(When) Does Money Growth Help to Predict Euro-area Inflation at Low Frequencies?[a]

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Abstract

Short answer: It helps a lot when other important variables are excluded from the information set.

Longer answer: We revisit claims in the literature that money growth is Granger-causal for inflation at low cycle frequencies. Applying frequency-specific tests to euro-area data in a system with various potentially important variables, money growth is not a significant low-frequency predictor of inflation. A general-to-specific testing strategy reveals a recursive structure where only the unemployment rate and long-term interest rates are directly Granger-causal for low-frequency inflation movements, and all variables affect money growth. We therefore interpret opposite results from bivariate inflation/money growth systems as spurious due to omitted-variable biases. We also analyze the resulting four-dimensional system in a cointegration framework and find structural changes in the long-run adjustment behavior, which do not affect the main conclusions, however.

Keywords: money growth, Granger causality, quantity theory, unemployment

JEL codes: E31, E40

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1 Introduction

Many central banks have ceased to regard the developments of monetary aggregates to be of utmost importance, in contrast to a traditional view that considers inflation to be predominantly a monetary phenomenon. The empirical reason for not focussing on money anymore is that money growth and inflation can be quite disconnected, especially in the short term, see figure 1 for the euro-area example.

The argument in favor of paying close attention to monetary aggregates therefore shifted to the slower low-frequency, longer-run changes of inflation. The intellectual founders of the European Central Bank (ECB) invented the "two-pillar" approach which reserves a whole pillar and thus a "prominent role for money" (ECB, 2004) for the longer run, but which acknowledges that other forces than money growth cause inflation in the short to medium term. Empirical two-pillar Phillips-curve equations adopted this view by adding money growth to reduced-form models of inflation (Gerlach, 2004; ECB, 2004; Beck and Wieland, 2007).

However, from the theoretical side Woodford (2008a,b) argued that modern (New Keynesian) models were already comprehensive without monetary aggregates. He concluded that money is simply not necessary in those models, but whenever it is included a long-run relationship between money (growth) and inflation arises naturally as an additional result. He therefore criticized any prominent role for money as being theoretically unfounded and attributes only a minor role to "monetary statistics ... as indicator variables" (p. 77) for the state of the economy, among many others.

The answer from economists advocating a more prominent role for money was that *empirically* the dynamics of money growth appear to contain more information than in the New Keynesian theoretical framework, for whatever reason. Most notably, in a series of papers Assenmacher-Wesche and Gerlach (2007, 2008a,b, AW&G) have argued that the longer-run (Granger-) causal impact of money growth on inflation can be empirically established if appropriate econometric techniques are used. These results attribute to money growth a low-frequency role in the inflation process and therefore directly support the approach of the ECB.[1]

The present study revisits this latter empirical claim for the case of the euro area by taking a closer look at the result of low-frequency Granger-causality (Granger, 1969, G-causality) running from money growth to inflation. We use a broad set of potential low-frequency predictors of inflation, considering also possible influences coming from goods, labor, and financial markets, and from abroad. We employ essentially the same econometric methods as AW&G, especially low-frequency G-causality tests in a system conditional on other persistent variables.[2]

Our main result is that based on the richer information set in our analysis we cannot confirm AW&G's claim. Money growth first turns out as G-non-causal, and then a search routine reveals unemployment and long-term interest rates as the only significant predictors for the low-frequency movements of inflation. Since we can replicate (qualitatively) AW&G's results in a bivariate dataset with money growth and inflation only, this indicates that their findings suffer from an omitted-variable bias.[3]

After briefly defining our dataset (section 2), we introduce the frequency-domain causality measures and tests, and we report the empirical details of the search routine and the main results of the paper in section 3. In addition, in section 4 we provide a detailed specification of vector error-correction models (VECM), because in this parametric framework we can also check whether the error-correction mechanisms (which correspond to the G-causality at

[1]See also the comments by Nelson (2008). Benati (2009) also confirms a long-run 1:1 relation between money growth and inflation, but seems to rule out the inverse causation direction a priori. Another recent result about the predictive content of money growth for regime switches of inflation is given in Amisano and Fagan (2013).

[2]As an empirical concept, the notion of G-causality rests on predictive power and of course does not necessarily coincide with any structural notion of causality. Nevertheless, we agree with AW&G that G-causality represents an immensely useful tool exactly because it shows us how to obtain better predictions of the variables of interest.

[3]See Lütkepohl (1982) for the theory of omitting variables in Granger-causality tests.

the zero frequency) have changed within the sample. Both sections 3 and 4 contain separate subsections showing that the respective bivariate results from AW&G can be replicated with our dataset. Finally, section 5 concludes.

2 Data

We attempt a fairly comprehensive coverage of potential influences and therefore include more variables than AW&G whose analysis was mainly built on a quantity relationship of money.[4] Defining the term spread as $s = i - i_{short}$ the complete list of included variables along with their labels in the figures and tables is therefore as follows:

Inflation (Δp), M3 growth rate (M3_gr, Δm), unit labor cost growth (ULC_gr, Δulc), labor share $(ulc - p)$, unemployment (u), yield for 10-year bonds (Yield_10yr, i), term spread (Spread, s), real output growth (RealOutp_gr, Δy, openness $(open)$, change of differentials between real and nominal exchange rates (FXdiff_gr, $\Delta E_{real} - \Delta E$), growth of total hours worked (TotalHours_gr, Δh).

A complication arises with the data on total hours h, which are not usually available for the euro area as longer time series. Here we resort to the dataset provided by Ohanian and Raffo (2012): we use the sum of the available series for six euro area countries as a proxy for the whole euro area. Openness is included due to the hypothesis by Romer (1993) and evidence in IMF (2006); Pain et al. (2008) and is measured as the ratio of imports and exports over output. The equilibrium real interest rate is of course unobservable but would also be helpful otherwise, especially given that it is likely to be time-varying and possibly even persistent.

For long-term interest rates we use 10-year government bond yields, for money we use

[4] A possible relationship in the levels of prices, money stocks, and real output is analyzed in the money demand cointegration literature. For euro-area data, Bruggeman et al. (2003) instead use real M3 and inflation (i.e. they impose price homogeneity in the long run) and find that inflation does not react to money demand equilibrium deviations, whereas real M3 does. Holtemöller (2004) analyzes nominal M3 and price levels in a double-integration I(2) framework and finds that both variables are adjusting. More recently Dreger and Wolters (2010) find a stable money demand relationship, but do not report unconditional adjustment estimates.

the M3 aggregate, and inflation is CPI-based. The data are taken from the ECB's area-wide model (AWM) dataset which is extended using available equivalent data from the OECD and the IMF (IFS). We discard the early 1970s to circumvent the problems of dealing with the final years of the Bretton Woods system and the period before the productivity slowdown. Also we do not include the recent financial and economic crisis because the episode of a binding zero lower bound of nominal interest rates represents a different regime for inflation dynamics. Apart from historical interest, the conclusions from our analysis should be relevant again for the time after the end of the current (at the time of writing) period of unconventional monetary policies.

The resulting sample is 1974-2008 with roughly 140 quarterly observations. It should be noted that the sample is dominated by the synthetic AWM data referring to the period prior to the actual formation of the euro area. While the aggregation to a virtual euro area before 1999 may of course be problematic, there is no obvious way around this issue; furthermore that dataset is widely used in policy analysis. Figures 1 through 4 plot all included variables, even starting in 1971.

3 Frequency-domain analysis[5]

3.1 Spectra

Before we apply the frequency-wise causality tests we turn to the fundamental properties of the variables in the frequency domain, i.e. we look at their spectra. The spectrum

$$f_{xx}(\omega) = \frac{1}{2\pi} \sum_{\tau=-\infty}^{\infty} R_{xx}(\tau) \exp(-j\omega\tau), \qquad -\pi \leq \omega \leq \pi \tag{1}$$

[5] All empirical results were produced with gretl, see Cottrell and Lucchetti (2009). The frequency-wise G-causality tests used Breitung's Gauss code that was ported to the Hansl programming language by the authors. The code is available as a gretl function package "BreitungCandelonTest" from the official gretl package server.

is a Fourier transform of the autocovariances $R_{xx}(\tau)$ (with j as the imaginary unit and τ indicating the lag) and can be interpreted as measuring the contributions of different cycle components (at different frequencies ω) for the total variation of the process x_t. The typical spectral shape for many macroeconomic processes is that low frequencies (long-run variations) dominate the spectrum. In the panels of figure 7 we can confirm this phenomenon for eight of the eleven considered variables. Only real output growth has an almost flat spectrum, and so it is clearly stationary, and the spectra of total hours growth and the interest rate spread only have moderate mass in the low-frequency band.

