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Ansgar Belke • Ingo G. Bordon • Torben W. Hendricks

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DIW Berlin
German Institute for Economic Research
Mohrenstr. 58
10117 Berlin
Tel. +49 (30) 897 89-0
Fax +49 (30) 897 89-200
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Global Liquidity and Commodity Prices –
A Cointegrated VAR Approach for OECD Countries

by

Ansgar Belke (University of Duisburg-Essen and DIW Berlin)

Ingo G. Bordon (University of Duisburg-Essen)

Torben W. Hendricks (University of Duisburg-Essen)

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Abstract

This paper examines the interactions between money, consumer prices and commodity prices at a global level from 1970 to 2008. Using aggregated data for major OECD countries and a cointegrating VAR framework, we are able to establish long run and short run relationships among these variables while the process is mainly driven by global liquidity. According to our empirical findings, different price elasticities in commodity and consumer goods markets can explain the recently observed overshooting of commodity over consumer prices. Although the sample period is rather long, recursive tests corroborate that our CVAR fits the data very well.

JEL codes: E31, E52, C32, F42

Keywords: Commodity prices, cointegration, CVAR analysis, global liquidity, inflation, international spillovers

Authors:

Prof. Dr. Ansgar Belke, Chair for Macroeconomics and Director of the Institute of Business and Economics, Department of Economics, University of Duisburg-Essen, Campus Essen, Universitaetsstr. 12, 45117 Essen, Germany, e-mail: ansgar.belke@uni-due.de, phone: +49 201 183 2277, fax: +49 201 183 4181

Ingo G. Bordon, Department of Economics, University of Duisburg-Essen, Campus Essen, Germany

Torben W. Hendricks, Department of Economics, University of Duisburg-Essen, Campus Essen, Germany

1. Introduction

Against the background of steadily increasing global liquidity since the beginning of the century in most industrial countries as well as in numerous emerging market economies with a dollar peg, especially China, broad money growth has been running well ahead of nominal GDP. Surprisingly enough, for long time consumer price inflation has remained widely unaffected by the strong monetary dynamics in many regions in the world. Over the same time horizon, however, many countries have experienced sharp but sequential booms in asset prices, such as commodity, real estate or share prices (Schnabl and Hoffmann, 2007). Between 2001 and 2007, for instance, house prices strongly increased by 40 to 60% in a number of OECD countries, the CRB commodity price index surged by 84% in the same period and stock prices more than doubled in nearly all major markets from 2003 to 2007. Many observers interpret the sequence of increases of asset prices as the result of liquidity spill-overs to certain asset markets (Adalid and Detken, 2007, Greiber and Setzer, 2007).

From a monetary policy perspective, the different price dynamics of assets and goods prices in recent years raises the question as to whether the money-inflation nexus has been changed (thereby calling into question the close long-term relationship between monetary and goods price developments that was observed in the past) or whether effects from previous policy actions are still in the pipeline. To investigate the relative importance of these developments, this study tries to establish an empirical link between money, asset prices and goods prices. For this purpose, we estimate a variety of cointegrated VAR (CVAR) models including a measure of global liquidity, proxied by a broad monetary aggregate in the OECD countries under consideration (United States, Euro area, Japan, United Kingdom, Canada, South Korea, Australia, Switzerland, Sweden, Norway and Denmark) and analyse the impact of global liquidity on commodity and goods price inflation. The basic idea is that different price elasticities of supply lead to differences in the dynamic pattern of price adjustment to a global liquidity shock. While goods prices adjust only very slowly to changing global monetary conditions due to plentiful supply of consumer goods from emerging markets, asset prices such as commodity prices react much faster since the supply of commodities cannot be easily expanded. Thus disequilibria on these markets are generally balanced out by price adjustments.

The main emphasis is on globally aggregated variables, which implies that we do not explicitly deal with spill-overs of global liquidity to national variables. The main motivation for this specific way of proceeding is heavily related to recent research according to which inflation appears to be a global phenomenon. So far, the relationship between money growth, different categories of asset prices and goods prices has been little studied in an international context. Only recently, a number of authors suggested specific interactions of global liquidity with global consumer price and asset price inflation (Baks and Kramer, 1999, Sousa and Zaghini, 2006, and Rüffer and Stracca, 2006). However, so far no study has tried to systematically analyze differences between an asset class such commodities and goods in the dynamic pattern of price adjustment to a global liquidity shock.

The remainder of the paper is organised as follows: in section 2, we convey an impression of the global perspective of the monetary transmission process. In section 3, we develop some simple theoretical considerations to illustrate the potential role of different supply elasticities as potential drivers of commodity- and goods-specific price adjustments to global liquidity shocks. In section 4 we turn to the technical details on our estimation strategy using the CVAR technique on a global scale and reports on our estimation results. Section 5 finishes with some policy conclusions.

2. The global perspective of monetary transmission

Both with respect to global inflation and to global liquidity performance, available evidence becomes stronger that the global instead of the national perspective is more important when the monetary transmission mechanism has to be identified and interpreted. For instance, Ciccarelli and Mojon (2005) find empirical evidence in favour of a robust error-correction mechanism, meaning that deviations of national inflation from global inflation are corrected over time. Similarly, Borio and Filardo (2007) argue that the traditional way of modeling inflation is too country-centred and a global approach is more adequate. Considering the development of global liquidity over time, the question is often raised whether and to what extent global factors are responsible for it. Rüffer and Stracca (2006) investigate this aspect for the G7 countries in the framework of a factor analysis and conclude that around fifty percent of the variance of a narrow monetary aggregate can be traced back to one common global factor. One prominent example of such a global factor is, for instance, the expansionary monetary policy stance of the Bank of Japan (BoJ) during the last years. It has been characterised by a significant accumulation of foreign reserves and by extremely low interest rates - at some time even approaching zero. By means of carry trades, financial investors took up loans in Japan and invested the proceeds in currencies with higher interest rates. Such kind of capital transactions has impacts on the development of monetary aggregates far beyond the special case of Japan and national borders in general (see, e.g., Schnabl and Hoffmann, 2007).

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. Simply accounting for the external sources of money growth and then mechanically correcting for cross-border portfolio flows or M&A activity, on the presumption of their likely less relevant direct effects on consumer prices, is not a sufficient reaction. Instead, these transactions have to be investigated with respect to their information content and potential wealth effects on residents' income and on asset prices which might backfire to goods prices as well (Papademos, 2007, p. 4, Pepper and Olivier, 2006). In the same vein, Sousa and Zaghini (2006) argue that global aggregates are likely to internalize cross-country movements in monetary aggregates - due to capital flows between different regions - that may make the link between money, inflation and output more difficult to disentangle at the country level. Giese and Tuxen (2007) stress the fact 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 asset and goods price inflation.

Some critics might argue that global liquidity, as measured in one currency, can only change in quantitative terms if one assumes a fixed exchange rate system worldwide. Note, however, that international liquidity spill-over effects may occur regardless of the exchange rate system. Under pegged exchange rate regimes official foreign exchange interventions result in a transmission of monetary policy shocks from one country to another. In a system of flexible exchange rates, the validity of the "uncovered interest rate parity" relationship should in theory prevent cross-border monetary spill-overs.

According to this theory, the expected appreciation of the low-yielding currency in terms of the high-yielding currency should be equal to the difference between interest rates in the two economies. However, the enduring existence of carry trades can be taken as evidence that exchange rates diverge from fundamentals for lengthy periods, as the exposure of a carry trade position involves a bet that uncovered interest rate parity does not hold over the investment period. Note as well that exchange rates might quite rarely be considered as truly flexible across our estimation period anyway, as, for instance, Reinhart and Rogoff (2004) classify only 4.5% of the exchange rate regimes under their investigation as "freely floating".

