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Monetary Policy, Global Liquidity and Commodity Price Dynamics

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Abstract

This paper examines the interactions between money, interest rates, goods and commodity prices at a global level. For this purpose, we aggregate data for major OECD countries and follow the Johansen/Juselius cointegrated VAR approach. Our empirical model supports the view that, when controlling for interest rate changes and thus different monetary policy stances, money (defined as a global liquidity aggregate) is still a key factor to determine the long-run homogeneity of commodity prices and goods prices movements. The cointegrated VAR model fits with the data for the analyzed period from the 1970s until 2008 very well. 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 indices.

The empirical evidence is in line with theoretical considerations. The inclusion of commodity prices helps to identify a significant monetary transmission process from global liquidity to other macro variables such as goods prices. We find further support of the conjecture that monetary aggregates convey useful information about variables such as commodity prices which matter for aggregate demand and thus inflation. Given this clear empirical pattern it appears justified to argue that global liquidity merits attention in the same way as the worldwide level of interest rates received in the recent debate about the world savings and liquidity glut as one of the main drivers of the current financial crisis, if not possibly more.

JEL-Classification: E31, E52, C32, F42

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

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Monetary Policy, Global Liquidity and Commodity Price Dynamics

by

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Berlin and Essen, February 5, 2010

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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 a long time, consumer price inflation has remained largely unaffected by the strong monetary dynamics in many regions in the world. Over the same time period, however, many countries have experienced sharp but sequential booms in asset prices, such as commodity, real estate or share prices.[1]

In the period from 2001 to mid 2008, for instance, house prices increased by 40 to 60 percent in a number of OECD countries, the CRB commodity price index surged by 105 percent in the same period, and also stock prices more than doubled in nearly all major markets from 2003 to 2008. A similar evolution can be found for oil prices. The oil price was still low in 2001, but the next six years saw a steady increase that tripled the price by the middle of 2007. Subsequently, oil prices continued to rise sharply reaching an all-time high on July 3, 2008, only to be followed by an even more spectacular price collapse.[2] Around the turn-of-year 2008-09, the oil price started to rebound and has now reached values of around $75 which is about twice as much as at the beginning of 2009. Many observers feel that the sequential increase of asset prices is the result of liquidity spillovers to certain asset markets.[3]

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.[4] To investigate the relative importance of these developments, this study tries to establish an empirical link between

[1] See Schnabl and Hoffmann (2007).
[2] See Hamilton (2008).
[3] See Adalid and Detken (2007) and Greiber and Setzer (2007).
[4] The main emphasis in these kinds of studies is on globally aggregated variables, which implies that they do not explicitly deal with spillovers of global liquidity to national variables. The main motivation for this way of proceeding is 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 have a number of authors suggested specific interactions of global liquidity with global consumer price and asset price inflation. See Baks and Kramer (1999), Sousa and Zaghini (2006) and Rüffer and Stracca (2006).

money, interest rates, asset prices and goods prices. For this purpose, we apply the cointegrated VAR (CVAR) framework and analyse the impact of global liquidity on commodity and goods price inflation. While goods prices adjust only slowly to changing global monetary conditions due to plentiful supply of consumer goods especially from emerging markets, asset prices such as commodity prices react much faster since the supply of commodities cannot be easily expanded and new information is relatively fast incorporated in these auction-based traded markets. Thus disequilibria on these markets are generally balanced out by price adjustments.

The paper is structured as follows: Section 2 provides an impression of the global perspective of the monetary transmission process. In section 3, we present an overview of the literature and apply some simple theoretical considerations to illustrate the potential impacts of monetary policy on commodity prices. Section 4 turns to the technical details using the CVAR technique on a global scale and reports on the estimation outcomes. The final section offers conclusions as well as some policy implications of the results.

2. The global perspective of monetary transmission

Both with respect to global inflation and global liquidity performance, available evidence is strong that the global rather than national perspective is more important when the monetary transmission mechanism has to be identified and interpreted.[5] Considering the development of global liquidity over time, the question is often raised whether and to what extent global factors are responsible for it.

A few studies investigate this aspect for the G7 countries and conclude that around 50 percent of the variance of a narrow monetary aggregate can be traced to one common global factor such as the expansionary monetary policy stance of the Bank of Japan during the last few years,[6] which has been characterized 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

[5] For instance, Ciccarelli and Mojon (2005) find that deviations of national inflation from global inflation are corrected over time. Similarly, Borio and Filardo (2007) argue that the traditional way of modelling inflation is too country-centred and a global approach is more adequate.
[6] See Rüffer and Stracca (2006).

the proceeds in currencies with higher interest rates. This kind of capital transaction has impacts on the development of monetary aggregates far beyond the special case of Japan and national borders in general.[7]

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 in 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.[8]

The concept of global liquidity has attracted growing attention in the empirical literature in recent years[9] and there is empirical evidence of the existence of a global business cycle.[10] D'Agostino and Surico (2009) find that forecasts for US inflation based on global liquidity are significantly more accurate than those based solely on domestic data. Some 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 discover significant and distinctive reaction of consumer prices to a global liquidity shock. In contrast, the relationship between global liquidity and asset prices is mixed. For instance, in the study by Rüffer and Stracca (2006), 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.

