices of commodity and food prices for the period 1980-2011 to investigate the interactions between global liquidity and commodity and food prices on a global level. For our analysis we use a global cointegrated vector-autoregressive (CVAR) model. Our results support the hypothesis that there is a positive long-run relation between global liquidity and the development of food and commodity prices, and that food and commodity prices adjust significantly to this cointegrating relation. Global liquidity, in contrast, does not adjust, it drives the relationship. Our findings highlight the dilemma between short- and long-term policy effects that arises when the central banks of virtually all major economies engage in expansionary monetary policies at the same time, causing a large global liquidity shock that feeds into commodity and food price inflation.

The remainder of this paper is structured as follows. The next section provides a brief overview of previous studies concerned with the link between monetary policy and asset price inflation. Section 3 outlines our empirical analysis, including descriptions of our econometric framework, the construction of the global liquidity and output measures, identification of the long-run structure and hypothesis testing. Section 4 concludes.

2. Literature review

Before turning to the empirical analysis, we briefly review previous theoretical and empirical contributions regarding the linkages between money growth (and thus, liquidity) and asset prices. Research by Fisher (1932), Kindleberger (1978), Borio and Lowe (2002), Congdon (2006), Gerdesmeier, Reimers and Roffia (2009) and others suggests that, historically, boom and bust cycles in asset markets have been closely associated with large movements in money and credit aggregates.[7] The different measures of "excessive credit creation" (i.e., the deviation of the global credit-to-GDP ratio from its trend level and global credit growth)

[7] For an overview of the links between asset price bubbles and monetary policy see ECB (2005; 2010).

used in these studies appear to be good indicators of the build-up of financial imbalances and asset price busts.

Congdon (2006), for instance, investigates the relationship between money supply (specified as broad money) and asset price booms. He analyses the portfolio management of financial institutions like pension funds, finding evidence in favour of a long-run stability of the money/asset ratio (percentage of money in their portfolios) and argues – similar to Meltzer (1995) – that increases in the money supply lead to "too much money chasing too few assets" (Congdon, 2007), suggesting that asset prices rise in order to restore the money/asset ratio.

Several studies investigating the impact of monetary policy and liquidity on asset prices find a special role for housing in the monetary transmission process (Giese and Tuxen, 2007; Adalid and Detken, 2007; Cecchetti et al., 2000). From a theoretical point of view, one can argue that it is a characteristic feature of housing markets that the supply of real estate cannot be easily expanded (Belke and Gros, 2007, OECD, 2005; and Shiller, 2005). Therefore, housing markets should exhibit a lower price elasticity of supply than, for instance, stock markets, which means that additional demand (caused by global excess liquidity) will be reflected to a higher degree in house price increases than on stock markets. Similarly, consumer goods are – not least due to low-cost producers from the emerging markets – nowadays supposed to be largely price-elastic on the supply side, so that additional demand has mainly materialised as additional quantity and not in price increases in recent years.

The role of commodity prices in setting monetary policy has been debated among economists (e.g., Frankel, 1986; Angell, 1992; IMF, 2010) over the last three decades.[8] We would like to highlight some important strands of this literature which also play a major role for our investigations. Drawing on Dornbusch's (1976) theory of exchange rate overshooting, Frankel (1986) pointed to the overshooting in commodity prices. Commodities are exchanged on fast-moving auction markets and, accordingly, are able to respond instantaneously to any pressure impacting on these markets. Following a change in

[8] Recently, the IMF's Global Financial Stability Report from April 2010 investigated the effects of global liquidity expansion, finding strong links between global liquidity expansion and asset prices, such as equity returns, in "liquidity-receiving" economies, as well as official reserve accumulation and portfolio inflows (IMF, 2010).

monetary policy, their price reacts more than proportionately, i.e., they overshoot their new long-run equilibrium, because the prices of other goods are sticky.[9]

Hence, there is some doubt that commodity prices can be used effectively in formulating monetary policy because they tend to be subject to large and market-specific shocks which may not have macroeconomic implications (Marquis and Cunningham, 1990; Cody and Mills, 1991). More importantly in our context and according to a more monetarist view, other researchers argue that commodity price movements are at least to some extent the result of monetary factors and, hence, the causality should run from monetary variables to commodity prices (Bessler, 1984; Pindyck and Rotemberg, 1990; and Hua, 1998).

We believe that this controversy can only be settled as a matter of empirical testing. To do so, we build on Belke, Bordon and Hendricks (2010), who apply a CVAR framework to examine the interactions between money, consumer prices and commodity prices at the global level for major OECD countries for the period 1970-2008. Belke et al. establish long- and short-run relationships among these variables with the process being mainly driven by global liquidity. Moreover, they find that different price elasticities in commodity and consumer goods markets can explain overshooting of commodity over consumer prices. We consolidate and develop their econometric approach by broadening the information set as well as expanding the period under investigation to incorporate the dynamics during the global financial crisis until the current edge, as will be outlined in the next section. The focus of our analysis is on establishing equilibrium relations between global money aggregates and commodity and food prices. Investigating the adjustment behaviour to the long-run relation we seek to examine driving factors of the equilibrium relationships.