The concept of “global liquidity" has attracted growing attention in the empirical literature in recent years. One of the first studies in this field is Baks and Kramer (1999) who use different indices of liquidity in seven industrial countries to explore the dimension of the relationship between liquidity and asset returns. The authors find evidence that there are important common components in G7 money growth and that an increase in G7 money growth is consistent with higher G7 real stock returns and lower G7 real interest rates.

Recently, a number of studies have applied VAR or VECM models to data aggregated on a global level. Important contributions include Rüffer and Stracca (2006), Sousa and Zaghini (2006) and Giese and Tuxen (2007). These studies find significant and distinctive reaction of consumer prices to a global liquidity shock. In contrast, the relationship between global liquidity and asset prices is mixed. In the study by Rüffer and Stracca (2006), e.g., a composite real asset price index that incorporates property and equity prices does not show any significant reaction to a global liquidity shock. Giese and Tuxen (2007) find no evidence that share prices increase as liquidity expands; however, they cannot empirically reject cointegrating relationships which imply a positive impact of global liquidity on house prices.

3. The price adjustment process

Notwithstanding the fact that the focus of this paper is clearly on the empirical aspect of the subject, we will address some theoretical issues regarding the linkages between money growth (and thus, liquidity) and asset prices. While there is a vast amount of literature available on the impact of commodity price developments on the macroeconomy (Cody and Mills, 1991) and on the role of fundamental factors other than monetary policy for commodity price developments (Hua, 1998), studies specifically dealing with the impacts of monetary policy on commodity prices are evenly distributed over the last decades but - especially for countries except the US - still surprisingly scarce.[1]

Over the last three decades the role of commodity prices in setting monetary policy has been debated among economists (Angell, 1992). We would like to highlight three important main strands of this literature which also play a major role in our investigation. First, Barsky and Kilian (2002) have looked at the role of monetary fluctuations in explaining oil and consumer prices in the 1970s. They argue that major oil price hikes were not the causal mechanism which triggered the stagflation of the 1970s, since any theoretical presumption that oil supply shocks are stagflationary and corresponding robust empirical evidence for this is absent. In contrast, Barsky and Kilian demonstrate that monetary expansions and contractions have the potential to generate stagflation of realistic magnitudes even if supply shocks are not relevant. According to their results, monetary fluctuations contribute to trace the historical pattern of the movements of prices of oil and other commodities and, above all, the boost of the prices of industrial commodities that preceded the 1973/74 oil price increase. Thus, they are able to deliver a persuasive explanation of the striking coincidence of major oil price increases and worsening stagflation.

Second, one of the main combatants in the field, Jeffrey A. Frankel (1986), has contributed a kind of overshooting theory of commodity prices. This piece heavily draws on Dornbusch's (1976) theory of exchange rate overshooting. 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 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. 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),

[1] It has been argued above that commodity prices might represent an early indicator of the current state of the economy because they are usually set in continuous auction markets with efficient information (Cody and Mills, 1991). Hence, some researchers as, for instance, Christiano et al. (1996) act for the inclusion of commodity prices as an explanatory variable in monetary VAR models.

Boughton and Branson (1990, 1991) and Fuhrer and Moore (1992). However, our contribution differs from these papers with respect to the way of modeling and the empirical methodology.

Third, there is a strand of literature which turns the causality of its research interest on its head and checks for the impact of commodity price developments on the conduct of monetary policy. For instance, Bhar and Hamori (2008) empirically investigate the information content of commodity futures prices for monetary policy. They employ a cross correlation function approach to empirically analyze the relationship between commodity futures prices and economic activity as, for instance, consumer prices and industrial production. They come up with the result that commodity prices can serve as suitable information variables for monetary policy. This study also clearly supports the view taken by Bernanke et al. (1997) who take a look at the oil price shocks to analyze the role of monetary policy in postwar U.S. business cycles. They find that an important part of the effect of oil price shocks on the economy results not from the change in oil prices, per se, but from the tighter monetary policy resulting from the change in oil prices. In the same vein, Awokuse and Yang (2003) claim that commodity price indices serve as important information variables for the conduct of monetary policy because they represent signals of future movements in macroeconomic variables.

However, 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 (Bessler, 1984, Pindyck and Rotemberg, 1990, and Hua, 1998) 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. However, we would like to argue in this paper that this controversy can only be settled as a matter of empirical testing.

Finally, a more general strand of literature investigates the impact of monetary policy on more generally defined asset price developments. One example is Congdon (2005) who investigates the relationship between money supply (specified as broad money) and asset price booms and finds empirical evidence in many cases. For instance, he analyses the portfolio management of (other) financial institutions like pension funds. There, he finds evidence in favor 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,” suggesting that asset prices rise in order to restore the money/asset ratio. 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 materialized as additional quantity and not in price increases in recent years. In the following, we will therefore present a simple model of price adjustment for illustration purposes.

Some insights into the relationship between money, commodity prices, and consumer prices can be derived from the dynamic price adjustment to a liquidity shock across the commodity sector and the goods market. In the short-term, an expansionary monetary policy providing the markets with ample liquidity may trigger an immediate price reaction in the commodity sector, but a more subdued price reaction in the consumer goods market. Over time, however, consumer prices also adjust to the new equilibrium by proportional changes of the consumer price level. In other words, it is plausible to argue that in the long term changes in money supply do not lead to any real effects in money or output. As will become clear below, the possibility of different dynamic adjustments of commodity prices and consumer prices to a monetary shock may also provide an explanation for the recent shift in relative prices between commodities and consumer goods. In order to formalize these considerations, the quantity theory of money might serve as a starting point:

$$m_t v_t = p_t y_t \qquad (1),$$

where m denotes the money stock, v represents the velocity of money, and p and y stand for the price level and real output, respectively. Equation (1) is simply an identity and is valid for all time periods t. Money can be spend either for commodities (y^{COM}) or on consumption goods (y^{CPI}) with prices p^{COM} and p^{CPI}, respectively. The distinguishing features of y^{COM} and y^{CPI} are different price elasticities of supply.[2] On the one hand, commodities are generally assumed to be restricted in supply and cannot be easily ex-

[2] See Browne and Cronin (2007) for a similar model on the relationship between commodity prices, consumer prices and money.

panded, and with high transaction expenses due to transportation costs. Hence, the elasticity of commodity supply vis-à-vis commodity price changes should be quite limited. On the other hand, consumption has infinite price elasticity so that additional demand can be satisfied without any price increase. This assumption is based on the recent developments in international trade. The emergence of low-cost producers in emerging markets and developing countries may have prevented firms from increasing consumer prices in response to a liquidity shock while supply in commodity markets was subject to natural constraints. The general price level is then a weighted combination of the prices of both goods:

$$p_t = \lambda p_t^{COM} + (1-\lambda) p_t^{CPI} \quad (2),$$

with 0<λ<1. Similarly, output consists of the production of both commodities and consumer goods:

$$y_t = \lambda y_t^{COM} + (1-\lambda) y_t^{CPI} \quad (3).$$

In the following, the effects of a one-off increase (of μ percent) in money supply in period t+1 are analyzed against this background. Assuming that v is constant and has a value of one, the relationship between money and the general price level in period t+1 can be written as follows:

$$(1+\mu) m_t = p_{t+1} y_{t+1} = (1+\mu) p_t y_t \quad (4).$$

Due to high competition in international goods markets and the vast supply of cheap labor in many emerging regions of the world, which weighs heavily on the prices of manufactured goods, consumer price inflation remains unaffected by the increase in aggregate demand:

$$P_{t+1}^{CPI} = P_t^{CPI} \quad (5).$$

Rather, the liquidity shock fully translates into an increase in output:

$$y_{t+1}^{CPI} = (1+\mu) y_t^{CPI} \quad (6).$$

By contrast, commodities are short-run supply restricted, which drives prices upwards as a result of the liquidity shock, but keeps output in the commodity sector constant:

$$p_{t+1}^{COM} = (1 + \mu)p_t^{COM} \quad (7),$$

$$y_{t+1}^{COM} = y_t^{COM} \quad (8).$$

Combining equations (4) to (9), the money-price relationship in period t+1 can be described as follows:

$$\begin{aligned}(1 + \mu)m_t &= [(1 + \mu)\lambda p_t^{COM} + (1 - \lambda)p_t^{CPI}][\lambda y_t^{COM} + (1 + \mu)(1 - \lambda)y_t^{CPI}] \\ &= (1 + \mu)p_t y_t \quad (9).\end{aligned}$$