[7] See Schnabl and Hoffmann (2007).
[8] 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. See Papademos (2007) and Pepper and Olivier (2006). 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.
[9] See IMF (2007).
[10] See Canova, Ciccarelli, and Ortega (2007).

3. Overview of the literature and theoretical considerations

Although the focus of this paper is clearly on the empirical aspect of the subject, we will address some theoretical issues regarding the linkages between interest rates, 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.[11]

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 some important main strands of this literature which also play a major role in our investigation. First, one of the main combatants in the field, Jeffrey A. Frankel (1986), has contributed a kind of overshooting theory of 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 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), 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.

Furthermore, 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 analyse the relationship between

[11] 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.

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 analyse 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.

Some insights into the relationship between money, interest rates, 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. 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 formalise these considerations, we apply the model by Frankel and Hardouvelis (1985) and begin with a money demand equation as starting point:

$$m_t - p_t = y_t - \lambda r_t \qquad (1),$$

where m_t and p_t are the logs of the money supply and the price level, y_t represents the influence of real income, r_t is the short-term interest rate and λ represents the semi-elasticity of the money demand with respect to the interest rate. The commodities market is subject to the arbitrage condition that the expected rate of change of commodity prices $E_t(cp_{t+1} - cp_t)$, minus storage costs sc, is equal to the short-term interest rate:

$$E_t cp_{t+1} - cp_t - sc = r_t \qquad (2).$$

The risk premium is assumed to be either zero or contained in the constant storage costs. For the hypothetical case that commodity and all goods prices in the consumption basket are perfectly flexible, the relative price of commodities and other goods is consequently invariant with respect to monetary developments. The general price level in this situation, $\overline{p_t}$, is proportional to the price of commodities. Substituting (2) into (1) and setting $\overline{p_t}$ equal to cp_t results in

$$m_t - \bar{p}_t = y_t - \lambda(E_t\bar{p}_{t+1} - \bar{p}_t - sc) \qquad (3).$$

Solving for $\overline{p_t}$ gives

$$\bar{p}_t = \left(\frac{1}{1+\lambda}\right)(m_t - y_t) + \left(\frac{\lambda}{1+\lambda}\right)(E_t\bar{p}_{t+1} - sc) \qquad (4),$$

and assuming rational expectations yields

$$E_t\bar{p}_{t+1} = \left(\frac{1}{1+\lambda}\right)E_t(m_{t+1} - y_{t+1}) + \left(\frac{\lambda}{1+\lambda}\right)(E_t\bar{p}_{t+2} - sc) \qquad (5).$$

Substituting (5) into (4), then replace for $E_t\bar{p}_{t+2}$ and continue recursively results in

$$\bar{p}_t = \left(\frac{1}{1+\lambda}\right)\sum_{\tau}^{\infty}\left(\frac{\lambda}{1+\lambda}\right)^{\tau} E_t(m_{t+\tau} - y_{t+\tau}) - \lambda sc \qquad (6).$$

Therefore, $\bar{p_t}$ should be viewed as the present discounted sum of the expected future path of the money supply. Equation (6) could be used directly to interpret the reactions of commodity prices to monetary developments provided that the hypothesis of perfectly flexible goods prices in the economy is correct and adjusting directly in response to monetary conditions is given.

Considering the other setting in which the prices of most goods and services are assumed to be sticky in the short run, this equation cannot be used to indicate the reaction of either the general price level or of commodity prices. Assuming for this situation that the general price level adjusts only gradually over time and only in the long run moves with $\bar{p}_t$, then it can be shown that commodity prices will react in the same direction as $\bar{p}_t$, but will move more than proportionally in the short run:

$$\Delta cp_t = \left(1 + \frac{1}{\theta\lambda}\right) \Delta\bar{p}_t \qquad (7),$$

where θ represents the fraction of the deviation from long-run equilibrium $\bar{p}_t$ that p_t can be expected to close each period. Equation (7) was first developed by Dornbusch (1976) in his famous overshooting model for exchange rate determination. Frankel and Hardouvelis (1985) adopt this to commodity prices to show how the spot price of commodities reacts more than proportional to a sudden permanent change in the money supply, that is, how commodity prices overshoot their long-run equilibrium compensating for the laggard movement in goods prices.[12] In the special case of perfectly flexible adjustment of all prices, θ is infinite and (7) reduces to the aforementioned case in which Δcp_t is equal to $\Delta\bar{p}_t$. Combining equations (6) and (7) results in

$$\Delta cp_t = \left(1 + \frac{1}{\theta\lambda}\right) \Delta \left[\left(\frac{1}{1+\lambda}\right) \sum_{\tau}^{\infty} \left(\frac{\lambda}{1+\lambda}\right)^{\tau} E_t(m_{t+\tau} - y_{t+\tau})\right] \qquad (8).$$