3.2 Granger causality in the frequency domain – the framework

The well-known notion of causality proposed by Granger (1969) rests on predictive power. If (and only if) the variable x_{cause} is Granger-causal (G-causal) for the variable x_{target}, then adding x_{cause} to the available information set gives better predictions of x_{target}. A generalization of this concept was introduced by Geweke (1982), who noted that causal effects can be different at different cycle frequencies. Using the vector moving average (VMA) representation $z_t = \Psi(L)\eta_t$ for $z_t = (x_{target,t}, x_{cause,t})'$ (with L as the lag operator, and η_t is a white noise innovation process) it is useful to partition the lag polynomial $\Psi(L)$ as:

$$\Psi(L) = \begin{pmatrix} \Psi_{11}(L) & \Psi_{12}(L) \\ \Psi_{21}(L) & \Psi_{22}(L) \end{pmatrix} \tag{2}$$

Geweke's causality measure for the frequency $\omega \in (0;\pi)$ is given by:

$$M(x_{cause} \to x_{target};\, \omega) = \log\left(1 + \frac{|\Psi_{12}(\exp(-i\omega))|^2}{|\Psi_{11}(\exp(-i\omega))|^2}\right), \tag{3}$$

An obviously interesting hypothesis to test is that of non-causality at a given frequency ω_0, i.e. that $M(x_{cause} \to x_{target};\, \omega_0) = 0$. Using the fact that $M = 0 \Leftrightarrow |\Psi_{12}(e^{-i\omega})| = 0$, Breitung and Candelon (2006, B&C) showed that this hypothesis is equivalent to two special but linear restrictions in the underlying VAR, and the test of non-causality therefore has standard asymptotics. It also allows to account for further conditioning variables which is desirable given the potential omitted-variable problems mentioned before.

And finally, the B&C test is also applicable to cointegrated systems without having to impose the cointegration restrictions. In this context, note that the B&C test is not applied to the infinite-run frequency zero itself; a test for non-causality at frequency zero can be simply achieved by testing the exclusion of the error-correction terms (zero restrictions on the adjustment/loading coefficients). We will turn to the analysis of the zero-frequency G-causality in the VECM framework in section 4.

3.3 Replicating the Assenmacher-Wesche & Gerlach results

When we analyze only a bivariate dataset comprising inflation and money growth we can replicate the findings by AW&G quite closely. Figure 8 shows that money growth seems G-causal for inflation at low frequencies (left panel) and no G-causality in the other direction. Therefore our different findings are due to the broader information set that we use, not to technical differences or to implementation details.

3.4 Test results and directed graphs

We will use the B&C test as a tool to clarify the possibly complex G-causal relationships between the variables in our dataset. Note that we only report G-causality relationships after the influence of other variables has been taken into account in the system, to avoid spurious findings. Because the underlying information set is quite broad, our aim is to reduce the model to reflect only the truly relevant relationships. Note that we have also experimented with bootstrapped small-sample critical values for some setups, but there were no qualitative changes.

Our empirical strategy to achieve this is as follows:

1. Start with all potentially G-causal variables.

2. Determine significant low-frequency G-causality relations (and their directions) with the B&C test.

3. Drop the following variables (except M3 growth and inflation, which are always retained):

   (a) Those which are completely non-causal (at low frequencies),

   (b) or if there are no non-causal variables, those whose effect on inflation is "most indirect" in the following intuitive graph-theoretic sense, see for example Eichler (2007): The system can be represented as a directed graph where each variable becomes a "node", and the low-frequency causality connections become "edges" that connect the nodes. Since we are dealing with directed graphs, the edges will have one or two arrowheads indicating the causality directions. A causal effect from variable A to B is direct if the two nodes are connected by an edge, with the appropriate arrowhead. An indirect causal effect is given when there is no edge between nodes A and B, but there are appropriate directed edges running for example from node A to node C, and from node C to node B. A causal variable X has the "most indirect" effect on Z if the shortest possible directed path from node X to node Z is longer than from any other causal variable to the target Z.

   (c) If there are more than one most indirect causal variables with respect to inflation, drop the one which is least connected to money growth.

4. Go back to step 2 until no further variables can be dropped.

"Non-causal" variables are those which do not G-cause any other variable in a frequency band ranging from zero to roughly 0.25 (where the cycle periodicity is roughly 25 quarters, or about six years). For all underlying VAR systems a uniform lag length of three was chosen, which in most cases was the recommendation by standard information criteria.

In figures 9 and 10 we report the detailed test results for all frequencies with inflation and money growth as target variables, respectively. (In the rest of the paper we will not report these detailed test results again, but instead we will consolidate the results into figures of directed graphs.) At low frequencies, the only significantly G-causal variable is unemployment for both target variables, but due to the large dimension of the system it remains to be seen whether this may be due to a loss of power of the tests.

The "directed graph" summarizing the information contained in all the similar (non-reported) test plots is shown in figure 11. Such a graph may in principle contain some direct feedback G-causality or indirect circular G-causality paths. In this case it turns out that total hours growth (Δh) is the only completely non-causal variable (at low frequencies). According to our empirical reduction strategy described above, we will therefore drop total hours growth from the dataset.

In the second iteration the G-causality graph in figure 12 does not contain any completely non-causal variables anymore. Note also that more variables than before now appear as G-causal for inflation and money growth: the long-term bond yield affects both, and furthermore inflation also appears as a low-frequency predictor for money growth (but not vice versa). The "most indirect" causal effects come from the spread and from openness, but since the openness variable is a node on a causal path from money to inflation and we do not wish to bias our results against money growth as a predictor, we drop the spread.

After having dropped the spread, in the 3rd iteration (figure 13) among the secondary variables the difference of exchange rate changes (FXdiff_gr) is one of the most indirect effects and is least connected to money growth, and thus we drop it next.

The picture in the next (4th) iteration is still quite complex, although now the only directly significant effect on inflation stems from the unemployment rate. The candidate variables for dropping are the labor share and openness; since the only G-causal effect from money growth works through the labor share, according to our rule we decide to keep that and drop openness instead.

In order to save space, we report the next iterations only textually, without further figures: In the seven-dimensional system of the next step the unit-labor cost growth (ULC_gr) becomes non-causal –along with money growth– and is dropped; afterwards the labor share becomes also non-causal (along with real output growth) and is dropped. Next, in the system with only inflation, money growth, unemployment, the long-term yield, and real output growth remaining, the latter variable only G-causes unemployment at low frequencies, whereas unemployment and the yield directly affect inflation. Thus we finally arrive at a four-dimensional system where from the initial broad information set only unemployment and the yield are kept, apart from the central variables inflation and money growth. The resulting directed graph is shown in figure 15.

Somewhat surprisingly, the resulting low-frequency G-causality graph has a recursive structure without any feedback effects. At the end of the G-causality chain stands money growth, which is significantly affected by all three remaining variables, but which itself is not a low-frequency predictor.

4 Error-correction model estimates

The previous section presented test results concerning the existence of low-frequency G-causality relationships, but remained silent on the quantitative dimension. In this section we provide quantitative models of the long-run determinants of inflation. We employ the standard tool of a vector error-correction model (VECM, i.e. a suitable representation of a cointegrated VAR with K lags):

$$\Delta x_t = \alpha \beta^{*'} (x'_{t-1}, 1)' + \sum_{k=1}^{K-1} \Gamma_k \Delta x_{t-k} + \varepsilon_t \tag{4}$$

Here we have already imposed the deterministic specification of a constant term restricted to the cointegration space, and $\beta^* = (\beta', \beta'_c)'$ is the $(n+1) \times r$ matrix holding the cointegration coefficients β as well as the coefficients of the restricted constant β_c, where n is the dimension of the system and r is the cointegration rank. If $0 < r < n$, the system is truly cointegrated, and α and β will have reduced rank.

Strictly speaking this choice means that we are not analyzing the open frequency band from 0 to 0.25 anymore but that we are analyzing the zero frequency itself. Now, when we model a cointegrated system, we are indeed assuming that the included variables are I(1), i.e. have a spectral peak (singularity) at the zero frequency.[6] In that sense the different methods are theoretically complementary; in practice, however, we expect similar features of the data because in finite samples the difference between low but positive frequencies and the zero frequency are usually blurred.

4.1 The bivariate system of money growth and inflation

In section 3.3 we showed that in a bivariate setup the results of AW&G reappear, namely that money growth seems to be long-run G-causal for inflation. Now we investigate the characteristics of the corresponding bivariate VECM.