In the long term, however, the theoretical proposition of long-run neutrality must hold, i.e., the increase in money supply affects prices without changing long-run equilibrium real values:

$$p_{t+2}^{CPI} = (1 + \mu)p_t^{CPI} \quad (10)$$

$$y_{t+2}^{CPI} = y_t^{CPI} \quad (11)$$

$$p_{t+2}^{COM} = (1 + \mu)p_t^{COM} \quad (12)$$

$$y_{t+2}^{COM} = y_t^{COM} \quad (13)$$

$$(1 + \mu)m_t = p_{t+2}y_{t+2} = (1 + \mu)p_t y_t \quad (14).$$

Equations (1) to (14) illustrate the price-quantity changes in the commodity and consumer goods markets when aggregate demand changes. On the goods market, one would expect an increase in the production of consumer goods if the demand for consumer goods increases as a result of a positive liquidity shock. By contrast, commodity supply is insensitive to price changes and thus the additional demand for commodities is fully reflected in an increase of commodity prices. In the long term, the neutrality of money holds, such that any change in the money supply is met with a proportional change in the price level that keeps real money and real output in both sectors unchanged.

Figure 1 illustrates (in an extreme form) the price-quantity changes as a result of a monetary expansion in markets with high (left graph) and low (right graph) price elasticity of supply. The aggregated supply of price elastic goods S_e in the short run (SR) is characterized by infinite price elasticity so that additional demand triggered by a liquidity shock (from D_{e1} to D_{e2}) can be satisfied without any price increase. Consequently, the liquidity shock translates into an increase in output achieving a new short-term equilibrium at p_{e1}. In contrast, goods characterized by restrictions in supply (S_i) cannot be expanded easily and are thus quantity insensitive to a monetary expansion. Additional demand (shift from D_{i1} to D_{i2}) is then fully reflected in a rise of commodity prices to p_{i1}.

In the long run (LR), prices will also react on the price elastic good market as the well-documented neutrality of money holds; any change in money supply is met with a proportional change in the price level that keeps real money and real output in both sectors unchanged (at p_{e2} and p_{i2}).

Figure 1 about here

The possibility of different dynamic adjustments of price elastic and inelastic goods to a monetary shock may provide an explanation for the recent upward shift in relative prices between assets and consumer goods. This assumption can be well motivated with developments in international trade. Due to high degree of competition in international goods markets and vast supply of cheap labour in many emerging markets around the world, which weighs heavily on the prices of manufactured goods, in the short-term goods prices remain unaffected by the increase in aggregate demand. Only in the long-term, increasing capacity utilization will translate into higher wages, putting upward pressure on prices.

In contrast, assets such as commodities are generally assumed to be restricted in supply. A number of constraints in the commodity market such as finite supply prevent producers in the commodity market from adjusting quantities to short-term price incentives. Moreover, as argued by Browne and Cronin (2007), the price adjustment process in commodity markets is relatively fast because participants are more equally empowered with more balanced information and resources than their consumer goods counterparts. Being auction-based traded in markets with efficient information, commodities could be characterized as flexible goods in contrast to consumer goods. This enables them to react quickly to changes in monetary conditions. Thus, additional demand for commodities is immediately reflected in a rise of commodity prices, so that in response to a money supply shock, commodity prices could also overshoot their long-run equilibrium compensating for the laggard movement in consumer prices (Frankel and Hardouvelis, 1985, and Frankel, 1986). Frankel formalizes his arguments by applying Dornbusch's (1976) famous overshooting model on commodity prices and monetary policy. Hence, commodity prices might influence consumer prices through a money-driven overshooting and the deviation has explanatory power for subsequent consumer price inflation.

4. Empirical analysis

4.1. Data description and aggregation issues

In our CVAR analysis, we make use of quarterly time series ranging from Q1-1970 to Q2-2008 for the United States, the Euro Area, Japan, the United Kingdom, Canada, Korea, Australia, Switzerland, Sweden, Norway, and Denmark. Hence, in our analysis 69.7% of the world GDP in 2007 and presumably a considerably larger share of the global financial markets are represented.[3] For the aforementioned countries, we have collected data for real GDP (Y), the consumer price inflation (CPI), a broad monetary aggregate (M) and two commodity price indices. The selected monetary aggregate is M2 for the U.S. and Japan, M3 for the Euro Area, and mostly M3 or M4 for the other countries. The data are taken from the IMF, the BIS, Thomson Financial Datastream and the EABCN database and are seasonally adjusted where available or treated with the X12-ARIMA procedure.

The first commodity price index we take into account in our analysis is the Commodity Research Bureau (CRB) index. The CRB provides an encompassing gauge of price trends in commodity markets because the most important 19 commodities are involved in this index. These markets are presumed to be amongst the first to be influenced by changes in economic conditions and would, therefore, be expected to be sensitive to developments in the monetary environment. It consists of energy (39%), softs/ tropicals (21%), grains/ livestock (20%), and industrial/ precious metals (20%). Along with this most broadly defined CRB index, a major division of the index, the Raw Industrials

[3] Own calculations based on IMF data.

(CRBRI) index, is used for robustness analysis. It comprises raw industrial materials/ metals but excludes the volatile food and energy parts. 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.

We start with aggregating the country-specific time series to produce a global series, strictly following the guidelines provided by Beyer et al. (2000) and applied by Giese and Tuxen (2007) in the same context. First, we calculate variable weights for each country by using PPP exchange rates to convert nominal GDP into a single currency.[4] Hence, the weight of country i in period t is given by:

$$w_{i,t} = \frac{GDP_{i,t} e_{i,t}^{PPP}}{GDP_{agg,t}} \quad (15).$$

Secondly, we start with the growth rates of the variable in the domestic currency and amalgamate them to global growth rates by applying the weights calculated above:

$$g_{agg,t} = \sum_{i=1}^{11} w_{i,t} g_{i,t} \quad (16).$$

Finally, aggregate levels are then obtained by choosing an initial value of 100 and multiplying with the computed global growth rates. This gives the level of each variable as an index:

$$index_T = \prod_{t=2}^{T} \left(1 + g_{agg,t}\right) \quad (17).$$

This method is applied to all variables except the commodity price indices, which already represent price developments at a global level.