In a money supply process with permanent disturbances to the trend and transitory disturbances to the level, Mussa (1975) has shown that this linear form is the rational one to take for market expectations. As a result, the reaction in commodity prices is linearly related to monetary conditions. Accordingly, the possibility of different dynamic adjustments of price elastic and inelastic goods to a monetary shock may provide an ex-

[12] See also Frankel (1986) for the detailed version of the model.

planation 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 should be characterized as flexible goods in contrast to consumer goods. This enables them to react quickly to changes in monetary conditions. As a result, 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. Consequently, commodity prices might influence consumer prices through a money-driven overshooting and the deviation has explanatory power for subsequent consumer price inflation.

In the following, we check for the empirical evidence of the implied transmission mechanism from global money via global interest rates to global commodity and global goods prices within a Cointegrating VAR framework.

4. Empirical analysis

4.1. Data description and aggregation issues

In our CVAR analysis, we make use of quarterly time series ranging from 1970, first quarter, to 2008, second quarter, for the United States, the Euro Area, Japan, the United Kingdom, Canada, Korea, Australia, Switzerland, Sweden, Norway, and Denmark. By this, our country set represents approximately 70% of the world GDP in 2008 and presumably a considerably larger share of the global financial markets.[13] For the aforementioned countries, we have collected data for real GDP (Y), the consumer price inflation (CPI), the three-month Treasury bill rate (TBR) as the short-term interest rate, broad money aggregates and two commodity price indices. The selected monetary aggregates are M2 for the U.S. and Japan, M3 for the Euro Area, and mostly M3 or M4 for the other countries.[14] By the method described below, we compute the ratio of global nominal money to nominal world GDP (LIQ) as global liquidity indicator, a measure commonly used as a sensor of excess liquidity (see, e.g., Rüffer and Stracca, 2006). The two commodity price indices we take into account in our analysis are the Commodity Research Bureau (CRB) and the CRB Raw Industrials (CRBRI) 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 CRBRI index, is used for robustness analysis. It comprises raw industrial materials/ metals but excludes the volatile food and energy parts.[15] 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.[16] Hence, the weight of country *i* in period *t* is given by:

$$w_{i,t} = \frac{GDP_{i,t} e_{i,t}^{PPP}}{GDP_{agg,t}} \qquad (9).$$

[13] Own calculations based on IMF data.
[14] 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.
[15] In the following, we present mostly the results for the broad CRB index as the robustness analysis yields comparable outcomes using the CRBRI index instead of the CRB. The results are available upon request by the authors.
[16] 1999 is our base year for the PPP exchange rates.

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:

$$g_{agg,t} = \sum_{i=1}^{11} w_{i,t} g_{i,t} \qquad (10).$$

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:

$$index_T = \prod_{t=2}^{T} \left(1 + g_{agg,t}\right) \qquad (11).$$

This method is applied to all variables except the commodity price indices, which already represent price developments at a global level and the aggregation for the short-term interest rate is performed without calculating growth rates. With respect to the measure of global liquidity, 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 under-represent countries with narrower 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 the continuous depreciation from 2001 to 2008 would result in an overestimation of global monetary growth.

Figure 1 about here

Figure 1 displays the time path of the globally aggregated time series under investigation. The CPI series reflects the great inflation of the 1970s until the mid 1980s, followed by modest inflation rates afterwards. Even more moderate CPI inflation started to emerge around the mid 1990s and has persisted until the end of the sample although the liquidity measure expanded heavily in the last decade. Global short-term interest rates started to decrease in 1982 and were at historically low levels since 2001 due to the monetary loosing starting at this time. A closer inspection of the global time series reveals that in recent years global excess liquidity has been accompanied by strong price increases in the commodity markets. The ongoing discussion about the linkage of global excess liquidity and asset price inflation is not at least based on the co-movement in this period. In the following econometric analysis we will examine this co-movement of global liquidity and commodity price inflation more deeply.

4.2. Econometric framework and univariate 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 loosing information on the levels of the data generating process. The cointegrated VAR framework allows avoiding the loss of information by modeling non-stationary data through linear combinations of the levels of the variables in consideration. Thus the dynamic system of time-series variables of the cointegrated VAR approach enables us to model short and long-run dependencies. The basic representation is a p-dimensional vector autoregressive model with Gaussian errors ($\epsilon_{it} \sim iidN(0, \quad)$):

$$X_t = A_1 X_{t-1} + \cdots + A_k X_{t-k} + \Phi D_t + \epsilon_t, \quad t = 1, \ldots, T \qquad (12),$$

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

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k-1} + \Phi D_t + \epsilon_t, \quad t = 1, \ldots, T, \qquad (13).$$

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) 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 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. Therefore the ECM model allows for the variables contained in X_t to be integrated of order 1 (I(1)).