The Johansen cointegration test indeed finds cointegration between inflation and money growth, although it does so only at the 10% significance level with asymptotic critical values (see the upper panel of table 1), and thus the bivariate evidence is weak. However, the lower panel reports that it is statistically acceptable to restrict the corresponding cointegration vec-

[6]We can back up this claim by formal unit root tests, but reporting the results yields no value added over what is known in the literature.

tor to a 1:1 relationship. If cointegration is really present, the super-consistent coefficients of this irreducible cointegration vector would enjoy the property that they are asymptotically invariant to extensions of the information set.

Mirroring the results of the bivariate B&C tests in section 3.3, the adjustment coefficients (loadings) also appear to support the hypothesis that inflation adjusts to long-run deviations while money growth is not caused by it. But note that the loading coefficients may be misleading if the system is mis-specified, because they are attached to stationary terms and thus the standard omitted-variables bias applies.

4.2 The system with the low-frequency predictors of inflation

We build on the system reduction analysis in the previous section and consider the four-dimensional, potentially cointegrated, VAR with inflation, money growth, unemployment, and the long-term bond yield. The full sample is still specified as 1974:1-2008:2.

First we run the standard Johansen cointegration test procedure to determine the cointegration rank of the system; in order to avoid the known finite-sample size distortions of this test we apply the Bartlett correction of Johansen (2002).[7] The results are shown in table 2 (upper panel), clearly indicating two linearly independent cointegration relationships in this system at the 1% level of significance. In this case it is even irrelevant whether the Bartlett correction is applied or not.[8]

The estimates of the long-run structure of the system is shown in the lower panel of table 2, where we have applied a number of statistically and economically acceptable coefficient restrictions. First of all, the first cointegration vector is again restricted to be a 1:1 relationship

[7]An implementation of this test procedure has recently been made available as a gretl function package on the standard gretl package server ("coint2finite").

[8]We have also checked a specification including centered seasonal dummies, because even though the series are supposed to be seasonally adjusted there might have been some remaining seasonality. There were no qualitative differences.

between inflation and money growth:

$$\Delta m = 3.71 + \Delta p \tag{5}$$

If interpreted from the perspective of the quantity equation, the constant term in this relationship captures the (differences of the) averages of real output growth and velocity changes.

The second cointegration vector is identified by setting the coefficient of money growth to zero, whereas the remaining coefficients for unemployment and the yield are freely estimated (apart from the normalization of the inflation coefficient):

$$\Delta p = -8.45 + 0.43\, i - 0.90\, urate \tag{6}$$

The fact that both slope coefficients in this trivariate relation are highly significant means that no further bivariate long-run relationships can be recovered from the data. In particular this means that the ex-post real interest rate ($i - \Delta p$) is not empirically stationary, due to the persistent gap between the variables in the central part of the sample that was already visible in 3. The flip side of this statement is that a simple form of the Fisher equation cannot hold, where the unobservable equilibrium real rate would be proxied by a constant or stationary term.

A similar relationship like (6) was found by Benati (2012, after an early version of the present study had been published, which he probably did not notice). He explicitly discounted the evidence in favor of cointegration of these variables, first on theoretical grounds against the background of his model, and secondly by arguing that the Johansen test is oversized in small samples. However, we have applied the cited small-sample corrected version of the test and still get a rejection at the 1% level of significance. Even factoring in some possibly remaining size issues of the test, it seems that the null hypothesis of no cointegration is rejected at conventional significance levels used in empirical economic research, and that is how we interpret the test result in order to learn from the data.

With respect to the equilibrium-correcting behavior of the system, the most important feature is that the unemployment rate and the yield are not adjusting at all. This means that these two variables are weakly exogenous, i.e. not being G-caused in the long run by the other variables, and they drive the system in the long-run by feeding the two stochastic trends into it. Therefore we have a clearcut separation in this four-dimensional system between two long-run driving variables (unemployment and the yield) and two adjusting variables (inflation and money growth). The remaining issue is the detailed adjustment behavior of inflation. Here we see that inflation is not reacting to the equilibrium deviations in the inflation-money growth relationship, which is in quite stark contrast to the results of the bivariate system in section 4.1. Thus the adjustment burden with respect to this first cointegration vector is exclusively borne by money growth. Inflation in turn is the variable which corrects the deviations from the second long-run equilibrium relationship, and quite strongly so. Money growth also reacts to these second equilibrium deviations, even though it is not part of that relationship.

4.3 Changes in the long-run G-causality structure

In figure 16 we display the results of the Hansen and Johansen (1999) test for stability of the cointegration coefficients; for this test the unrestricted cointegration space estimates are used, not the restricted ones reported in table 2. There are nominal rejections of stability around 2004 at the 5% level of significance. However, the maximum test statistics just barely exceed the critical value, and thus we do not interpret this test result as strong evidence against the stability of the long-run relationships.

On the other hand we suspect that there still may have been structural breaks in the dynamic adjustment behavior of the system. Such a break would in general have direct implications for the long-run G-causality patterns between the variables. Therefore in the following we analyze a generalized model where the adjustment matrix α is allowed to change. The timing of the potential break is fixed exogenously at the middle of the sample, i.e. in the year 1991.[9] We also take as given the full-sample estimates of the cointegration relationships because of the weak evidence against stability as discussed before, and also because the subsamples would be too short to estimate separate long-run relationships reliably. We thus take the two-dimensional error-correction terms $ect_t = \hat{\beta}^{*'}(x_t', 1)'$ as given. Let $s_{1991q1,t}$ be a step dummy taking the value 1 in and after 1991, and zero elsewhere. Then we estimate the following system:

$$\Delta x_t = \alpha_1\, ect_{t-1} + \alpha_2 s_{1991q1,t} ect_{t-1} + \sum_{k=1}^{K-1} \Gamma_{break,k} \Delta x_{t-k} + \varepsilon_{break,t} \tag{7}$$

In this extended VECM the parameter α_1 contains the adjustment coefficients for the first half of the sample, while the loadings for the second half are given as $\alpha_1 + \alpha_2$. Without a structural break, the corresponding element of α_2 would not be significantly different from zero. In a first step, the system 7 can be estimated efficiently with OLS. After restricting some elements of α_1 or α_2 to zero, we estimate the system efficiently by feasible GLS (SUR).

The upper panel of table 3 contains the full estimates of model 7. In this unrestricted specification no break terms (α_2) are significant, and the only qualitative difference with respect to the earlier full-sample analysis is that unemployment now also seems to adjust significantly to equilibrium deviations. (A finding which would not change the conclusions with respect to the G-causality relationships between inflation and money growth.)

Then we proceed to apply various restrictions: the yield is still weakly exogenous, i.e. $\alpha_{1;4,1} = \alpha_{1;4,2} = \alpha_{2;4,1} = \alpha_{2;4,2} = 0$, which is clearly still acceptable ($\chi^2_4 = 3.47$, $p = 0.48$); in addition, the adjustment of money growth does not change, $\alpha_{2;2,1} = \alpha_{2;2,2} = 0$, which jointly gives $F_{6,496} = 0.80$, $p = 0.57$; next, the adjustment of inflation to the first cointegra-

[9]This date roughly coincides with German unification, but it also broadly coincides with the end of the disinflation process in the euro area. Having split the sample simply in half to avoid pretesting issues, we leave the question of the driving force behind any potential break for further research.

tion vector –the inflation-money growth relation– does not break, $\alpha_{2;1,1}=0$, $F_{7,496}=0.69$, $p=0.68$; furthermore, the adjustment of unemployment breaks in such a way that unemployment becomes weakly exogenous (not long-run G-caused) in the second subsample, $\alpha_{1;3,1}+\alpha_{2;3,1}=0$, $\alpha_{1;3,2}+\alpha_{2;3,2}=0$, $F_{9,496}=0.561$, $p=0.829$, and finally, inflation does not adjust at all to the first long-run equilibrium, $\alpha_{1;1,1}=0$, yielding $F_{10,496}=0.625$, $p=0.793$. The final result of the long-run structure is reported in the lower panel of table 3.

The most important features of this final specification are the following:

- The adjustment behavior of money growth is essentially unchanged and not subject to the structural break.

- In the recent (post-1991) subsample unemployment and the long-term yield are weakly exogenous and drive the system at the zero frequency, coinciding with the full-sample results. In the earlier subsample, however, the unemployment rate was also equilibrium-correcting.

- For the adjustment of inflation we confirm the previous result that it does not react to deviations from the inflation-money growth relationship. The reaction of inflation to the second long-run relationship, however, becomes quite a bit stronger in the second subsample, rising (in absolute terms) from 0.64 to 1.04. It appears that some of the equilibrium adjustment has shifted from unemployment to inflation.