In our analysis, we take a global perspective on monetary liquidity. The concept of "global liquidity" has attracted considerable attention in recent years. One of the first studies is Baks and Kramer (1999), who apply different indices of liquidity in seven industrial countries to investigate the direction of the relationship between liquidity and asset returns more deeply. They find evidence in favour of important common components in G7 money growth. Moreover, their results indicate that an increase in G7 money growth is consistent with higher G7 real stock returns. Rüffer and Stracca (2006) estimate that for the G7

[9] Other studies checking for the potential theoretical and empirical importance of monetary conditions for the relationship between commodity prices and consumer goods prices are, for instance, Surrey (1989), Boughton and Branson (1990) and Fuhrer and Moore (1992).

countries around 50% of the variance of a narrow monetary aggregate can be traced back to one common global factor. For instance, the Bank of Japan's extremely expansionary monetary policy stance over the last years is a prominent example of such a global factor. It has been characterised by a significant accumulation of foreign reserves and by extremely low interest rates. By means of carry trades, financial investors took out loans in Japan which they invested in currencies with higher interest rates which in turn should have had an impact on the development of monetary aggregates beyond Japan (Belke and Gros, 2010).

An additional argument in favour of focusing on global instead of national liquidity is that national monetary aggregates have become more difficult to interpret due to the huge increase of international capital flows (Papademos, 2007). Sousa and Zaghini (2006) argue that global aggregates are likely to internalise cross-country movements in monetary aggregates that may make the link between money and inflation and output more difficult to disentangle in the single country case. Giese and Tuxen (2007) further argue that in today's linked financial markets shifts in the money supply in one country may be absorbed by demand elsewhere, but simultaneous shifts in major economies may have significant effects on worldwide goods price inflation. Not only with respect to global liquidity but also as far as the global inflation performance is concerned, available evidence becomes increasingly stronger that the global instead of the national perspective is more important when monetary transmission mechanisms have to be identified and interpreted. For instance, Ciccarelli und Mojon (2005) apply a factor analysis to macroeconomic data of 22 OECD countries and establish that 70% of the variance of the inflation rates of these countries can be traced back to a common factor. The same authors find some empirical evidence in favour of a robust error-correction mechanism, meaning that deviations of national inflation from global inflation are corrected over time. They finally conclude that inflation is to a large degree a global phenomenon. Borio and Filardo (2007) corroborate these results and find that the importance of global factors has increased significantly in recent years. They hence argue that the traditional way of modelling inflation is too country-centred, and that a global approach is more adequate.

3. Empirical analysis

3.1 Data description and aggregation of the international liquidity and output measures

As pointed out by Carney (2011: 2), global liquidity is "an amorphous concept", which "has no agreed definition and, as a consequence, there has been no coherent policy approach to tame its more violent tendencies." In order to make our analysis not dependent on one single way of measuring global liquidity, we apply different indicators. For our baseline analysis we use the aggregate of nominal money for major economies (details of calculation are provided below). We also use two alternative measures for global liquidity: US M0 (seasonally adjusted) plus total foreign exchange reserves excluding gold (Chinn, 2011; Matsumoto, 2011); and total foreign exchange reserves excluding gold. The results we obtain with these different global liquidity measures yield similar results for the long-run relationship between global liquidity and commodity and food prices. For the sake of convenience, in the following we present only the empirical models attained with the international nominal money stock (M_G) variable. The other results are available upon request from the authors.

The selected monetary aggregates to construct the global money measure M_G are M2 for the U.S. and Japan, M3 for the Euro Area, and mostly M3 or M4 for the other countries.[10] We use quarterly data ranging from the first quarter of 1980 to the first quarter of 2011. The aggregated data for the international liquidity and output measures contain broad money aggregates for the United States, the Euro Area, the United Kingdom, Japan, Canada, Australia, New Zealand, Denmark, Norway, Sweden, Switzerland and the BRIC countries. The BRIC data enter the global time series for Brazil in 1990, for Russia in 1994, for India in 2007 and for China in 1999 in the first quarter respectively. The country set under consideration represents approximately 80% of world GDP in 2011 and presumably a considerably larger share of the global financial markets.[11]

Further variables included in the empirical analysis are nominal GDP (Y_G) and the consumer price index CPI (CPI_G) on a global level. We also include the nominal effective exchange rate of the US dollar (USD_EER) to account for dollar valuation effects, and data on exports

[10] The data are taken from the IMF' IFS, the BIS, Thomson Financial Datastream and the EABCN database and are seasonally adjusted.
[11] According to our calculations based on IMF data for the GDP aggregate considering the BRIC countries. Approximately 65% of the world GDP in 2011 is covered when BRIC data are not included.

of emerging and developing economies (EX_EC)[12] to the rest of the world as proxies for demand driven impulses for the development of commodity and food prices, which are represented by indices of the Commodity Research Bureau (CRB). Our main results are reported for the CRB's food prices index (CRB Foodstuffs, CP_FOOD) and the price development of a broad commodity prices index (CRB Spot Index, CP).[13] An advantage of using indices of commodity groups rather than individual commodity prices is that idiosyncratic factors impacting on individual commodity markets should have far less influence at the level of a multi-commodity, broadly-based index.

When aggregating the country-specific time series we follow the approach suggested by Beyer et al. (2000) and applied by Giese and Tuxen (2007) in the same context. The international measures are generated by transforming the country specific series considering market rates as well as PPP rates, with 2005 as the base year. The aggregation of the global money and output series reflects the weight of the respective economy calculated as a proportion of the summation of the GDPs. Forming the international aggregates as weighted sums avoids the under-representation of countries with narrower definitions of their monetary aggregates and vice versa.