With respect to the monetary aggregate which plays a central role in our analysis, this method lowers the bias resulting from different national definitions of broad money which obviously exist. Forming a simple sum of national monetary aggregates – as often conducted in the related literature - would underrepresent countries with narrower

[4] 1999 is our base year for the PPP exchange rates.

definitions of the monetary aggregate and vice versa. Using this methodology we also avoid the so-called 'dollar bias,' which results from converting national monetary aggregates with actual exchange rates into U.S. dollar and constructing a simple unweighted sum to obtain global money. For instance, the sharp fall of the dollar between 1985 and 1988 or 2000 to mid 2008 would result in an overestimation of global monetary growth.

4.2. Econometric framework and univariate properties of the data

The econometric framework applied is a cointegrated vector autoregressive (CVAR) model which allows us to model for the impact of monetary shocks on the economy while taking care of the feedback between the variables.
The basic representation is the p-dimensional vector autoregressive model with Gaussian errors ($\epsilon_{it} \sim iidN(0, \Omega)$):

$$X_t = A_1 X_{t-1} + \cdots + A_k X_{t-k} + \Phi D_t + \epsilon_t, \quad t = 1, \dots, T \qquad (18),$$

where X_t are the variables of interest and D_t is a vector of deterministic variables, containing the constant of the model and dummy variables. Reformulating the model in an error correction form allows distinguishing between stationarity that is created by linear combinations of the variables and stationarity created by first differencing:

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k-1} + \Phi D_t + \epsilon_t, \quad t = 1, \dots, T, \qquad (19).$$

The ECM representation of the VAR model provides a favorable 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, categorizing the effects in long (Π) and short (Γ_1) 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 $r \times p$ matrices α and β ($\Pi = \alpha\beta'$). The factorization provides r stationary linear combinations of the variables (cointegrating vectors) and $p - r$ common stochastic trends of the 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. Hence the ECM model allows for the variables in X_t to be integrated of order 1 (I(1)). To access the unit root properties of the individual time series we employ Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) statistics with the natural logs of the variables.

Table 1 about here

Table 1 reports that the levels of all series are clearly non-stationary using standard ADF tests, where the appropriate lag length is selected by the Akaike Information Criterion (AIC) and by the Schwarz Bayesian Criterion (SBC). The Phillips–Perron tests corroborate these results. The first-differences give evidence that most of the series can be assumed to be integrated of order one, whereas the only exception is the CPI data where the indication of the test statistics is mixed. The PP test clearly indicates that CPI is I(1) and this is confirmed by the ADF test with respect to the SBC at the 10% significance level. However, the ADF using the AIC does not reject the null hypothesis of a unit root. It is noted by Greene (2008) and Hamilton (1994) that the ADF test can fail to distinguish between a unit root and a near unit root process and will too often indicate that a series contains a unit root and, furthermore, that the SBC is superior to the AIC for large samples. Given these arguments and the fact that we have a sample size of 154 observations and an ADF statistic with respect to the AIC only marginally greater than the 90 per cent critical value of -3.146, we pursue on the basis that all series are each integrated of order one.

4.3. Lag length selection and diagnostic testing on the unrestricted VAR model

The empirical analysis is conducted with both the broader definition of the commodity index (CRB) and the raw industrials index (CRBRI) giving strong evidence of the findings being robust to varying definitions of commodity price developments.

Specifying the lag length has strong implications for subsequent modeling choices. Choosing too few lags could lead to systematic variation in the residuals whereas if too many lags are chosen comes with the penalty of fewer degrees of freedom (as adding another lag, adds $p \times p$ variables). In macroeconomic modeling it is hard to imagine agents using information that reaches much further back than two to four quarters. In general, a lag length of two is encouraged.

Table 2 about here

The results in Table 2 support the above reasoning for our data. For determining the appropriate lag length for the CVAR model the Schwarz's Bayesian, Akaike's and the Hannan-Quinn information criteria were utilized. In the calculation of AIC, HQC and SBC the number of observations is kept constant for all lags of the endogenous variables, i.e. the CPI, the CRB, M and Y. The Schwarz and Hannan-Quinn criteria suggest lag 2, while AIC suggests lag 3. However, based on the p-values of the final prediction error (FPE) test, lag 2 is a better choice to model the time series of interest. The supplementary figures in Table 3 for the CRBRI prop up choosing a lag length of order 2 for both datasets.

Table 3 about here

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 cause a violation of the normality assumption. The deviation from the normality assumption leads to incorrect statistical inferences. Thus it is important to identify the dates of such shocks and to correct them with intervention dummies (Juselius, 2006). The global data we apply seem overall well behaved in respect to big outliers, thus the necessity of correcting by dummy variables is not given.

Table 4 about here

Table 4 presents univariate and multivariate residual analysis of the unrestricted VAR(2). Based on these results the multivariate LM(1) and LM(2) tests reject autocorrelation in the first and second lag of the residuals. The null of normality for the multivariate model is rejected. Considering the univariate residual analyses, there are deviations from normality in skewness and/or kurtosis for the global money and the global real GDP time series. The results for the CRB index are in line with those for the

CRBRI presented in Table 5. Although the commodity price indexes show high fluctuations especially in the last period of the data sample, according to the univariate statistics there are hardly any evidence of ARCH effects. Even moderate ARCH-effects are not considered as highly problematic as Rahbek et al. (2002) show that the cointegration rank testing is still robust. The formal misspecification tests indicate rejection of multivariate normality due to non-normality in the global money and the global GDP variable. Altogether the VAR(2) model seems to provide a fair description of the information contained in the data.

Table 5 about here

4.4. Rank determination and testing restrictions on the cointegrated VAR model

The cointegration rank is determined according to Johansen LR trace test (Johansen, 1988, 1991, 1994). The results of the LR trace test are presented in Table 6 for the CRB index and in Table 7 for the CRBRI. The trace test statistics fails to reject the hypotheses of $p - r = 2$ common trends and $r = 2$ cointegrating relations on a 1% significance level for the CRB. The choice of a rank of $r = 2$ is even more supported for the dataset including the CRBRI as a measure for commodity price development. In the latter case the trace test statistic fails to reject the hypothesis of 2 cointegrating relations on a 5% 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. Hence Juselius (2006) suggests using additional information, e.g. recursive graphs of the trace statistic and t-values of the adjustment coefficient in order to choose the appropriate rank.

Table 6 and 7 about here

The recursive graphs of the trace test statistic are calculated by $trace(r) = -T\sum_{i=r+1}^{p}\log(1-\hat{\lambda}_i)$. The primary interest is in the time paths of the statistics. The visual inspection is not affected by the scaling of each statistic by the 95% critical value of the trace test distribution. The $trace(j)$ is expected to show upward sloping behavior for $j \leq r$ and for $j > r$ to be constant, as λ_i tends to a constant for $i \leq r$ and to zero for $i > r$.

Figure 2 and 3 about here

Figure 2 and Figure 3 illustrate the recursive estimated trace statistics for both indexes. The graphs based on the concentrated model R1 give support to our choice of a rank of 2 with 2 linearly growing trace statistics and the third being a borderline case. Not including the third cointegrating relation is also supported by examining the t-values of the α coefficients of the third cointegrating vector of the unrestricted VAR(2) model which are all smaller than 2.66 with the CRB index and even smaller (t < 2.29) with the CRBRI data. Hence we do not expect to gain additional information by including the third vector as a cointegrating relation in the model, also in line with our theoretical hypothesis.

Figure 4 and 5 about here

As pointed out above, the appropriate rank is chosen given the evidence from formal trace testing as well as other indicators, e.g., plots of the estimated cointegrating relations. Figure 4 and Figure 5 illustrate the equilibrium errors corrected for short-run effects, depicting the long-run disequilibrium error of the first and second estimated cointegrating relation. The variation in time indicates a fairly stable and stationary pattern supporting the choice for a rank of 2.