- With respect to the long-run G-causality of money growth, the second subsample appears qualitatively as the full-sample estimates, with money growth being purely long-run non-causal. In the first subsample, an isolated rise (fall) of money growth would produce a fall (rise) of unemployment through the partial reaction to the negative (positive) deviation from the inflation-money equilibrium; this fall (rise) of unemployment would imply a positive (negative) deviation from the equilibrium relationship linking inflation, unemployment, and the yield, to which inflation would in turn react by falling (rising). However, this channel from money to inflation only runs indirectly through unemployment again, and secondly, the sign of this partial effect is inconsistent with the long-run 1:1 relationship between the two variables. Therefore the full system dynamics would still have to be taken into account for long-run inflation predictions even in the first subsample.

Also note that the negative reaction of money growth to the second cointegration relationship involving unemployment and long-term interest rates may lead to short-run "perverse" dynamics of money growth. For example, after a positive shock to inflation, money growth would at first shrink and thereby the difference $\Delta p - \Delta m$ would be amplified. Afterwards however, the combined reactions of inflation and money growth would restore the two long-run equilibria.

5 Conclusions

The main result of this paper is that the predictive content of money growth for low-frequency movements of inflation in the euro area vanishes once other driving variables are taken into account. Secondly, we found that unemployment and long-term interest rates are the actual low-frequency Granger-causal variables.[10]

Therefore there appears to be an omitted-variable bias in studies such as Assenmacher-Wesche and Gerlach (2007, 2008a) (or with a different method Amisano and Fagan, 2013) who found money growth to be leading inflation at low frequencies. Instead of supporting a prominent role for monetary stance analysis for the longer run, our findings appear more compatible with the skeptical view of Woodford (2008b).

[10]Other considered variables included growth of unit labor costs, hours worked, real output, and of the gap between nominal and real exchange rates; the level of the spread, a trade openness indicator, and the labor share. In a reduced bivariate dataset we could also replicate the result by that the Granger causality at low frequencies misleadingly appears to run from money to inflation.

The positive low-frequency effect of long-term interest rates on inflation is plausible, because changes of long-term interest rates likely signal movements of long-run unobservable inflation expectations which later materialize in observed inflation rates. In general, the nucleus of the structure with interest rates, unemployment, and inflation forming a cointegrated system suggests an interpretation along the lines of the Farmer monetary model (Farmer, 2013), which rests on hysteresis and multiple steady-state equilibria.[11] On the other hand, while it seems difficult to reconcile our empirical results with a New Keynesian model, our main research question is not affected by any ranking of these theoretical models, as money balances or their growth do not appear in either of them.

Checking for structural shifts in these patterns, we found that unemployment used to be (and afterwards ceased to be) long-run Granger-caused in the first part of the sample, and that in the second part of the sample inflation reacts more strongly to disequilibria with respect to long-term interest rates and unemployment. The main finding that money growth is largely irrelevant as a predictor for long-run inflation movements remained intact.

Our results do *not* mean that inflation and money growth are unrelated. Indeed we could confirm that a bivariate equi-proportional (1:1) long-run relation between money (M3) growth and inflation in the euro area is compatible with the data. Nor would monetary aggregates be non-causal for inflation under all circumstances. In a hypothetical scenario of large and exogenous "helicopter drops" of money it would of course be expected that inflation reacts. Also, we do not claim that our analysis is applicable to the special situation of the great recession since 2009, with a binding zero lower bound on nominal interest rates and unconventional monetary policies, which we believe to represent a different regime. But after the

[11] For the USA Beyer and Farmer (2007) had also found nominal interest rates, unemployment, and inflation to constitute a cointegrated system. The difference here is that for the euro area the only bivariate cointegration vector is the add-on relationship between inflation and money growth. Within the trivarate set of nominal rates, unemployment, and inflation we only find a single cointegration relationship. The non-stationary real interest rate in our dataset prevents a higher cointegration rank. Providing an explanation for the non-stationarity of the real interest rate is beyond the scope of this paper, but one candidate reason would be the different demography in continental Europe.

return to standard monetary policy operations our findings would again be relevant.