To access the unit root properties of the individual time series used by us we apply Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) test statistics to the natural logs of our variables of consideration, except the interest rate, which is verified in its level.

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 (PP) tests corroborate these results. Looking at the results for the first-differences conveys empirical evidence that most of the series can be assumed to be integrated of order one. However, the only exception is the CPI data for which the empirical realization of the test statistics gives mixed results. The PP test clearly indicates that CPI can be considered as integrated of order 1 (I(1)), an assessment which is confirmed by the ADF test with respect to the SBC at the 10% significance level. However, the ADF test specified according to the AIC does not reject the null hypothesis of a unit root. However, Greene (2008) and Hamilton (1994) observe that the ADF test tends to fail to distinguish between a unit root and a near unit root process and too often indicates that a series contains a unit root. Furthermore, they argue that the SBC is superior to the AIC in the case of a large sample. Given these arguments and the fact that we dispose of a sample size of 154 observations and an empirical realization of the ADF test statistic based on the AIC only marginally larger than the 90 per cent critical value of -3.146, we feel legitimized to continue assuming that all series are each integrated of order one. We now proceed with the lag length selection and diagnostic testing on the unrestricted VAR model.

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

Since all asymptotic results as the choice of the cointegrating rank depend on the appropriateness of specification of the underlying model, we now focus on testing the adequacy of the econometric model used by us.

Specifying the lag length of the VAR has strong implications for our subsequent modeling 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 $p \times p$ variables). In macroeconomic modeling it is hard to imagine agents using information that reaches back much further than two to four quarters. Hence, a lag length of two is generally encouraged. 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 “Schwartz” and “Hannan-Quinn” information criteria.

Estimation of our 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 incorrect statistical inferences. Hence, it is quite important to identify the dates of such shocks and to correct them with intervention dummies (Juselius, 2006). The global data used here seem overall well behaved. However, our formal residual analysis suggests the inclusion of dummy variables to deal with potentially non-Gaussian properties of the residuals.

Table 2 about here

Table 2 reports a 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. We reject the null of normality for the multivariate model due to empirical evidence of deviations from normality in skewness and/or kurtosis for the global liquidity and the global interest rate series. Although the commodity price time series display high fluctuations especially in the last period of the data sample, there is hardly any evidence of 2nd 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 global liquidity and interest rates series. Overall the VAR(2) model seems to provide a reasonable description of the information contained in the data. We now estimate our global CVAR after having determined its rank.

4.4. Estimation and rank determination of the global CVAR

The complex determination of the cointegration rank of the Π-matrix 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 our global CVAR being presented in Table 3. The trace test statistics fails to reject the hypotheses of $p - r = 2$ common trends and $r = 3$ 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 3 about here

The unrestricted estimates of the α-coefficients do not draw clear picture in favour of choosing a rank of 3 by themselves, whereas the graphical inspection of the recursively calculated trace test statistics and stationarity of the cointegrating relations suggest that a rank of 3 seems an appropriate choice. 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 about here

Figure 2 illustrates the recursive estimated trace statistics. The graphs based on the concentrated model R1 render support to our choice of a rank of 3 with 2 linearly growing trace statistics and the third being a borderline case. As will be pointed out below, including the third cointegrating relation is yet favoured by the identification of the long-run structure and our recursive tests of parameter constancy.

Figure 3 about here

Figure 3 illustrates the long-run disequilibrium error of the first, second and third cointegrating relation. The time paths indicate a fairly stable and stationary pattern supporting our choice of a rank of 3.

4.5. Identification of the long-run structure and hypothesis testing of the restricted model

As mentioned in section 4.3., the existence of outliers in the underlying data set leads to autocorrelation and distortion of the residual distribution. In addition to accounting for a deterministic trend in the data by specifying the model including a trend term restricted to the cointegration space, we correct for innovational outliers. More specifically, we include three permanent impulse dummies taking the value one in the given quarter of the respective year and zero elsewhere:

$$D_t = [D7401p, D7802p, D8004p]_t \quad (14).$$

The information set is defined by the variable vector

$$x_t' = [CPI, CRB, LIQ, TBR, Y]_t \quad (15).$$

Table 4 about here

The restrictions for the long-run specification for r=3 and a lag length of k=2 are not rejected with a p-value of 0.309 ($\chi^2(1) = 1.034$) and the results are presented in Table 4.

$$\hat{\beta}_1': CRB - CPI - 3.059\ [-6.071] LIQ + 4.416\ [9.181] TRB + 0.007\ [7.977] t \sim I(0) \tag{16}$$

$$\hat{\beta}_2': TRB + 0.206\ [4.919] Y - 0.337\ [-12.576] CPI + 0.204\ [10.671] CRB \sim I(0) \tag{17}$$

$$\hat{\beta}_3': CRB - 4.057\ [-5.699] LIQ - 0.088\ [-0.904] CPI \sim I(0) \tag{18}$$