Before proceeding with the identification of the long-run structure and the adjustment dynamics a test for variable exclusion is performed. Based on the results in Table 8 and Table 9 illustrating the LR-test for long-run exclusion, we find that with a rank of 2 none of the variables included in the information set can be excluded from the long run relationships. Drawing on the evidence of the above results the empirical analysis is pursued for both commodity index specifications with a rank of 2.

Table 8 and Table 9 about here

A first step towards identifying the long run structure of the CVAR(2) model is done by imposing restrictions on the two cointegrating relations and formally identifying the vectors. The model is estimated with 2 lags and no deterministic terms. The variable vector x_t is defined by:

$$x_t' = [CPI, M, Y, COM]_t \qquad (20).$$

Imposing zero restrictions (e.g. in a formulation of $\beta = H\varphi$, with H being a design matrix and φ a vector of the estimated parameters) yields a first notion of the long run structure.

Table 10 about here

The two cointegrating relationships suggested by theory with CPI being cointegrated with Y and M on the one hand and cointegration occurring between COM and the money and output variable on the other hand are supported by the results of the just identified long-run relations in Table 10. In addition we would expect long run homogeneity between the monetary and the price variables and thus the restriction of the coefficient on M being -1 to be accepted jointly for both cointegrating relations. For each cointegrating vector the normalization is chosen on the price indexes and the commodity prices are restricted to zero in the first long-run relation while the numerical ordering is reversed for the second long-run relation restricting the coefficient on the consumer price index to be zero. The results of the exactly identified restrictions on the first and second cointegrating relation are given as:

$$CPI - 0.909\ [-5.612]\ M + 0.182\ [0.776] Y \ \sim I(0)$$
$$COM - 1.193\ [-5.367]\ M + 0.778\ [2.143] Y \ \sim I(0) \qquad (21).$$

The M coefficient has the expected negative sign and is significant (with t-values in brackets) across the two reported vectors for the CRB commodity index. The coefficients on Y have the correct positive sign but are insignificant in the vector normalized on CPI.

Table 11 about here

According to the theoretical hypothesis of the price adjustment process we would expect long-run proportionality between the monetary aggregate and the respective price index. In addition we would also suppose in line with the theoretical model the output coefficient to be positive and statistically significant. Testing these restrictions gives the following cointegrating vectors:

$$CPI - M + 0.322\ [14.928] Y \sim I(0)$$
$$COM - M + 0.480\ [21.633] Y \sim I(0) \qquad (22).$$

The LR-test statistic of 4.017 with a p-value of 0.134 ($\chi^2(2) = 4.017$ [0.134]) indicates that the restrictions imposed are not rejected and support the theoretical hypothesis. With the economically proposed long run proportionality between prices and money, the coefficients on real GDP are highly significant and with the correct sign supporting the determination of both prices by monetary terms in the long-run. Moreover the empirical findings for the CRB index are supported by the results for the CRBRI depicted in Table 12 and in Table 13 with the over-identifying restrictions being not rejected at a 5% significance level ($\chi^2(2) = 5.194$ [0.075]).

Table 12 and 13 about here

Following the identification of the cointegrating relations the short-run dynamics are analyzed by accounting for the two error correction terms of the long-run relations in a vector error-correction model (VECM). The first error-correcting term (CE_1) is the residual of the cointegrating vector including the consumer price index, the monetary aggregate and the output measure. According to theoretical reasoning above, we would expect the first lag of the CE_1 term to have a negative coefficient, measuring the deviation of the variable from its long-run relation, in an equation where the CPI is the dependent variable.

$$CRB\colon \Delta CPI = -0.0233\ [-5.49] CE_{1(-1)} + 0.0289\ [1.84] CE_{2(-1)} + \ldots$$
$$CRBRI\colon \Delta CPI = -0.0187\ [-5.10] CE_{1(-1)} + 0.000798\ [0.56] CE_{2(-1)} + \ldots \qquad (23).$$

The second error term (CE_2) is the deviation of the commodity price index from its long-run equilibrium, which in turn is expected to have a positive coefficient. Regressing the first difference of the CPI on the lagged error-correction terms and the lags of the changes of the other right-hand side variables included in the CVAR model (which are not reported here in order to keep a parsimonious result presentation) yields the coefficients to have the expected signs. The coefficient on the first error correcting term in eq. (23) is highly significant (indicated by the z-values in brackets) while the second error correcting term is significant on a 10% level for the CRB index and insignificant for the CRBRI.

The analysis is broadly supportive of the model and the theoretical hypotheses. The long-run proportional relationship between global money and prices is underlined by the cointegration analysis. The cointegration error-terms have explanatory power for ensuing consumer price inflation. The deviation of commodity prices from their long-run equilibrium explains subsequent consumer price inflation. By establishing the monetary driven commodity price development in cointegration analysis there is support for deducing that the feedback from commodity prices to consumer prices is a monetary phenomenon.

5. Conclusions and policy implications

So does the inclusion of commodity prices help to identify a significant monetary transmission process from global liquidity to macro variables? And more specifically: Does global liquidity spill over to commodity prices? The main empirical results of our paper in this respect are the following: At a global level, we find further support of the conjecture that monetary aggregates may convey some useful information on variables such as commodity prices which matter for aggregate demand and hence inflation. Thus, we conclude that global liquidity is a useful indicator of commodity price inflation and of a more generally defined inflationary pressure at a global level. To put things differently, we corroborate the results gained by Browne and Cronin (2007) for the US on a global level. Therefore we would like to argue that global liquidity merits some attention in the same way as the worldwide level of interest rates as in the recent hot debate about the world savings versus liquidity glut as the main drivers of the current financial crisis, if not possibly more.

Expressed on a more technical level, this paper has analyzed the relationship among money and commodity prices at a global level. On an OECD level, we find further support to the conjecture that monetary aggregates may convey some useful information on the future development of commodity prices which matter for aggregate demand and hence consumer price inflation. Our empirical results appear to be overall robust since they pass inter alia a series of recursive tests and are stable for varying compositions of the commodity indexes.

To the extent, that our findings do also provide some support for considering commodity price indexes along with other information variables as early indicators of more general inflation and, by this, emphasize rather early claims by Furlong (1989) and Garner (1985).[5] One further advantage might be the more timely availability of commodity price data relative to those on overall prices. Thus, we conclude that liquidity is a useful indicator of commodity price inflation. In our view, one important reason for these quite unbalanced findings is the wide array of different price elasticities of supply.

Against the background, a high level of global liquidity can generally be seen as a threat to future asset price inflation and financial stability.[6] Since global liquidity is found to be an important determinant of commodity prices there might be at least two implications. First, monetary authorities have to be aware of likely spill-overs from commodity to consumer prices. Second, when commodity prices reach an unsustainable level and a potential bubble is created, the implications are risks not only to price stability but also to the economy at large - as seen in the current subprime crisis which apparently has partly spread from the U.S. to other parts of the world.

We also see some implications for policy makers emanating from our empirical results. In the first place, our CVAR analysis indicates that commodity prices might well serve

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5 Bhar and Hamori (2008) and Furlong and Ingenito (1996) focus less on the role of monetary policy in a relationship like presented in our CVAR and more on the signaling or predictive power of commodity prices for consumer price inflation. Accordingly, Sims (1998) and Sims and Zha (1998) emphasize the importance of introducing the commodity price variable in designing monetary policy rules.

6 See the early and continuous publics about the latter by the ECB Observer group as expressed, for instance, in Belke et al. (2004). For details see http://www.ecb-observer.com.

as indicators of future inflationary pressures. Moreover, strong monetary growth might be a good indicator of emerging bubbles in the commodity sector. Hence, asset price movements should certainly play a role in policy.