Figure 2 illustrates the development of the global money (M_G) and the global output (Y_G) measure under consideration. The inspection of the time series reveals that global liquidity and global output has strongly grown, not only in the last 8 quarters, when the co-movement with increasing commodity and food prices was a main argument in the sense that excess liquidity has been considered as an important factor for the explanation of this development. In the following econometric analysis we will examine this co-movement of global liquidity and commodity price inflation more thoroughly. We consistently report the results for the conversion of the national series by market rates. Applying the global measures derived from the national series converted by PPP rates yields comparable empirical results.[14]

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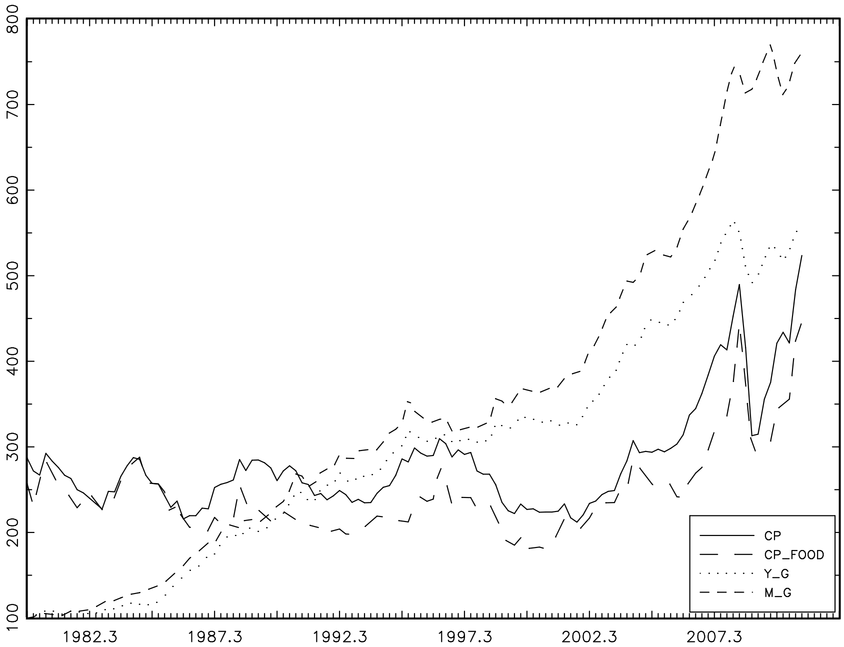
[12] We follow the IMF's country classification, according to which Emerging and Developing Economies comprises countries of Central and Eastern Europe, the Commonwealth of Independent States, Developing Asia, ASEAN-5, Latin America and the Caribbean, Middle East and North Africa, and Sub-Saharan Africa. The BRICS economies are hence included in this classification. For details and a list of countries see: http://www.imf.org/external/pubs/ft/weo/2011/02/weodata/groups.htm.

[13] For composition of these indices see the Figure 1.

[14] For the following description of the empirical analysis we present the results for the food and commodity prices spot indices. As for the international money measure we report the results for the nominal broad money

Figure 2: Global money, global output and commodity prices, 1980-2011



Source: Compiled by authors with data from CRB, Thomson Reuters/Jeffries, IMF, EABCN.

3.2 Econometric framework and time series properties of the data

The econometric framework we apply is a cointegrated vector-autoregressive (CVAR) model. A pertinent problem in time-series econometrics is that of non-stationarity adversely affecting inference. The most common solution to this issue is differencing the data until it becomes stationary but at the same time this implies losing information on the levels of the data generating process. The CVAR framework allows avoiding the loss of information by modelling non-stationary data through linear combinations of the levels of the variables under consideration. Thus the dynamic system of time-series variables of the CVAR approach enables us to model short and long-run dependencies. The basic representation is a p-dimensional vector autoregressive model with Gaussian errors ($\epsilon_{it} \sim iidN(0, \Omega)$):

$$X_t = A_1 X_{t-1} + \cdots + A_k X_{t-k} + \Phi D_t + \epsilon_t, \quad t = 1, \dots, T \qquad (1),$$

where X_t are the variables of interest and D_t is a vector of deterministic components, containing the constant of the model and dummy variables. Reformulating the model in an

aggregate. Using alternative measures, e.g. global aggregates including data on BRIC countries as far as their availability is given, yields comparable outcomes. The results are available from the authors upon request. We do not regard employing proxy measures in a variable system as checking for robustness, but are aware that an altered information set could as well represent a different theory underlying the data.

error correction form allows distinguishing between stationarity that is created by linear combinations of the variables and stationarity created by first differencing:

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k-1} + \Phi D_t + \epsilon_t\,, \quad t = 1, \dots, T, \qquad (2).$$

The error correction model (ECM) representation of the VAR model provides a favourable transformation. Combining levels and differences, the multicollinearity often present in macroeconomic data is reduced. In addition the ECM form of the model gives an intuitive explanation of the data, categorising the effects in long- (Π) and short- (Γ) run information. The logical inconsistency with $X_t \sim I(1)$ is resolved by transforming the multivariate model and reducing the rank of Π to $r < p$, with p being the number of variables. The reduced rank matrix can be factorised into two $p \times r$ matrices α and β ($\Pi = \alpha\beta'$). The factorisation provides r stationary linear combinations of the variables (cointegrating vectors) and $p - r$ common stochastic trends that are creating the nonstationary property in the data system. Formulating the cointegration hypothesis as a reduced rank condition on the matrix $\Pi = \alpha\beta'$ implies that the processes ΔX_t and $\beta' X_t$ are stationary, while the levels of the variables X_t are nonstationary. Therefore the ECM model allows for the variables contained in X_t to be integrated of order 1 (I(1)). For the discussion of the deterministic components in D_t and their specification the error correction model (2) can be rewritten in a more concentrated notation as