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Figures and Tables

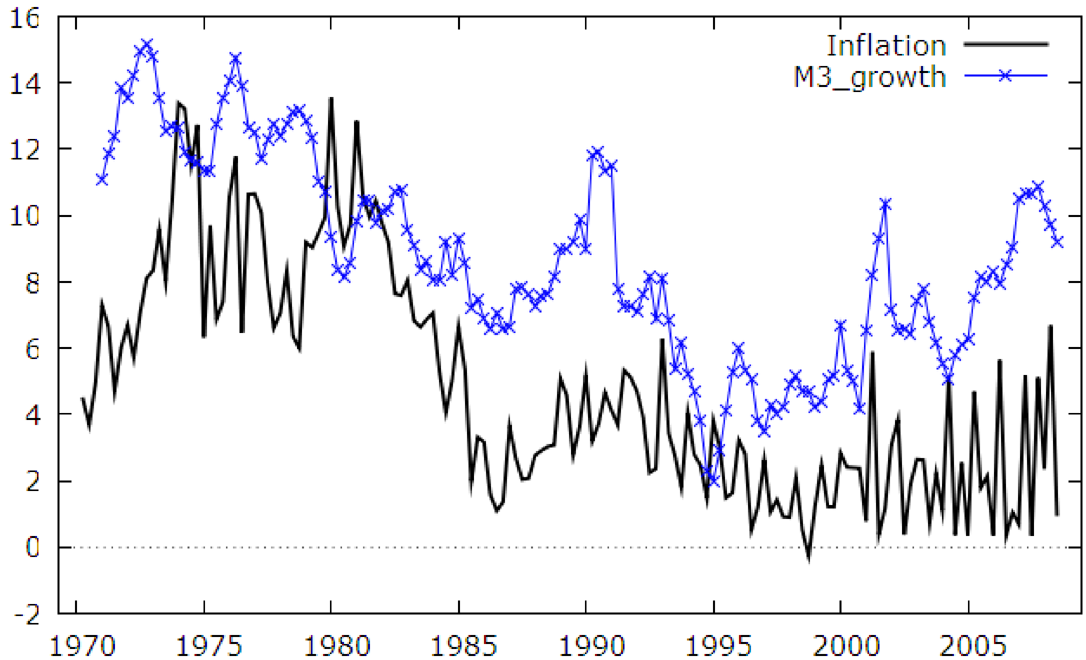

Figure 1: Money growth and inflation (CPI-based) in the euro area

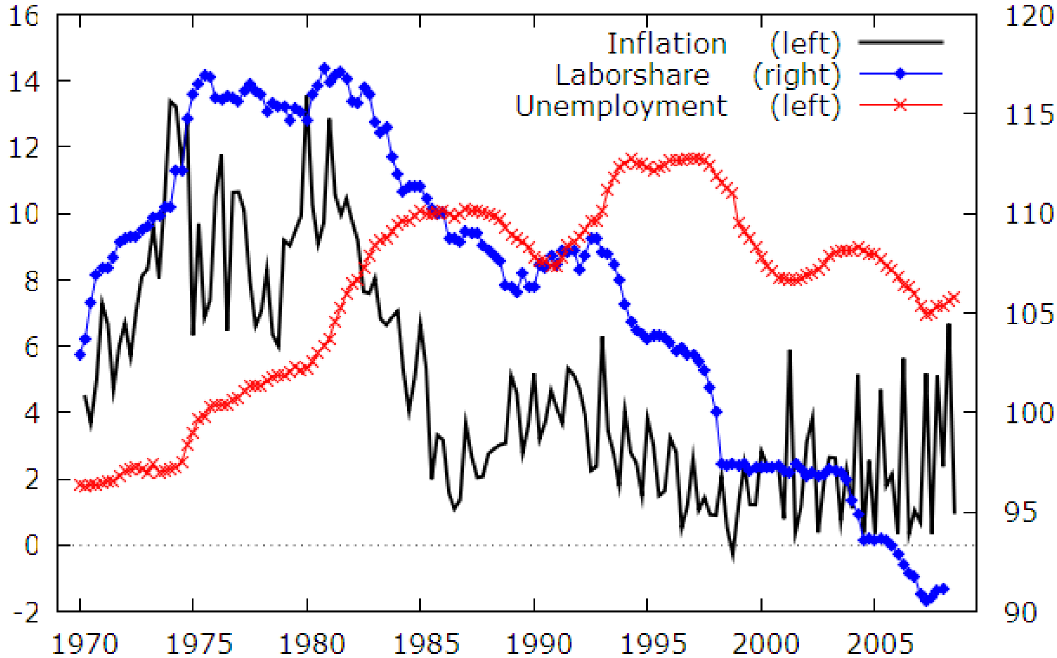

Figure 2: Inflation with labor share and unemployment

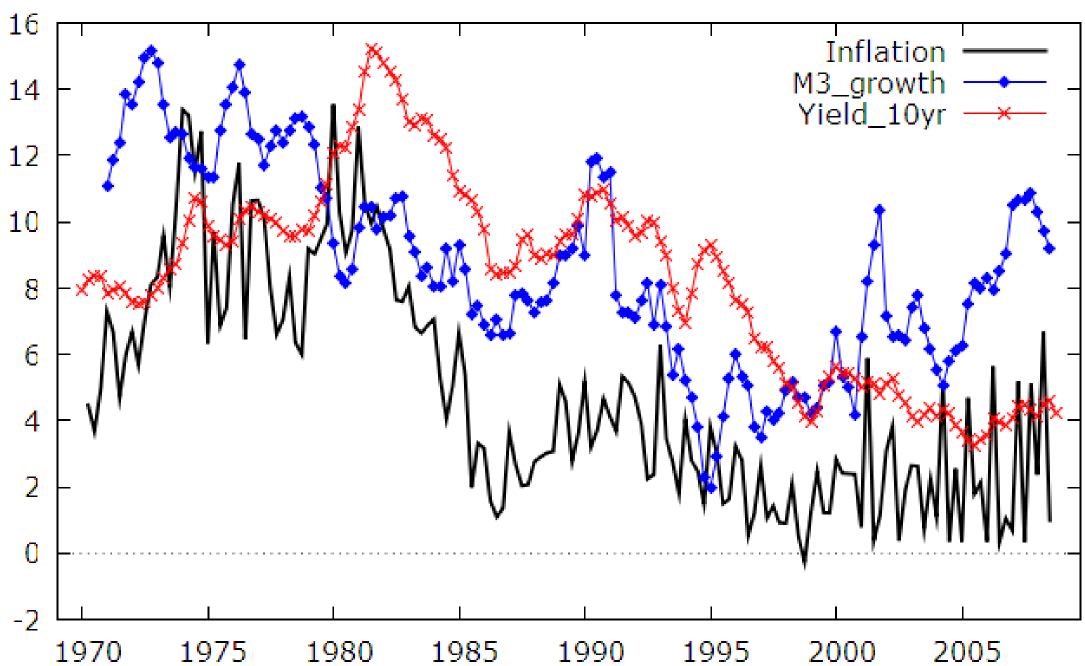

Figure 3: Inflation with money growth and long-term interest rates

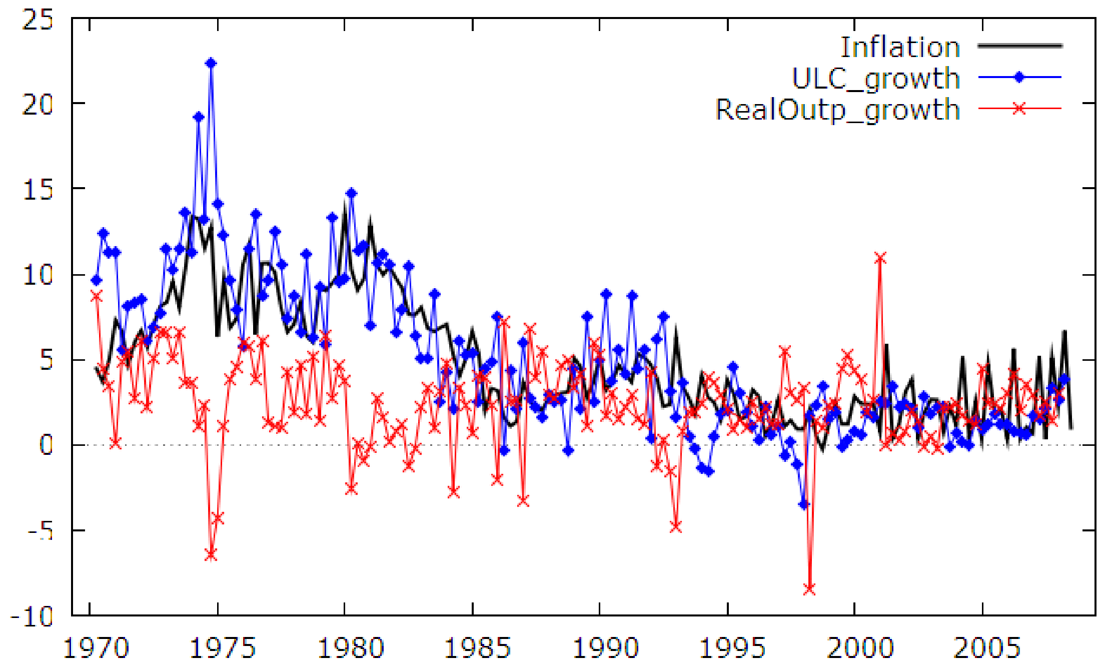

Figure 4: Inflation with unit labor cost growth, and real output growth

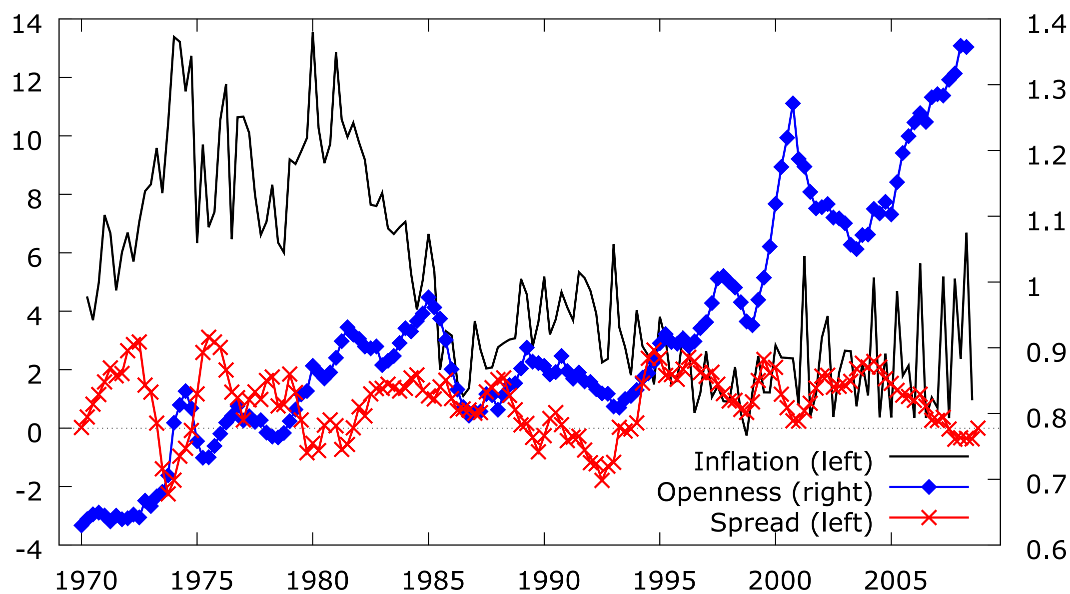

Figure 5: Inflation with the openness measure and the spread

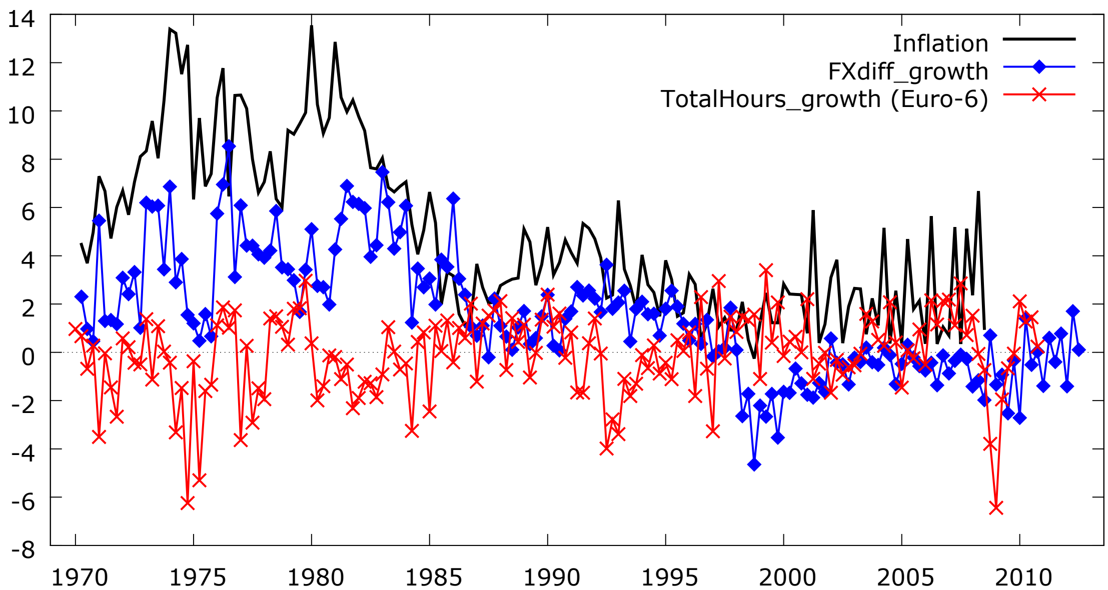

Figure 6: Inflation with the difference of nominal and real exchange rate changes (FXdiff_growth) and growth of total hours