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Figure 1: Short- and long-run impact of a liquidity shock to price elastic (left-hand side) and price inelastic good (right-hand side).

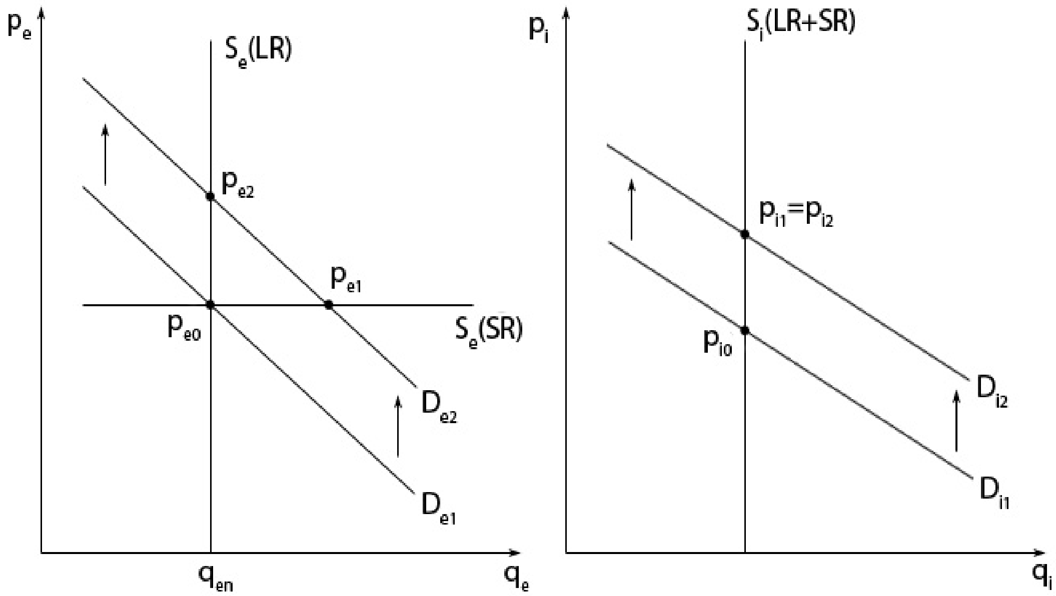


Figure 2: Recursive calculated trace test statistics based on the full and the concentrated model (Base sample 1970:03 to 1975:1) for CRB

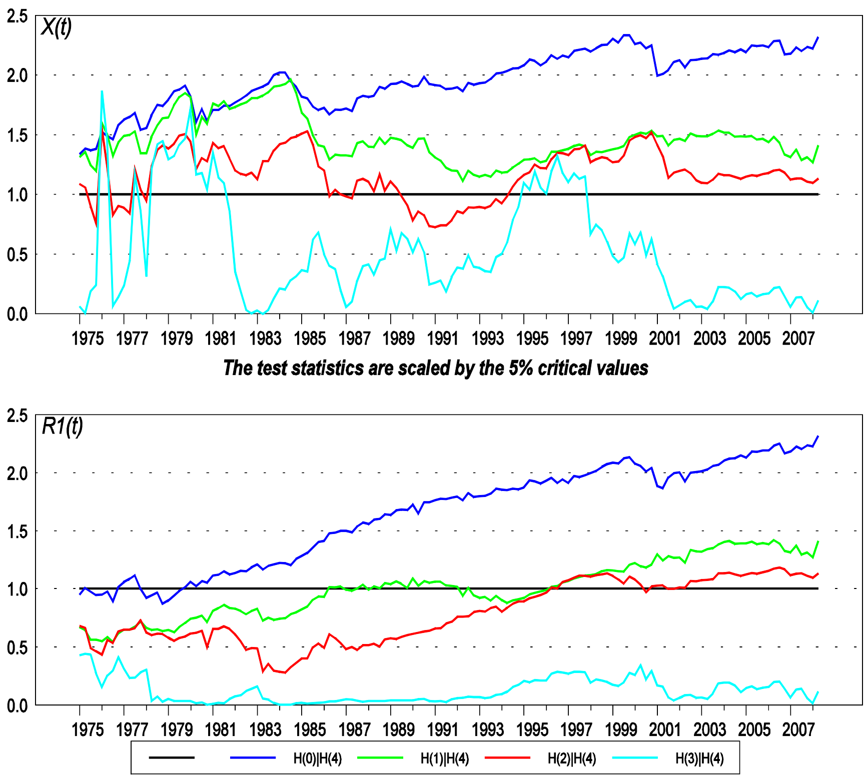


Figure 3: Recursive calculated trace test statistics based on the full and the concentrated model (Base sample 1970:03 to 1975:1) for CRBRI

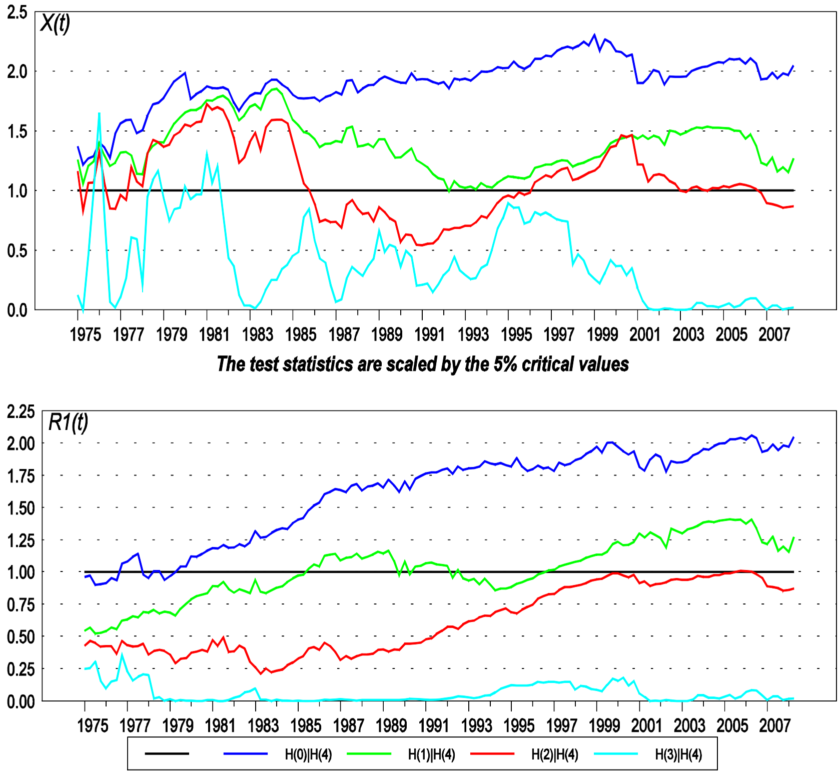
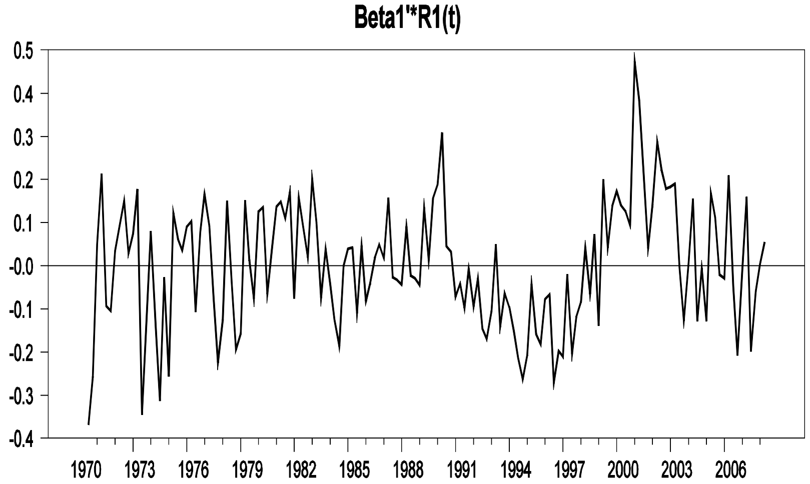