$$Z_{0t} = \alpha\beta' Z_{1t} + \Psi Z_{2t} + \epsilon_t\,, \quad \epsilon_{it} \sim iidN(0, \Omega) \qquad (3a)$$

with

$$Z_{0t} = X_t, \qquad Z_{1t} = \begin{pmatrix} X_{t-1} \\ D_{t-1}^R \end{pmatrix}, \quad Z_{2t} = \begin{pmatrix} \{\Delta X_{t-i}\}_{i=1}^{k-1} \\ D_t^U \end{pmatrix} \qquad (3b)$$

where $\Psi = [\Gamma_1, \Gamma_2, \dots, \Gamma_{k-1}, \Phi]$ and D_{t-1}^R is a d_R-dimensional vector of variables restricted to the cointegrated space and D_t^U are d_U unrestricted deterministic terms. The dimensions of Z_{0t}, Z_{1t} and Z_{2t} are *p, (p+*d_R*)* and *p(k-1)+* d_U respectively.

To ascertain the unit root properties of the individual time series we apply Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) test statistics (Table 1). The formal testing supports the application of the cointegration framework since the time series under consideration are integrated of order one.

Table 1: Unit root testing

	M_G	Y_G	CPI_G	CP	CP_FOOD	EX_EC	USD_EER
	Levels						
ADF	2.720	1.219	-4.36	0.378	0.017	1.442	-1.819
PP	2.533	1.199	-3.276	0.231	-0.245	1.159	-1.523
	First-Differences						
ADF	-5.964***	-8.249***	-5.791***	-5.836***	-6.027**	-7.506***	-4.871***
PP	-8.087***	-9.630***	-8.380***	-7.935***	-9.981***	-7.777***	-7.523***

Note: Asterisks refer to level of significance: *10%, **5%, ***1%. Testing specifications: Augmented Dickey-Fuller: 2 lags, constant term; Phillips-Perron: 2 lags (Newey-West), constant term, *tau*-statistic reported.

3.3 Lag length selection and diagnostic testing on the unrestricted VAR model

The asymptotic results depend on the adequate specification and the appropriateness of the choice of the cointegrating rank specification of the underlying model. Specifying the lag length of the VAR has strong implications for the subsequent modelling choices. Choosing too few lags could lead to systematic variation in the residuals whereas choosing too many lags comes with the penalty of fewer degrees of freedom (as adding another lag, adds $p \times p$ parameters). With regards to the formal testing based on the maximum of the likelihood function, the choice of a lag length of two is supported for our data by the "Schwarz" and "Hannan-Quinn" information criteria and lag reduction tests (see Tables 2 and 3).

Table 2: Lag length selection and residual analysis (x'_{t_1})

(Effective Sample: 1981:02 to 2011:01)

k	*Schwarz Criterion*	*Hannan-Quinn Criterion*	*LM Test LM(1)*	*LM Test LM(k)*
4	-40.391	-41.977	0.000	0.069
3	-41.086	-42.328	0.001	0.311
2	-41.849	-42.746	0.015	0.834
1	-42.142	-42.693	0.000	0.000

Table 3: Lag length selection and residual analysis ($x_{t_2}^{'}$)
(Effective Sample: 1981:02 to 2011:01)

K	*Schwarz Criterion*	*Hannan-Quinn Criterion*	*LM Test LM(1)*	*LM Test LM(k)*
4	-46.089	-48.406	0.001	0.513
3	-46.758	-48.579	0.033	0.101
2	-47.848	-49.172	0.771	0.573
1	-48.384	-49.212	0.000	0.000

Estimation of the VAR model is based on the assumption that the residuals display Gaussian properties. Extraordinarily large shocks corresponding to economic reforms or intervention and by those possibly marking structural breakpoints in the data series tend to cause a violation of the normality assumption. The deviation from the normality assumption leads to distorted statistical inferences. Hence, it is important to identify the dates of such shocks and to correct them with intervention dummies (Juselius, 2006). We correct for innovational outliers indicated by large residuals due to shocks to the innovation term that are diffused in the autoregressive structure of the data-generating process, which in terms of distorting inference on the cointegration rank are less problematic than additive outliers (Nielsen, 2004). Incorporating dummy variables ($D0804p, D0901p$) we account for unconventional large scale expansionary monetary policy implemented during the peak of the financial crisis in 2008 by most of the major central banks, not least the Federal Reserve.

The empirical analysis is presented for an information set focusing on the effects of global macro-aggregates on food prices ($x_{t_1}^{'} = [M\_G, CPI\_G, CP\_FOOD, Y\_G, USD\_EER]_t$) and variables ($x_{t_2}^{'} = [M\_G,\ CPI\_G, CP, Y\_G, USD\_EER, EX\_EC]_t$) that draw on the impact of global liquidity on a broader level of commodity prices and provides insights for the role of emerging economies' exports and their impact on the price variables.

Tables 4 and 5 report the univariate and multivariate residual analysis of the unrestricted VAR(2). The null hypothesis of normality for the multivariate model is rejected due to empirical evidence of deviations from normality in skewness and/or kurtosis for the commodity and food prices as well as the consumer price data. Although the commodity and food price time series display high fluctuations especially for the last periods of the data sample, there is hardly any evidence of second order ARCH effects according to the univariate statistics. We do not consider even moderate ARCH-effects as highly problematic since Rahbek et al. (2002) show that the cointegration rank testing is still robust in this case.

