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

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(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 frequencies. Applying frequency-specific tests in a comprehensive system setup for euro-area data we consider various theoretical predictors 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

[a]This paper has benefited from comments by Dieter Nautz, Jörg Breitung, Till van Treeck, Christian Proaño-Acosta, Jürgen Wolters, Helmut Lütkepohl, and by seminar participants at Goethe University Frankfurt, IWH Halle, Hamburg University, and Free University Berlin. Nonetheless errors are mine.

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1 Introduction

For all economic agents it is important to understand the causes behind inflation movements at low frequencies, where low-frequency developments may be thought of as the slower but long-lasting, longer-run changes of a variable. For monetary policy it is important to predict the long-run inflation developments in order to assess inflationary pressures and to be able to adjust its policy stance accordingly. Private agents of course also undertake long-term financial planning and must therefore forecast the more persistent movements of inflation.

An established and popular view of the forces behind inflation movements is based on the traditional quantity theory of money. According to that view, inflation is predominantly a monetary phenomenon, and therefore movements of money growth are supposed to cause inflation changes. A problem with this approach is that already a casual look at the data of many (developed) countries typically suggests that money growth and inflation indeed share long-run developments, but can be quite disconnected in the short term, see figure 1 for the euro-area example. This empirical assessment is also reflected in the practice of modern macroeconomics to build models without monetary aggregates and many central banks have abandoned looking closely at the developments of monetary aggregates.

However, other economists such as the intellectual founders of the European Central Bank (ECB) saved the monetary view by inventing 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 adopt this view by adding money growth to reduced-form models of inflation (Gerlach, 2004; ECB, 2004; Beck and Wieland, 2007). For a general discussion on the role of money in monetary policy and for inflation in the long run see the contributions in a special issue of the Journal of Money, Credit and Banking: Woodford (2008); McCallum (2008); Nelson (2008).

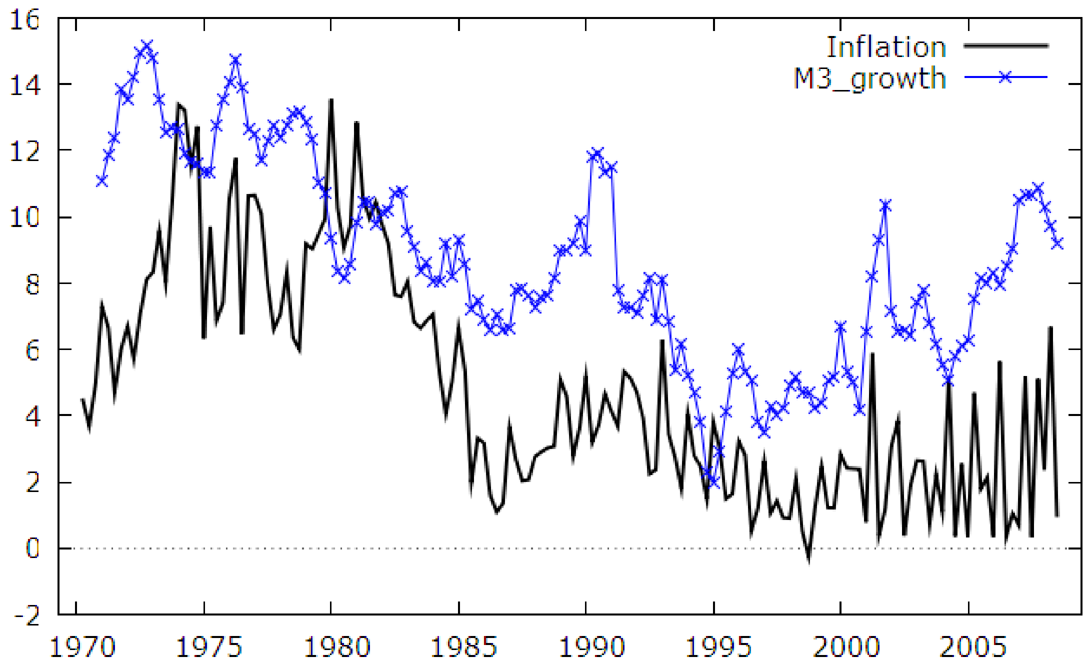

Figure 1: Money growth and inflation (CPI-based) in the euro area

Additionally, in a series of papers Assenmacher-Wesche and Gerlach (2007, 2008a,b, AW&G) have recently 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] Another very recent result about the predictive content of money growth for (different regimes of) inflation is given in Amisano and Fagan (2013).

In this paper we take a closer look at the result of low-frequency G-causality running from money growth to inflation, for the case of the euro area. We use a broader theoretical foundation for the low-frequency predictors of inflation, including the quantity theory of money but considering also other possible influences coming from goods, labor, and financial markets, and from abroad. Mirroring the approach of considering more than one theory, our

[1] 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.

empirical strategy is to perform a general-to-specific search routine where empirically non-causal variables are successively excluded from the analysis. We employ essentially the same econometric methods as AW&G, especially low-frequency causality tests in a system conditional on other persistent variables. As an empirical concept, the notion of Granger causality (Granger, 1969, 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.