Figure 4: The first and second estimated cointegrating relation $\widehat{\boldsymbol{\beta}}_1' \boldsymbol{R}_{1t}$ and $\widehat{\boldsymbol{\beta}}_2' \boldsymbol{R}_{1t}$ for CRB



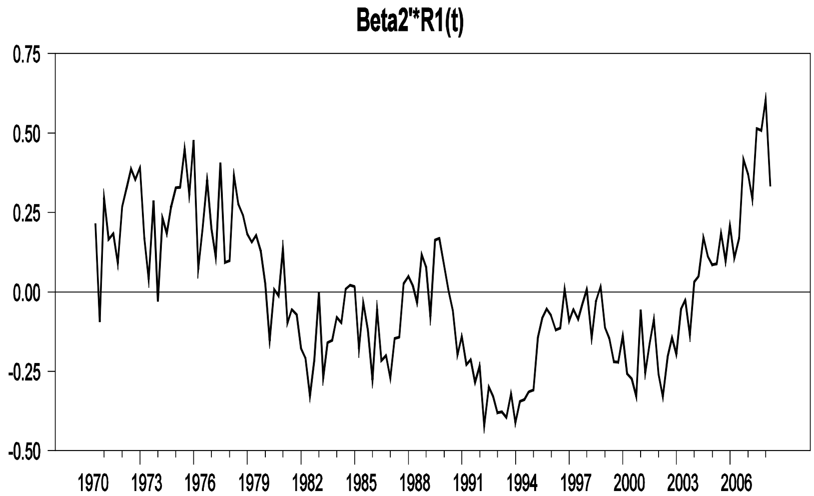
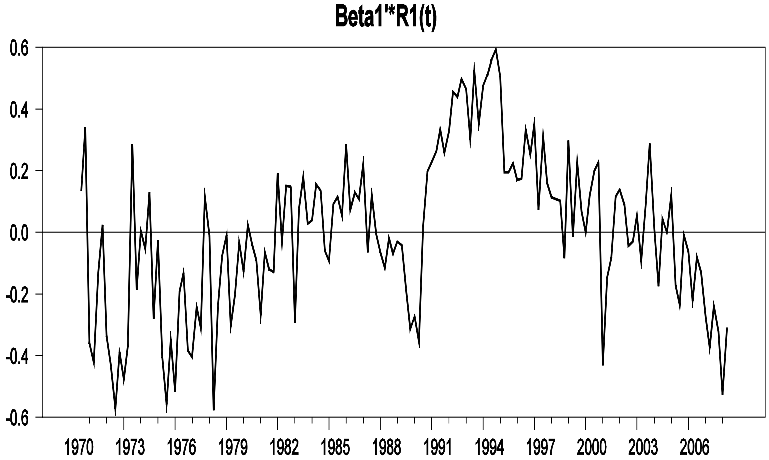


Figure 5: The first and second estimated cointegrating relation $\widehat{\boldsymbol{\beta}}_1' \boldsymbol{R}_{1t}$ and $\widehat{\boldsymbol{\beta}}_2' \boldsymbol{R}_{1t}$ for CRBRI



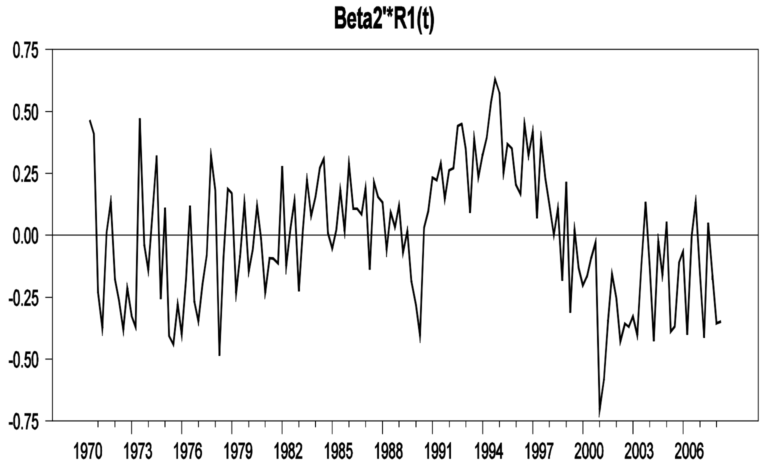


Table 1: Unit root tests

	CPI	CRB	CRBRI	GDP	M
Levels					
ADF (AIC)	-2.999	-2.511	-2.190	-2.627	-3.114
ADF (SBC)	-2.479	-1.612	-2.190	-2.627	-3.114
PP	-1.448	-2.229	-2.221	-2.843	-3.048
First-Difference					
ADF (AIC)	-3.271*	-6.069***	-6.285***	-5.538***	-3.756**
ADF (SBC)	-3.109	-9.134***	-9.003***	-4.137***	-3.756**
PP	-5.647***	-9.716***	-9.234***	-9.469***	-5.032***

Note: Asterisks refer to level of significance: *10%, **5%, ***1%.

Table 2: Lag length selection, CRB

lag	FPE	AIC	HQC	SBC
1	2.5e-17	-26.896	-26.766	-26.575
2	**1.3e-17**	-27.517	**-27.256**	**-26.875**
3	1.3e-17	**-27.518**	-27.126	-26.554
4	1.3e-17	-27.507	-26.986	-26.223

Table 3: Lag length selection, CRBRI

lag	FPE	AIC	HQC	SBC
1	3.6e-17	-26.524	-26.394	-26.203
2	**1.6e-17**	**-27.310**	**-27.049**	**-26.668**
3	1.6e-17	-27.309	-26.917	-26.345
4	1.7e-17	-27.260	-26.738	-25.975

Table 4: Residual analysis - diagnostic testing on the unrestricted VAR(2) model, CRB

Multivariate tests				
Residual autocorrelation				
LM(1)	χ^2 (16) = 20.912 [0.182]			
LM(2)	χ^2 (16) = 3.412 [0.999]			
Test for Normality	χ^2 (8) = 56.808 [0.000]			
Test for ARCH				
LM(1)	χ^2 (100) = 113.696 [0.165]			
LM(2)	χ^2 (200) = 314.037 [0.000]			
Univariate tests				
	ARCH(2)	Normality	Skewness	Kurtosis
ΔCPI	0.587 [0.746]	2.817 [0.244]	0.204	3.438
ΔM	0.914 [0.633]	20.921 [0.000]	0.742	5.633
ΔY	3.577 [0.167]	28.656 [0.000]	-0.192	5.410
ΔCRB	0.972 [0.615]	1.881 [0.390]	-0.258	3.125

Note: *p*-values in brackets.