Our formal misspecification tests indicate rejection of multivariate normality and ARCH effects due to the above mentioned features of the commodity and food price series. Overall the VAR(2) model seems to provide a reasonable description of the information contained in the data. The following estimation of our global CVAR is based on modelling the data process with two lags and the specified deterministic terms for outlier correction and linear trends in the variables that are restricted to the cointegrating relations as well.

Table 4: Residual analysis and diagnostic testing on the unrestricted VAR(2) model (x'_{t_1})

Multivariate tests				
Residual autocorrelation				
LM(1)	χ^2 (25) = 45.009 [0.008]			
LM(2)	χ^2 (25) = 20.528 [0.719]			
Test for Normality	χ^2 (10) = 29.707 [0.001]			
Test for ARCH				
LM(1)	χ^2 (225) = 198.666 [0.896]			
LM(2)	χ^2 (450) = 430.953 [0.733]			
Univariate tests				
	ARCH(2)	Normality	Skewness	Kurtosis
ΔM_G	3.380 [0.185]	2.196 [0.334]	0.230	3.348
ΔCPI_G	1.252 [0.535]	4.672 [0.097]	0.321	3.726
ΔCP_FOOD	1.154 [0.562]	3.759 [0.153]	0.285	3.597
ΔY_G	1.112 [0.573]	0.244 [0.885]	-0.009	2.637
ΔUSD_EER	3.001 [0.223]	0.152 [0.927]	0.014	2.674

Note: *p*-values in brackets.

Table 5: Residual analysis and diagnostic testing on the unrestricted VAR(2) model (x'_{t_2})

Multivariate tests

Residual autocorrelation	
LM(1)	χ^2 (36) = 29.939 [0.751]
LM(2)	χ^2 (36) = 34.740 [0.528]
Test for Normality	χ^2 (10) = 35.545 [0.000]
Test for ARCH	
LM(1)	χ^2 (441) = 558.082 [0.000]
LM(2)	χ^2 (882) = 1036.966 [0.000]

Univariate tests

	ARCH(2)	Normality	Skewness	Kurtosis
ΔM_G	2.765 [0.251]	1.516 [0.469]	0.172	3.250
ΔCPI_G	3.277 [0.194]	4.912 [0.086]	0.304	3.764
ΔCP	1.575 [0.455]	0.620 [0.733]	-0.013	2.538
ΔY_G	1.003 [0.606]	0.323 [0.851]	-0.082	2.681
ΔUSD_EER	2.486 [0.289]	0.198 [0.906]	0.058	2.694
ΔEX_EC	4.169 [0.124]	2.014 [0.365]	-0.299	3.118

Note: *p*-values in brackets.

3.4 Estimation and rank determination of the global CVAR

The complex determination of the cointegration rank of the Π–matrix, i.e. the cointegration space of the model, is subject to empirical evidence from various pre-testing indicators. The principal formal testing procedure is the Johansen LR trace test (Johansen, 1988, 1991, 1994) with the results being presented in Tables 6 and 7. The trace test statistic rejects the hypotheses of $p - r = 4$ ($p - r = 5$ respectively) common stochastic trends and $r = 1$ cointegrating relations but fails to reject the hypotheses of $p - r = 3$ ($p - r = 4$ respectively) common trends and $r = 2$ cointegrating relations on a 1% significance level. As there are cases for hypotheses that are close to the unit circle, the size of the test and the power of the alternative can be of almost the same magnitude.

Table 6: Trace test statistics for determination of the cointegration rank for the unrestricted VAR(2) model (x'_{t_1})

p - r	*r*	*Eigenvalue*	*Trace*	*95% Critical Value*	*P-Value*
5	0	0.370	132.963	88.554	0.000
4	1	0.289	76.122	63.659	0.003
3	2	0.136	34.099	42.770	0.288
2	3	0.101	16.142	25.731	0.489
1	4	0.025	3.058	12.448	0.859

Table 7: Trace test statistics for determination of the cointegration rank for the unrestricted VAR(2) model (x'_{t_2})

p - r	*r*	*Eigenvalue*	*Trace*	*95% Critical Value*	*P-Value*
6	0	0.438	179.542	117.451	0.000
5	1	0.381	108.601	88.554	0.001
4	2	0.169	49.580	63.659	0.436
3	3	0.140	26.820	42.770	0.694
2	4	0.052	8.220	25.731	0.972

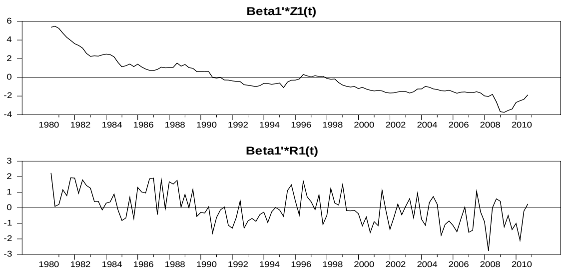
Hence Juselius (2006) suggests using additional information, e.g. recursive graphs of the trace statistic and *t*-values of the adjustment coefficients in the respective potential cointegrating relation in order to choose the appropriate rank. Figures 3 and 4 visualise the overall stationary time path of the candidate cointegrating relations and underpin the formal testing results in favour of a cointegration rank of 2.

The recursive estimation results for constancy of the beta-coefficients, the log-likelihood and the simulation of the trace statistics suggest for both of the information sets the choice of a rank of two. The inspection of the system's eigenvalues and roots of the companion matrix supports this specification.[15]

[15] The graphs and numerical results for the formal testing procedures are available from the authors upon request.