The empirically identified long-run structure represented by the cointegrating relations $\hat{\beta}_1'$, $\hat{\beta}_2'$ and $\hat{\beta}_3'$ is guided by economic reasoning. The first cointegrating relation represents the price spread of commodities and consumer goods with the former being characterized as the flexibly adjusting quantity. The deviation of commodity prices from consumer prices is significantly driven by excess liquidity and is negatively related with the interest rate. $\hat{\beta}_2'$ can be interpreted as a Taylor-rule-type relation where the interest rate is positively connected to inflation. Given that the interest rate is inversely reacting to output and commodity prices it seems more appropriate to read the second long-run relation in a sense as a "failing" Taylor-rule-type relation, as the interest rate does not seem to have been adjusted enough in the long-run to account for commodity price inflation and output growth on a global scale. The third long-run relation again conveys empirical evidence of commodity prices positively reacting to global liquidity and consumer price inflation, whereas the latter goes without statistically significance.

Figures 4 and 5 about here

We corroborate the appropriateness of the model in describing the long-run properties of the data by means of the recursively calculated test of beta-constancy and the log-likelihood constancy depicted in Figure 4 and 5. The time path of the tests indicate that overall the model performs well. Apart from identifying the long-run equilibriums, the understanding of the system's structure is enhanced by analyzing the α-vectors. Testing for a zero row of the commodity price variable in the α-matrix corresponding to the identified long-run structure is accepted with a p-value of 0.351 ($\chi^2(4) = 4.426$). Thus the commodity prices do not display error-correcting behavior to the cointegrating relations pointing to a dynamic of the commodity prices that potentially compensates the rather sluggish adjustment of consumer prices. The latter can be characterized as completely endogenous, as the respective test of the corresponding α-vectors is rejected with a p-value of 0.000 ($\chi^2(4) = 45.910$).

Our empirical analysis is broadly supportive of the model and the theoretical hypotheses. The impression of a long-run proportional relationship between global money and prices has been hardened by our 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 within the cointegration analysis framework, we have gained support for deducing that the feedback from commodity prices to consumer prices is a monetary phenomenon.

5. Conclusions and policy implications

The main empirical results of our paper 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. Moreover, we identify a negative relation between the world interest rate and commodity prices as proposed by Frankel and Hardouvelis (1985). Thus, we conclude that global liquidity and the world interest rate are useful indicators of commodity price inflation and of a more generally defined inflationary pressure at a global level.

As a by-product, we are able to identify a Taylor-rule-type relation where the interest rate is positively connected to inflation. Given that the interest rate is inversely reacting to output and commodity prices, it seems more appropriate to read the second long-run relation in a sense as a failing Taylor-rule-type relation, as the interest rate doesn’t seem to have been adjusted enough in the long-run to account for commodity price inflation and output growth on a global scale.

Therefore we would like to argue that global liquidity merits some attention in the same way as the worldwide level of interest rates received in the recent hot debate about the world savings and liquidity glut as the main drivers of the current financial crisis, if not possibly more.

Expressed on a more technical level, this paper has analysed the relationship among money, interest rates and commodity prices on a global scale. At the OECD level, we find further support of the conjecture that monetary aggregates may convey some useful information about 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 indices.

Our findings do also provide some support for considering commodity price indices 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).[17] 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.

Against this background, a high level of global liquidity can generally be seen as a threat to future asset price inflation and financial stability.[18] Since global liquidity is found to be an important determinant of commodity prices and there is long-run homogeneity among commodity and goods prices, there might be at least one implication. Monetary authorities have to be aware of likely spill-overs from commodity to consumer prices. 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 as indicators of future more general 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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[17] 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.

[18] 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.

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Figure 1: Time path of the aggregated global series from 1970:01 to 2008:02 (in logs, except for the interest rate)