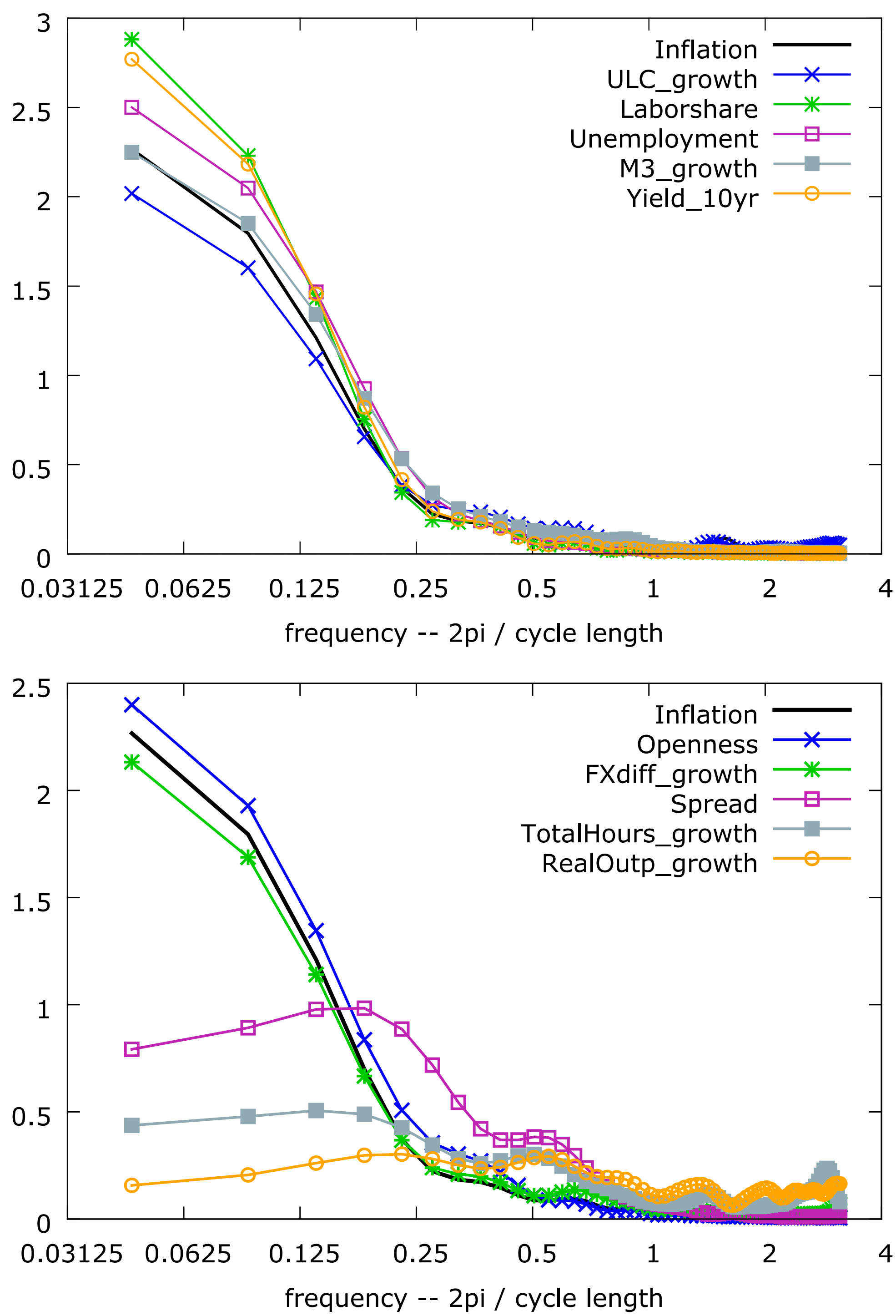

Figure 7: Spectra of the variables. The frequency axis is given in logarithmic scale to emphasize the low-frequency portion. Variables were normalized to have unit variance.

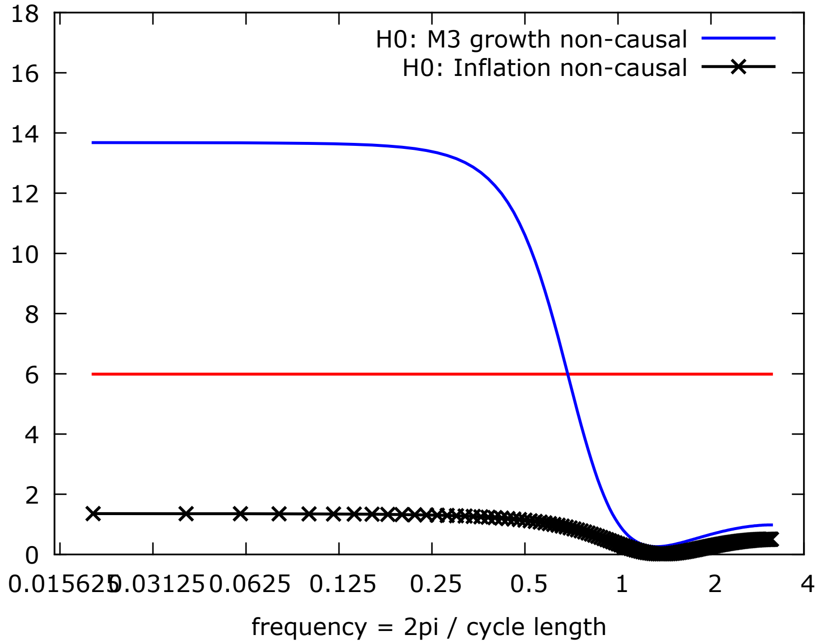

Figure 8: B&C test results, bivariate system of inflation and M3 growth replicating AW&G's results. The critical value is represented by the horizontal line.

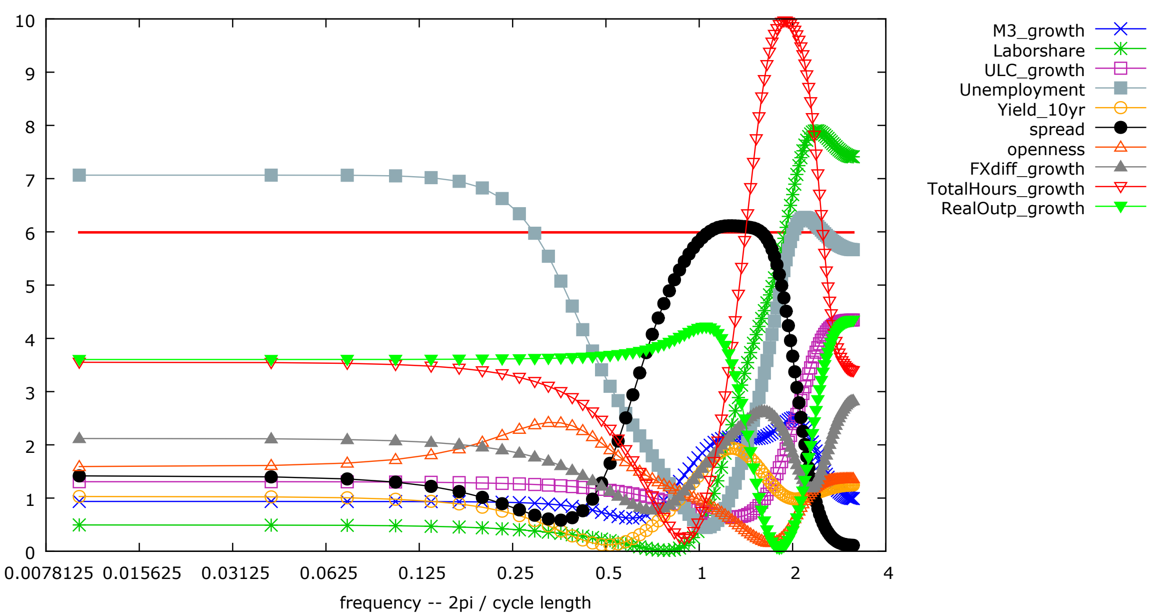

Figure 9: Detailed B&C test results, full variable set, target variable is the inflation rate. The critical value is represented by the horizontal line.