Figure 3: Plots of the two cointegrating vectors for the information set x'_{t_1}

Vectors $\widehat{\beta_1}X_{t_1}$



Vectors $\widehat{\beta_2}X_{t_1}$

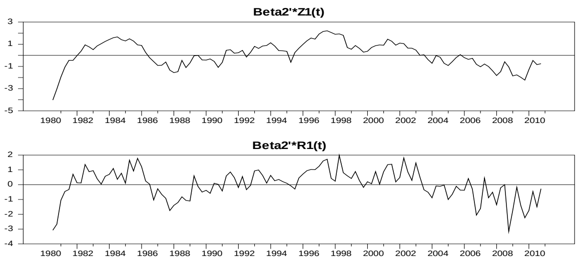
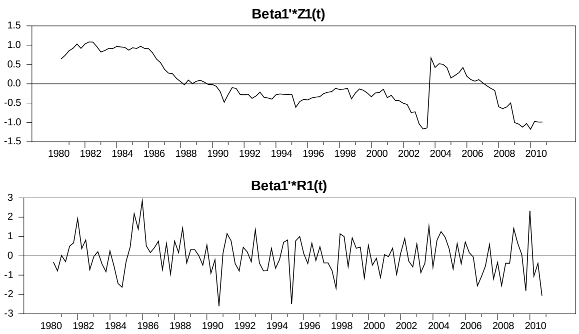
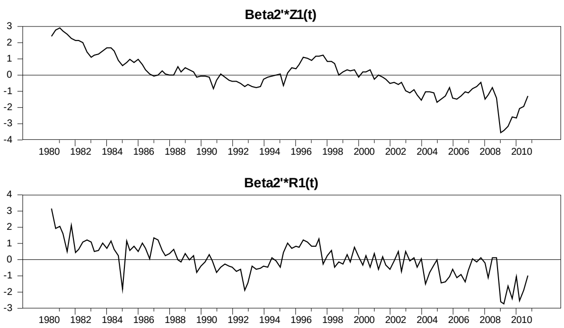


Figure 4: Plots of the two cointegrating vectors for the information set x'_{t2}

Vectors $\widehat{\beta_1}X_{t_2}$



Vectors $\widehat{\beta_2}X_{t_2}$



3.5 Identification of the long-run structure and adjustment to the stationary relations

We approach the identification of the interaction between the global aggregates and the price variables as well as the respective system dynamics by separately accounting for food (x_{t_1}') and commodity prices (x_{t_2}'). The identification of the systems is conducted by imposing restrictions on the long-run and short-run coefficients and thus characterising the equilibrium relations and the underlying error-correcting adjustment behaviour. The information set is defined by the variable vectors

$$x_{t_1}' = [M\_G, CPI\_G, CP\_FOOD, Y\_G, USD\_EER]_t, \quad (4)a$$

$$x_{t_2}' = [M\_G, CPI\_G, CP, Y\_G, USD\_EER, EX\_EC]_t \quad (4)b$$

$$D_t \;= [D0804p, D0901p]_t,\ D_t^U = 1 \text{ and } D_t^R = t \quad (4)c.$$

The system comprising food price data is restricted and specified following the above reasoning and formal testing to a cointegration rank of r=2 and a lag length of k=2 with the imposed restrictions on the long-run structure being not rejected with a p-value of 0.281 ($\chi^2(2) = 2.539$).

$$\hat{\beta}_{1_1}' : (\mathrm{M\_G} - Y\_G) - 0.592\, CPI\_G + 0.311\ \mathrm{CP\_FOOD} \sim \mathrm{I}(0) \quad (5)a$$

$$\hat{\beta}_{2_1}' :\ \mathrm{M\_G} + 0.299\ \mathrm{CP\_FOOD} + 0.811\ \mathrm{Y\_G} + 0.009\ \mathrm{USD\_EER} \sim \mathrm{I}(0) \quad (5)b.$$

The empirically identified long-run structure represented by the cointegrating relations $\hat{\beta}_{1_1}'$ and $\hat{\beta}_{2_1}'$ highlights the significant effect for global money on the development of food prices (cf. the *t*-values in Table 8). Both the stationary spread of global money and global output in the first relationship as well as the global money measure support the long-run influence of (“excess”-) liquidity on food prices. The global consumer price level significantly enters the stationary relationship yet not with the expected sign as for the nominal effective exchange rate of the USD. The expected and observable correlation of the prices of commodities in USD with the price of currencies in USD is not underscored in a steady-state relation with our global data. As the underlying causes of commodity and dollar cycles are not clear cut over a longer period under observation, cyclical trends in commodity prices (and food prices as a sub-aggregate) could have become more attenuated in the present global dataset and the measures we apply.