Table 5: Residual analysis - diagnostic testing on the unrestricted VAR(2) model, CRBRI

Multivariate tests				
Residual autocorrelation				
LM(1)	χ^2 (16) = 35.865 [0.003]			
LM(2)	χ^2 (16) = 14.058 [0.594]			
Test for Normality	χ^2 (8) = 52.321 [0.000]			
Test for ARCH				
LM(1)	χ^2 (100) = 106.729 [0.304]			
LM(2)	χ^2 (200) = 318.248 [0.000]			
Univariate tests				
	ARCH(2)	Normality	Skewness	Kurtosis
ΔCPI	0.668 [0.716]	3.574 [0.167]	0.193	3.549
ΔM	0.993 [0.609]	19.618 [0.000]	0.659	5.385
ΔY	5.110 [0.078]	25.644 [0.000]	-0.202	5.241
$\Delta CRBRI$	2.987 [0.225]	0.185 [0.912]	-0.057	2.958

Note: *p*-values in brackets.

Table 6: Trace test statistics for determination of the cointegration rank for the unrestricted VAR(2) model, CRB

r	p - r	Eigenvalue	Trace	95% Critical Value	P-Value
4	0	0.321	92.914	40.095	0.000
3	1	0.125	34.162	24.214	0.002
2	2	0.085	13.928	12.282	0.026
1	3	0.003	0.464	4.071	0.565

Table 7: Trace test statistics for determination of the cointegration rank for the unrestricted VAR(2) model, CRBRI

r	p - r	Eigenvalue	Trace	95% Critical Value	P-Value
4	0	0.287	82.057	40.095	0.000
3	1	0.124	30.746	24.214	0.006
2	2	0.067	10.671	12.282	0.093
1	3	0.001	0.078	4.071	0.842

Table 8: LR-test of long-run exclusion, $\chi^2(r)$, CRB

r	DGF	CPI	M	Y	CRB	5% Critical Value
1	1	9.727	1.439	9.470	22.665	3.841
		[0.002]	[0.230]	[0.002]	[0.000]	
2	2	15.542	7.767	9.746	27.341	5.991
		[0.000]	[0.021]	[0.008]	[0.000]	
3	3	24.171	11.196	22.079	36.827	7.815
		[0.000]	[0.011]	[0.000]	[0.000]	

Note: *p*-values in brackets.

Table 9: LR-test of long-run exclusion, $\chi^2(r)$, CRBRI

r	DGF	CPI	M	Y	$CRBRI$	5% Critical Value
1	1	5.724	4.942	10.725	15.068	3.841
		[0.017]	[0.026]	[0.001]	[0.000]	
2	2	13.190	11.995	10.870	21.907	5.991
		[0.001]	[0.002]	[0.004]	[0.000]	
3	3	18.406	12.984	21.368	30.195	7.815
		[0.000]	[0.005]	[0.000]	[0.000]	

Note: *p*-values in brackets.

Table 10: The just-identified long-run cointegration relations for r=2, CRB

Just-identified long-run relations				
	CPI	M	Y	CRB
$\hat{\beta}'_1$	1	-0.909 [-5.612]	0.182 [0.776]	0
$\hat{\beta}'_2$	0	-1.193 [-5.367]	0.778 [2.413]	1
	ΔCPI	ΔM	ΔY	ΔCRB
$\hat{\alpha}'_1$	-0.023 [-5.626]	-0.003 [-0.679]	0.009 [1.716]	-0.055 [-1.235]
$\hat{\alpha}'_2$	0.015 [6.274]	0.006 [2.477]	-0.001 [-0.312]	0.015 [0.576]
Combined estimates				
	CPI	M	Y	CRB
ΔCPI	-0.023 [-5.626]	0.003 [1.885]	0.008 [6.213]	0.015 [6.274]
ΔM	-0.003 [-0.679]	-0.004 [-3.034]	0.004 [3.395]	0.006 [2.477]
ΔY	0.009 [1.716]	-0.007 [-3.612]	0.001 [0.569]	-0.001 [-0.312]
ΔCRB	-0.055 [-1.235]	0.032 [1.935]	0.002 [0.130]	0.015 [0.576]

Note: *t*-values in brackets.

Table 11: The over-identified long-run cointegration relations for r=2, CRB

Over-identified long-run relations				
	CPI	M	Y	CRB
$\hat{\beta}'_1$	1	-1	0.322 [14.928]	0
$\hat{\beta}'_2$	0	-1	0.480 [21.633]	1
	ΔCPI	ΔM	ΔY	ΔCRB
$\hat{\alpha}'_1$	-0.019 [-5.279]	0.001 [0.352]	0.009 [1.881]	-0.013 [-0.334]
$\hat{\alpha}'_2$	0.016 [6.035]	0.004 [1.577]	-0.001 [-0.395]	-0.011 [0.380]

Combined estimates				
	CPI	M	Y	CRB
ΔCPI	-0.019 [-5.279]	0.002 [1.463]	0.002 [3.539]	0.016 [6.035]
ΔM	0.001 [0.352]	-0.005 [-3.616]	0.002 [4.583]	0.004 [1.577]
ΔY	0.009 [1.881]	-0.007 [-3.677]	0.002 [3.060]	-0.001 [-0.395]
ΔCRB	-0.013 [-0.334]	0.023 [1.461]	-0.009 [-1.650]	-0.011 [0.380]

Note: *t*-values in brackets.

Table 12: The just-identified long-run cointegration relations for r=2, CRBRI

	Just-identified long-run relations			
	CPI	M	Y	$CRBRI$
$\hat{\beta}'_1$	1	-1.181 [-5.448]	0.606 [1.927]	0
$\hat{\beta}'_2$	0	-1.628 [-5.645]	1.410 [3.370]	1
	ΔCPI	ΔM	ΔY	$\Delta CRBRI$
$\hat{\alpha}'_1$	-0.019 [-5.230]	-0.001 [-0.279]	0.009 [1.944]	-0.048 [-1.119]
$\hat{\alpha}'_2$	0.013 [5.761]	0.004 [1.857]	-0.002 [-0.702]	0.023 [0.850]
	Combined estimates			
	CPI	M	Y	$CRBRI$
ΔCPI	-0.019 [-5.230]	0.001 [0.577]	0.007 [5.299]	0.013 [5.761]
ΔM	-0.001 [-0.279]	-0.005 [-4.102]	0.005 [3.982]	0.004 [1.857]
ΔY	0.009 [1.944]	-0.007 [-4.051]	0.002 [1.465]	-0.002 [-0.702]
$\Delta CRBRI$	-0.048 [-1.119]	0.019 [1.144]	0.004 [0.220]	0.023 [0.850]

Note: *t*-values in brackets.

Table 13: The over-identified long-run cointegration relations for r=2, CRBRI

	Over-identified long-run relations			
	CPI	M	Y	$CRBRI$
$\hat{\beta}'_1$	1	-1	0.328 [14.968]	0
$\hat{\beta}'_2$	0	-1	0.442 [21.134]	1
	ΔCPI	ΔM	ΔY	$\Delta CRBRI$
$\hat{\alpha}'_1$	-0.014 [-4.560]	0.003 [0.982]	0.008 [2.272]	-0.019 [-0.534]
$\hat{\alpha}'_2$	0.014 [5.415]	0.003 [1.173]	-0.001 [-0.438]	-0.001 [-0.023]
	Combined estimates			
	CPI	M	Y	$CRBRI$
ΔCPI	-0.014 [-4.560]	0.000 [0.090]	0.002 [2.939]	0.014 [5.415]
ΔM	0.003 [0.982]	-0.005 [-4.191]	0.002 [4.398]	0.003 [1.173]
ΔY	0.008 [2.272]	-0.007 [-4.076]	0.002 [3.397]	-0.001 [-0.438]
$\Delta CRBRI$	-0.019 [-0.534]	0.020 [1.184]	-0.006 [-1.071]	-0.001 [-0.023]

Note: *t*-values in brackets.