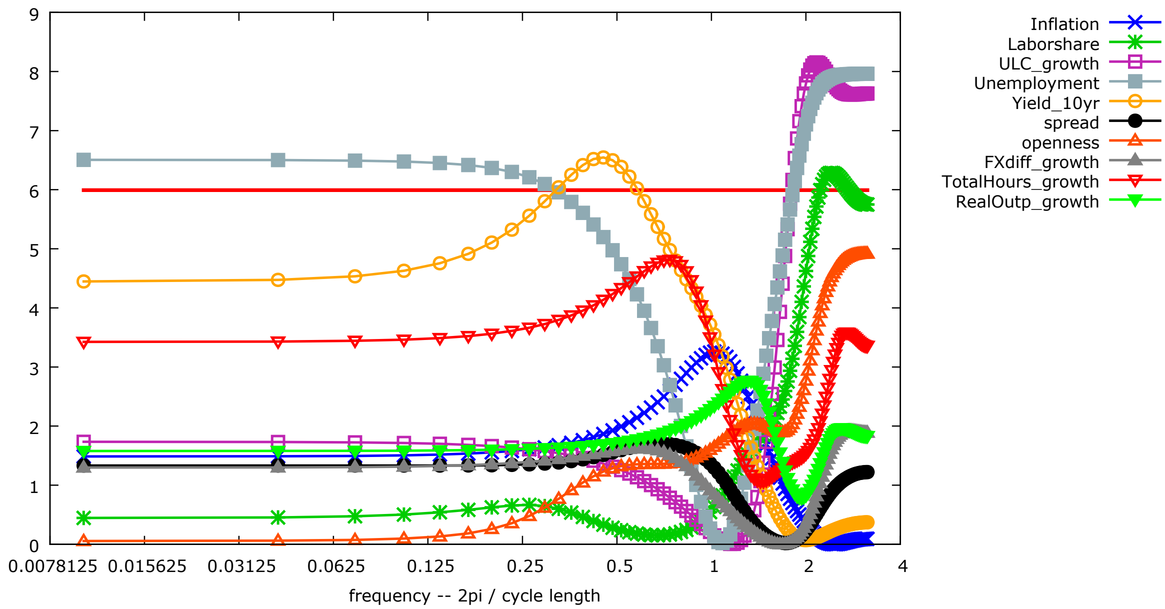

Figure 10: Detailed B&C test results, full variable set, target variable is money growth. The critical value is represented by the horizontal line.

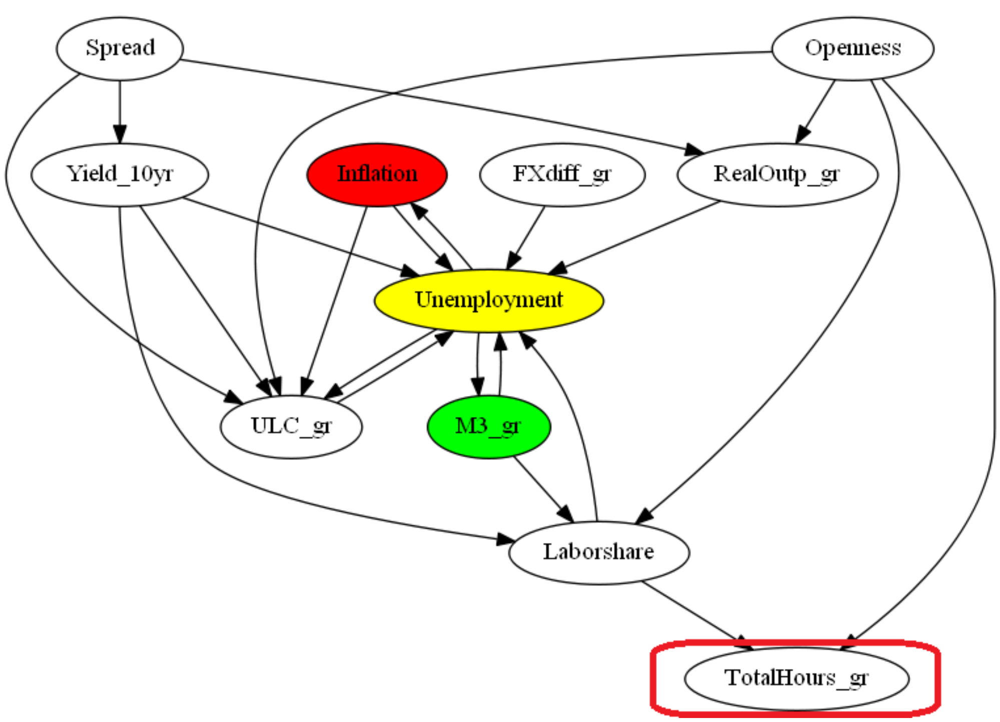

Figure 11: Low-frequency G-causality graph, full variable set. Drawn edges with arrows indicate significant low-frequency Granger causality (G-causality).