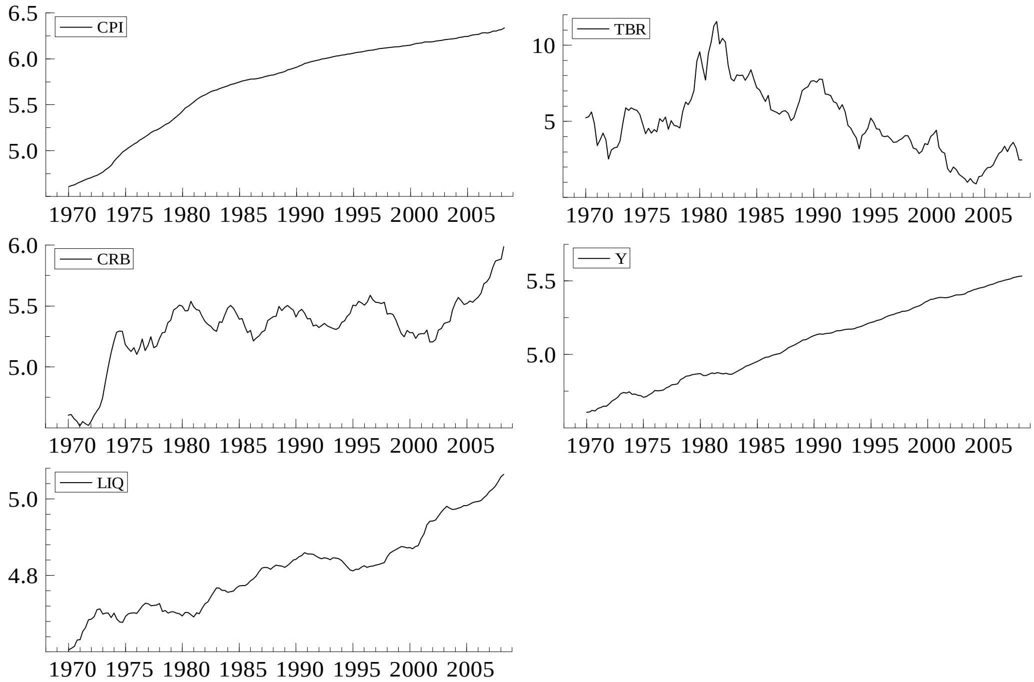

Figure 2: Recursively calculated trace test statistics based on the full and the concentrated model (Base sample 1970:04 to 1978:1)

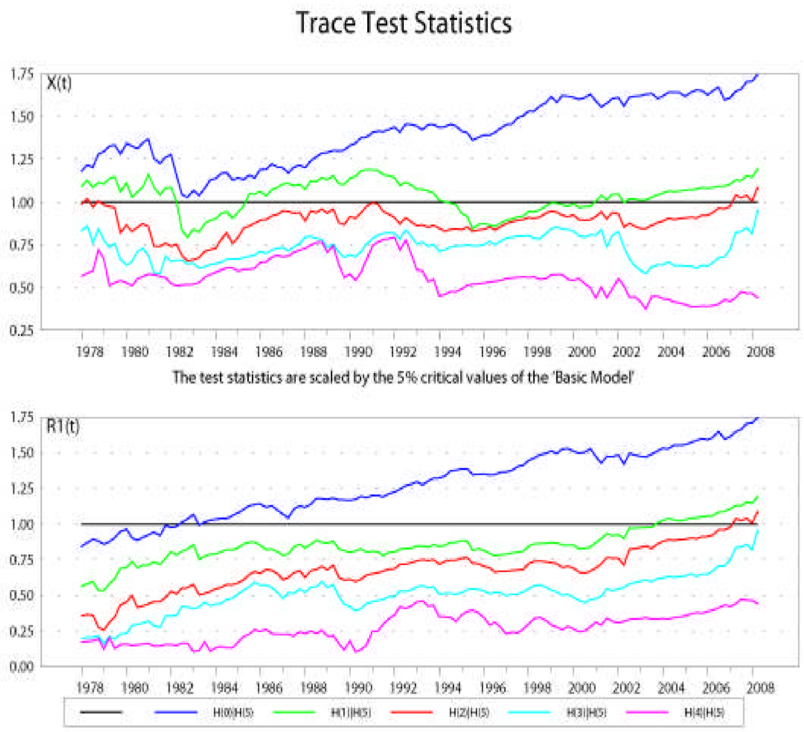

Figure 3: The cointegrating relations
$\widehat{\beta}_1' R_{1t}$, $\widehat{\beta}_2' R_{1t}$ and $\widehat{\beta}_3' R_{1t}$