Table 8: The long-run cointegrating relations (x'_{t_1})

$$\Delta X_t = \alpha\beta' X_{t-1} + \epsilon_t$$

$$\begin{pmatrix} \Delta M\_G_t \\ \Delta CPI\_G_t \\ \Delta CP\_FOOD_t \\ \Delta Y\_G_t \\ \Delta USD\_EER_t \end{pmatrix} = \begin{pmatrix} \underset{(1.149)}{0.174} & \underset{(-1.340)}{-0.235} \\ \underset{(-4.970)}{-0.110} & \underset{(4.505)}{0.115} \\ \underset{(-0.577)}{-0.198} & \underset{(1.198)}{0.476} \\ \underset{(0.834)}{0.118} & \underset{(-0.988)}{-0.162} \\ \underset{(-0.609)}{-0.100} & \underset{(0.544)}{0.103} \end{pmatrix} \begin{pmatrix} 1.000 & \underset{(7.417)}{0.592} & \underset{(-3.976)}{-0.311} & -1.000 & 0.000 & \underset{(-1.048)}{-0.000} \\ 1.000 & 0.000 & \underset{(-4.469)}{-0.299} & \underset{(-24.529)}{-0.811} & \underset{(-0.283)}{-0.009} & 0.000 \end{pmatrix} \begin{pmatrix} M\_G_{t-1} \\ CPI\_G_{t-1} \\ CP\_FOOD_{t-1} \\ Y\_G_{t-1} \\ USD\_EER_{t-1} \\ t \end{pmatrix} + \epsilon_t$$

Combined estimates

	M_G	CPI_G	CP_FOOD	Y_G	USD_EER	*Trend*
ΔM_G	-0.061 [-1.754]	0.103 [1.149]	0.016 [1.725]	0.016 [0.658]	0.002 [1.340]	-0.000 [-1.149]
ΔCPI_G	0.005 [1.050]	-0.065 [-4.970]	-0.000 [-0.216]	0.017 [4.613]	-0.001 [-4.505]	0.000 [4.970]
ΔCP_FOOD	0.278 [3.549]	-0.117 [-0.577]	-0.081 [-3.823]	-0.188 [-3.376]	-0.004 [-1.198]	0.000 [0.577]
ΔY_G	-0.044 [-1.353]	0.070 [0.834]	0.012 [1.340]	0.013 [0.569]	0.001 [0.988]	-0.000 [-0.834]
ΔUSD_EER	0.003 [0.088]	-0.059 [-0.609]	0.000 [0.018]	0.016 [0.611]	-0.001 [-0.544]	0.000 [0.609]

Note: *t*-values in brackets. Test statistic restricted model: CHISQR(2) = 2.539, corresponding p-value: 0.281.

Having identified the long-run stationary relations, the characteristics of the adjustment behaviour and the common driving trends as the cumulated sum of empirical shocks to the respective variable can be examined by imposing restrictions on the adjustment coefficients, e.g. exogeneity relative to the information set under consideration. Restrictions on the α-coefficients have implications for the common stochastic trends of the system as their orthogonal component multiplies the unit root components. The hypothesis that the cumulated residuals from a specific equation indicate a common driving trend can be specified as a zero row in the adjustment coefficients, implying that the respective variable is weakly exogenous for the afore-identified long-run relation. The opposite hypothesis, i.e. that the residuals of a relation have transitory but no permanent effects on the variables of the system, can be tested by specifying a unit vector in alpha. With the above reasoning in line, long-run weak exogeneity is accepted (with the x'_{t_1} information set based on a CHISQR(4) = 4.907 test statistic with a p-value of [0.297]) for the global liquidity aggregate, i.e. it is not error-correcting and a unit root component is “driving” the system. On the other hand food prices adjust to the long-run relation and exhibit error-correcting behaviour (based on CHISQR(4) = 17.637 test statistic with a p-value of [0.001]).

Focusing on the impact of global liquidity on commodity prices on a broader level and aggregated consumer prices, Table 9 shows the results of the long-run relationships ((6)a,b) for the $x_{t2}^{'}$ information set. The identification of the long-run structure imposed by the joint restrictions of the vectors of $\hat{\beta}_{1_2}^{'}$ and $\hat{\beta}_{2_2}^{'}$ is supported on common significance levels (based on a CHISQR(2) = 4.181 test statistic with a p-value of [0.124]). International money is significantly part of an equilibrium relation including commodity and consumer prices. The relevance of fast growing emerging economies and accordingly the demand for commodities is accounted for by including exports of emerging countries which enter the long-run relationship. Excluding consumer prices from the long-run relation but considering the nominal effective exchange rate of the USD does not yield the expected sign (though not significant) for commodity prices to enter the long-run equilibrium. The analysis of the adjustment coefficients emphasises the above results for food prices. The international liquidity measure is again found to be weakly exogenous (as indicated by a CHISQR(4) = 7.155 test statistic and the according p-value of [0.128]). As with food prices, commodity prices are correcting to the equilibrium errors and deviations from the system's long-run stationary path (as testing for a unit vector in the respective orthogonal complement of the adjustment coefficients indicates by a CHISQR(4) = 26.004 test statistic).

$$\hat{\beta}_{1_2}^{'}: \mathrm{M\_G} + 1.460\ \mathrm{CPI\_G} + 0.206\ \mathrm{CP} + 0.198\ Y\_G + 0.002\ \mathrm{t} \sim \mathrm{I}(0) \quad \text{(6)a}$$

$$\hat{\beta}_{2_2}^{'}: \mathrm{M\_G} - 5.324\ \mathrm{CP} - 5.296\ \mathrm{USD\_EER} + 10.855\ EX\_EC - 0.307\ \mathrm{t} \sim \mathrm{I}(0) \quad \text{(6)b.}$$