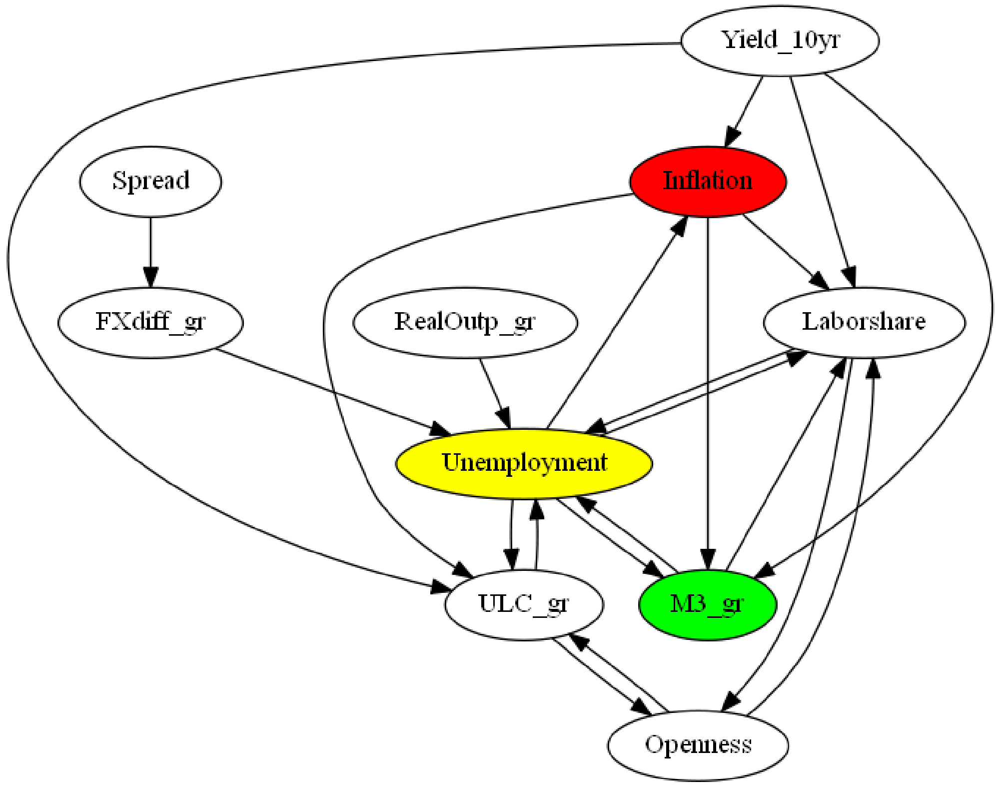

Figure 12: Low-frequency G-causality graph, 2nd iteration

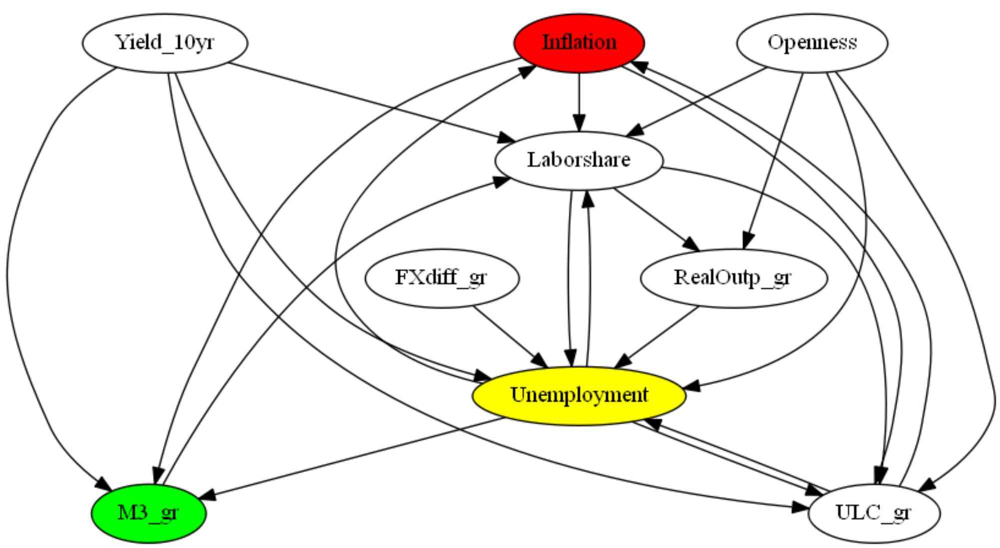

Figure 13: Low-frequency G-causality graph, 3rd iteration

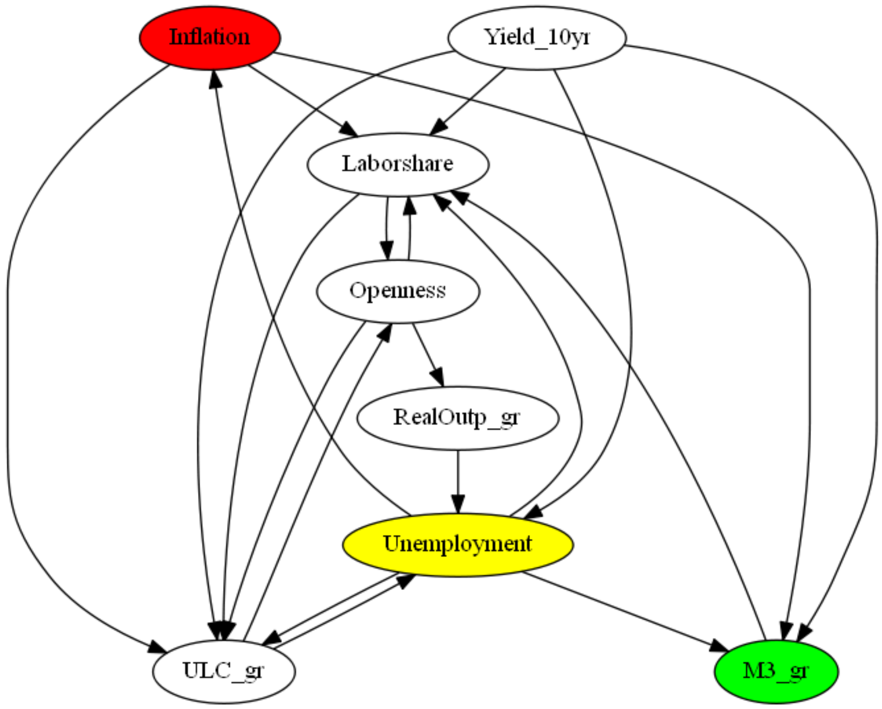

Figure 14: Low-frequency G-causality graph, 4th iteration

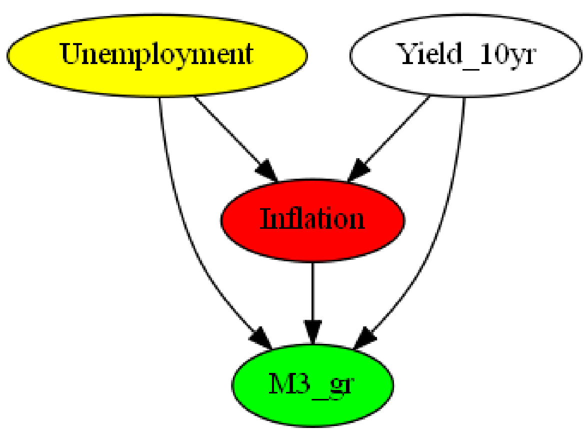

Figure 15: Low-frequency G-causality graph, final

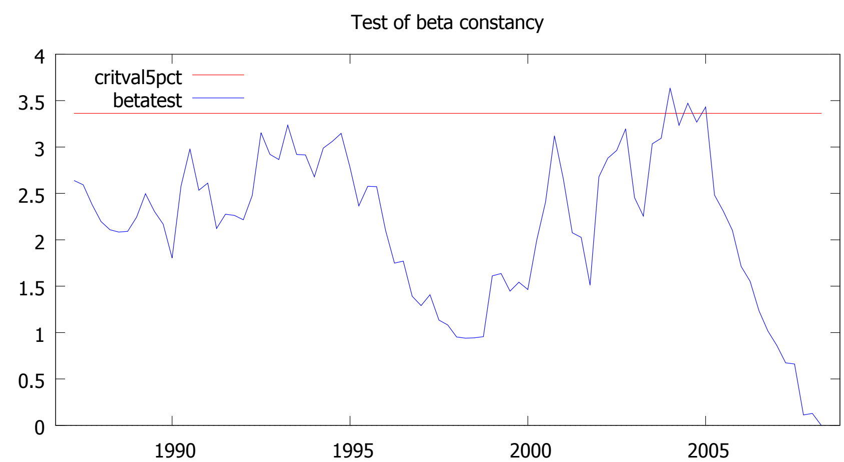

Figure 16: Stability test of the cointegration coefficients (four-dimensional system as in table 2, but based on unrestricted estimates)

Table 1: Bivariate cointegration analysis

rank	eigenvalue	trace stat.	λ-max stat.
0	0.10	18.54 [0.084]	14.57 [0.079]
1	0.03	3.98 [0.427]	3.98 [0.426]

Notes: Johansen cointegration rank test; p-values in brackets; lag order = 5; sample 1974:1 - 2008:3 (T=137), restricted constant.

	Inflation	M3 growth	constant
error-correction term	1	-1	3.98 (0.512)
loadings	-0.259 (0.069)	0	-

Notes: VECM estimates; standard errors in parentheses; LR test of the two restrictions $\chi^2(2) = 0.109$ (p = 0.947).

Table 2: Cointegration analysis and VECM estimation results

rank H0	trace stat.	asympt. p-val	Bartlett p-val	Bartlett trace stat.
0	98.8	0.000	0.000	87.9
1	47.8	0.001	0.005	43.3
2	15.3	0.213	0.309	13.8
3	4.39	0.369	0.517	3.41

Notes: Cointegration rank test, asymptotically and with small-sample Bartlett correction according to Johansen (2002); restricted constant, lag order = 3, sample: 1974:1 - 2008:2 (T = 138).

	Inflation	M3_growth	Unemployment	Yield_10yr	cnst
coint. relation 1	1	-1	0	0	3.71 (0.350)
adjustment coeff.	0	0.37 (0.059)	0	0	-
coint. relation 2	1	0	0.90 (0.038)	-0.43 (0.033)	-8.45 (0.451)
adjustment coeff.	-0.83 (0.135)	-0.64 (0.118)	0	0	-

Notes: Restricted VECM estimates, coefficients 1, 0, -1 are restricted and/or normalized; restriction test: $\chi^2(7) = 10.10$ (p = 0.183); VECM contains restricted constant, levels lag order 3, standard errors in parentheses, sample: 1974:1 - 2008:2 (T = 138).

Table 3: Estimates with breaks in the long-run adjustments

Unrestricted estimates		Inflation equ.	M3_growth equ.	Unempl. equ.	Yield equ.
α_1	adj. to ECT1	0.14 (0.168)	0.40** (0.087)	0.038* (0.015)	0.017 (0.036)
	adj. to ECT2	-0.83** (0.308)	-0.66** (0.159)	-0.061* (0.029)	0.0072 (0.067)
α_2	adj. to ECT1	-0.055 (0.197)	-0.067 (0.102)	-0.031 (0.018)	-0.039 (0.043)
	adj. to ECT2	-0.40 (0.310)	0.025 (0.160)	0.051 (0.029	-0.0099 (0.067)

Restricted estimates		Inflation equ.	M3_growth equ.	Unempl. equ.	Yield equ.
α_1	adj. to ECT1	0	0.37** (0.0574)	0.036* (0.0140)	0
	adj. to ECT2	-0.64** (0.143)	-0.64** (0.115)	-0.057* (0.0239)	0
α_2	adj. to ECT1	0	0	-0.036* (na)	0
	adj. to ECT2	-0.40* (0.158)	0	0.057* (na)	0

Restriction test: $F_{10,496} = 0.625$, $p = 0.793$.

Notes: ECT1 and ECT2 are the (deviations from the) cointegration relationships from table 2; notably $ECT1 = \Delta p - \Delta m + const$. The adjustment matrix for the subsample 1974:3-1990:4 is α_1, for the subsample 1991:1-2008:2 it is $\alpha_1 + \alpha_2$. Standard errors in parentheses (not separately available in the unemployment equation for α_2 because the coefficients are restricted to be equal to the negative of the ones of α_1).