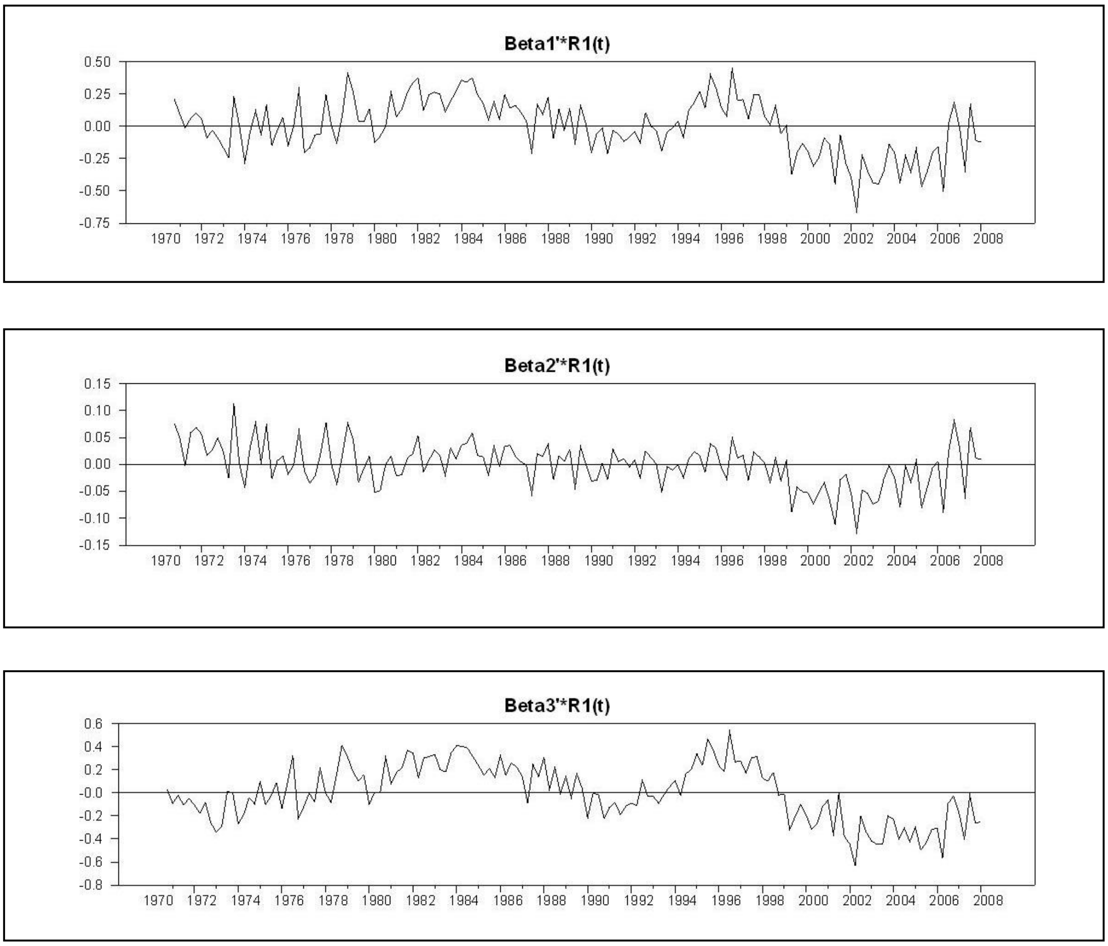

Figure 4: Recursively calculated max test of β constancy
(Base sample 1970:04 to 1978:1)

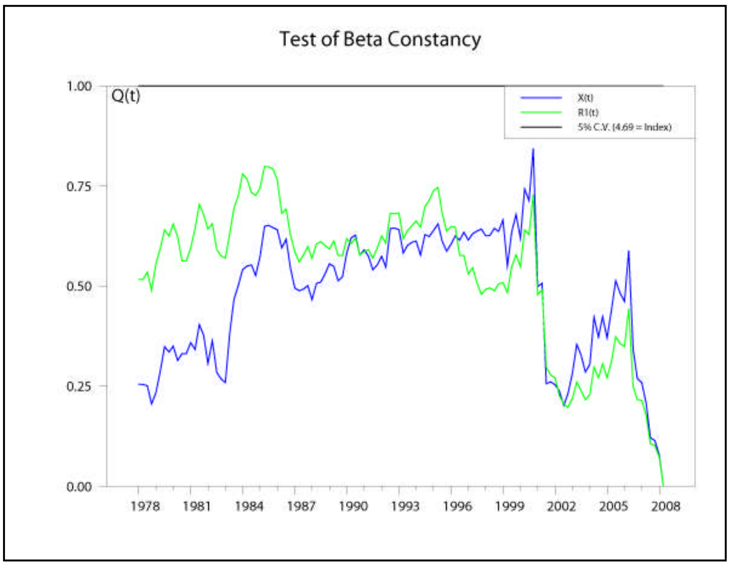

Figure 5: Recursively calculated test of constancy of the Log-Likelihood function (Base sample 1970:04 to 1978:1)