Table 9: The long-run cointegrating relations (x'_{t_2})

$$\Delta X_t = \alpha\beta' X_{t-1} + \epsilon_t$$

$$\begin{pmatrix} \Delta M\_G_t \\ \Delta CPI\_G_t \\ \Delta CP_t \\ \Delta Y\_G_t \\ \Delta USD\_EER_t \\ \Delta EX\_EC_t \end{pmatrix} = \begin{pmatrix} -0.080 & 0.000 \\ (-1.788) & (0.141) \\ 0.019 & -0.001 \\ (3.163) & (-4.575) \\ 0.353 & -0.002 \\ (5.229) & (-1.488) \\ -0.035 & -0.000 \\ (-0.832) & (-0.016) \\ -0.012 & -0.001 \\ (-0.249) & (-1.152) \\ 0.236 & 0.004 \\ (2.237) & (1.632) \end{pmatrix} \begin{pmatrix} 1.000 & -1.460 & -0.206 & -0.198 & 0.000 & 0.000 & -0.002 \\ & (-6.095) & (-4.615) & (-2.121) & & & (-2.068) \\ 1.000 & 0.000 & 5.324 & 0.000 & 5.296 & -10.855 & 0.307 \\ & & (1.881) & & (1.745) & (-6.184) & (7.284) \end{pmatrix} \begin{pmatrix} M\_G_{t-1} \\ CPI\_G_{t-1} \\ CP_{t-1} \\ Y\_G_{t-1} \\ USD\_EER_{t-1} \\ EX\_EC_{t-1} \\ t \end{pmatrix} + \epsilon_t$$

Combined estimates

	M_G	CPI_G	CP	Y_G	USD_EER	EX_EC	*Trend*
ΔM_G	-0.079 [-1.789]	0.116 [1.788]	0.017 [1.518]	0.016 [1.788]	0.001 [0.141]	-0.002 [-0.141]	0.000 [0.508]
ΔCPI_G	0.019 [3.056]	-0.028 [-3.163]	-0.008 [-4.916]	-0.004 [-3.163]	-0.004 [-4.575]	0.008 [4.575]	-0.000 [-5.042]
ΔCP	0.350 [5.203]	-0.515 [-5.229]	-0.086 [-4.994]	-0.070 [-5.229]	-0.013 [-1.488]	0.027 [1.488]	-0.001 [-2.515]
ΔY_G	-0.035 [-0.834]	0.051 [0.832]	0.007 [0.664]	0.007 [0.832]	-0.000 [-0.016]	0.000 [0.016]	0.000 [0.157]
ΔUSD_EER	-0.013 [-0.278]	0.017 [0.249]	-0.005 [-0.393]	0.002 [0.249]	-0.007 [-1.152]	0.015 [1.152]	-0.000 [-1.052]
ΔEX_EC_t	0.240 [2.282]	-0.344 [-2.237]	-0.026 [-0.966]	-0.047 [-2.237]	0.023 [1.632]	-0.046 [-1.632]	0.001 [1.098]

Note: *t*-values in brackets. Test statistic restricted model: CHISQR(2) = 4.181, corresponding p-value: 0.124.

Summing up, both food and commodity prices show significant adjustment behaviour to the identified long-run structure and corroborate the hypothesis of being driven by global liquidity in the long-term for the examined data. The long-run stochastic path of the system characterised by the adjustment coefficients is influenced by the international money variable, while at the same time it exhibits “no levels feedback”. That is, it is not influenced by the other variables of the system and is weakly exogenous for the long-run structure. Food and commodity prices instead are not found to be weakly exogenous for the long-run parameters and thus receive “long-run feed-back” from the system.

4. Conclusion

In this paper we sought to investigate the relationship between global liquidity and commodity and food prices using a global CVAR model. We use different measures of global liquidity and various indices of commodity and food prices. In order to understand the interactions between monetary aggregates, inflation and commodity prices on a global level, we primarily focus on long-run equilibrium relations and emphasise the role of monetary factors in explaining food and commodity price movements. Our results provide noteworthy insight into the links between monetary policy and commodity and food price inflation and support the hypothesis that there is a positive long-run relation between global liquidity and the development of food and commodity prices. Food and commodity prices significantly adjust to the cointegrating relations and exhibit a long-term co-movement with liquidity on an international level.

Our findings highlight the dilemma that arises when the central banks of virtually all major economies engage in expansionary monetary policies at the same time in order to stabilise their domestic economies and financial sectors, causing a large global liquidity shock that feeds into commodity and food price inflation. While such expansionary monetary policies may be warranted to adequately respond to financial crisis, economic contraction, high unemployment and deflationary tendencies, our analysis suggests that there are pronounced negative side-effects in terms of commodity and food price inflation.

Price increases in foodstuff and commodities can have serious implications for public and private budgets in developing, emerging and advanced countries alike. Whereas exporters of foodstuff and commodities benefit from rising prices, which should also boost the government's revenue position in commodity exporting countries (provided the state's ability tax these export earnings), the effects for households that do not derive their income from commodity producing sectors are negative in both exporting and importing countries. Especially poorer households usually suffer most from food price inflation and rising energy prices (which we did not analyse in this paper), since they tend to spend a larger proportion of their income on these items. Even though we are not forecasting food and commodity price developments in this study, our analysis suggests that further price hikes may be in store given current expansionary monetary conditions.

Over the period that we observed, 1980-2011, food and commodity price inflation were apparently driven by monetary expansion in the world's major economies. Our results can be seen as supporting the view of a "financialisation of commodities", where food and commodity prices are driven to a large extend by flows of portfolio investment seeking return in commodity markets and not merely by demand from the real economy. Policymakers should take into account the negative side-effects of loose monetary policy and consider stricter regulation of food and commodity markets – such as the imposition of tighter limits on speculative positions in food commodities – to prevent a further flow of liquidity into these markets.

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