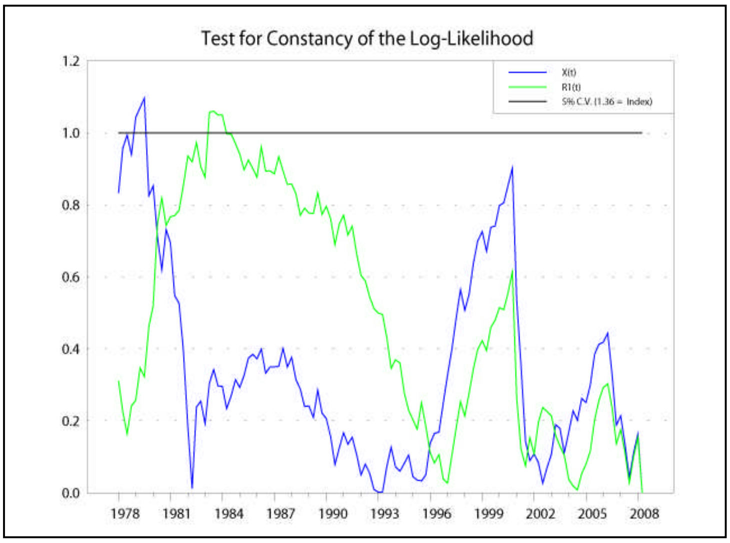

Table 1: Unit root tests

	CPI	*CRB*	*CRBRI*	*LIQ*	*TBR*	*Y*
Levels						
ADF (AIC)	-2.999	-2.511	-2.190	-0.968	-2.131	-2.627
ADF (SBC)	-2.479	-1.612	-2.190	-1.276	-1.834	-2.627
PP	-1.448	-2.229	-2.221	-1.557	-2.109	-2.843
First-Difference						
ADF (AIC)	-3.109	-6.069***	-6.285***	-3.661**	-10.189***	-5.538***
ADF (SBC)	-3.271*	-9.134***	-9.003***	-8.679***	-10.189***	-4.137***
PP	-5.647***	-9.716***	-9.234***	-8.969***	-10.161***	-9.469***

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

Table 2: Residual analysis and diagnostic testing on the unrestricted VAR(2) model

Multivariate tests				
Residual autocorrelation				
LM(1)	χ^2 (25) = 39.310 [0.084]			
LM(2)	χ^2 (25) = 24.409 [0.496]			
Test for Normality	χ^2 (10) = 31.414 [0.001]			
Test for ARCH				
LM(1)	χ^2 (225) = 315.365 [0.000]			
LM(2)	χ^2 (225) = 659.395 [0.000]			
Univariate tests				
	ARCH(2)	Normality	Skewness	Kurtosis
ΔCPI	1.166 [0.558]	4.612 [0.100]	0.424	3.348
ΔCRB	1.563 [0.458]	3.274 [0.195]	-0.308	3.413
ΔLIQ	2.402 [0.301]	6.469 [0.039]	0.267	3.879
ΔTBR	8.511 [0.014]	9.543 [0.008]	-0.017	4.117
ΔY	2.077 [0.354]	4.134 [0.127]	-0.010	3.630

Note: *p*-values in brackets.

Table 3: Trace test statistics for determination of the cointegration rank for the unrestricted VAR(2) model

r	p - r	Eigenvalue	Trace	95% Critical Value	*P*-Value
5	0	0.406	154.687	88.554	0.000
4	1	0.178	76.113	63.659	0.003
3	2	0.135	46.532	42.770	0.019
2	3	0.119	24.621	25.731	0.069
1	4	0.036	5.474	12.448	0.539

Table 4: The long-run cointegration relations

	CPI	CRB	LIQ	TBR	Y	*trend*
$\hat{\beta}'_1$	-1.000	1.000	-3.059	4.416	0.000	0.007
	[NA]	[NA]	[-6.071]	[9.181]	[NA]	[7.977]
$\hat{\beta}'_2$	-0.337	0.204	0.000	1.000	0.206	0.000
	[-12.576]	[10.671]	[NA]	[NA]	[4.919]	[NA]
$\hat{\beta}'_3$	-0.088	1.000	-4.057	0.000	0.000	0.000
	[-0.904]	[NA]	[-5.699]	[NA]	[NA]	[NA]

	ΔCPI	ΔCRB	ΔLIQ	ΔTBR	ΔY
$\hat{\alpha}'_1$	-0.004	-0.193	0.082	-0.036	-0.054
	[-0.404]	[-1.648]	[5.330]	[-2.738]	[-4.083]
$\hat{\alpha}'_2$	0.077	0.568	-0.159	0.057	0.116
	[3.474]	[2.255]	[-4.829]	[1.989]	[4.122]
$\hat{\alpha}'_3$	-0.001	0.050	-0.053	0.023	0.035
	[-0.136]	[0.649]	[-5.239]	[2.618]	[4.005]

Combined estimates

	CPI	CRB	LIQ	TBR	Y	*trend*
ΔCPI	-0.022	0.011	0.017	0.059	0.016	-0.000
	[-5.466]	[6.736]	[2.204]	[2.215]	[3.474]	[-0.404]
ΔCRB	-0.002	-0.027	0.388	-0.285	0.117	-0.001
	[-0.055]	[-1.508]	[4.555]	[-0.948]	[2.255]	[-1.648]
ΔLIQ	-0.024	-0.004	-0.035	0.202	-0.033	0.001
	[-3.984]	[-1.568]	[-3.172]	[5.132]	[-4.829]	[5.330]
ΔTBR	0.015	-0.002	0.018	-0.104	0.012	-0.000
	[2.983]	[-0.920]	[1.901]	[-3.048]	[1.989]	[-2.738]
ΔY	0.011	0.005	0.023	-0.120	0.024	-0.000
	[2.259]	[2.374]	[2.463]	[-3.577]	[4.122]	[-4.083]

Note: *t*-values in brackets.

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