In addition, we provide a detailed specification of vector error-correction models (VECM) to quantify the long-run relations. Within this parametric framework we also check whether the error-correction mechanisms (which correspond to the G-causality at the zero frequency) have changed within the sample.

Based on the richer information set in our analysis we arrive at conclusions that are quite different from AW&G's. Money growth turns out as non-causal, while unemployment and long-term interest rates are 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.[2]

In the following section 2 we discuss the underlying theories. Then in section 3 we briefly introduce the frequency-domain causality measures and tests, and we report the empirical details of the search routine and its results. After that we present the VECM analysis in section 4. 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]See Lütkepohl (1982) for the theory of omitting variables in Granger-causality tests.

2 Economic theory

In this section we consider several theories that are potentially relevant for low-frequency movements of inflation, and which will determine our information set for the empirical analysis. All relationships are presented in a bare-bones form suppressing constants, error terms and richer dynamics. The empirical methods in this paper account for that.

We start with the quantity theory of money in log-differenced form as in AW&G, where inflation Δp is related to money growth Δm, real output growth Δy, and changes of velocity Δv:

$$\Delta p = \Delta m - \Delta y + \Delta v \tag{1}$$

Of course equation (1) as such is an identity, not a theoretical hypothesis. Apart from possible additional assumptions about the properties of velocity such as relative stability or whether it is related to interest rate changes as in AW&G, the key theoretical issue is precisely given by the hypothesis that money growth tends to determine inflation and not vice versa.[3] Since (1) is an identity, we do not need to consider velocity changes once the other three variables are accounted for.

Next, it is natural to consider the New Keynesian Phillips curve (NKPC, see e.g. Galí and Gertler, 1999) as a modern theory of inflation, where inflation is driven by discounted expected marginal costs, with λ and β as parameters:

$$\Delta p_t = \lambda \sum_{k=0}^{\infty} \beta^k \mathsf{E}_\mathsf{t}(\text{marg. costs}_{t+k}) \tag{2}$$

The standard approach is to use (log) real unit labor costs $ulc - p$ (essentially the labor share)

[3] 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.

as a proxy for unobservable marginal costs. However, it should be noted that the theory is originally formulated for business-cycle frequencies and may fail at lower frequencies. Also, since it is *expected* future marginal costs that drive current inflation, if the theory and expectations are (roughly) correct then the predictive Granger causality should run the other way around, from realized inflation to realized marginal costs, given that we do not observe expected marginal costs.

Wage-curve models yield additional insights about which factors may affect inflation (see e.g. Blanchard and Katz, 1997, 1999). From the point of view of wage setters, expected real hourly wage growth can be written as $\Delta w_t - \Delta p_t^e$ (where Δw is hourly nominal wage growth, Δp^e is expected inflation, and we loosely denote with superscript e a more general concept of expectations than rational model-consistent expectations) and depends on lagged real unit labor costs, unemployment[4] u_t, and the growth of real labor productivity per labor input (hours) Δq:

$$\Delta w_t - \Delta p_t^e = -(1-\alpha)(ulc-p)_{t-1} - \beta_u u_t + (1-\alpha)\Delta q_t, \qquad \alpha \in [0;1], \beta_u > 0,$$

which can be extended with more complicated dynamics. Using the identities $\Delta ulc_t = \Delta w_t - \Delta q_t$ and $\Delta q_t = \Delta y_t - \Delta h_t$, where h is total labor inputs (log hours), we can rearrange the equation as follows:

$$\Delta ulc_t = \Delta p_t^e - (1-\alpha)(ulc-p)_{t-1} - \beta_u u_t - \alpha \Delta y_t + \alpha \Delta h_t \qquad (3)$$

In analogy to the discussion of the expectational NKPC, but reversing the argument, here the empirical G-causality would run in the "right" direction towards observed inflation.

Partially embedded in (3) is another theoretical relationship to justify the inclusion of the

[4]More generally it is the overall labor market tightness which matters. The unemployment rate serves as a reasonable proxy here.

growth rate of (nominal) unit labor costs, namely a simple differenced mark-up pricing rule:

$$\Delta p = \Delta ulc^{(e)}, \tag{4}$$

where we write "(e)" with parentheses to denote that it may be either realized or expected developments of unit labor costs which determine inflation, depending on the timing of information flows. Again, if the true relationship is expectational, $\Delta p = \Delta ulc^{e}$, then the empirical Granger causation would actually run from inflation to unit labor costs. Without differentiating between frequency bands this direction appears to be the empirical finding at least for US data (Mehra, 1991; Strauss and Wohar, 2004). Note that a 1:1 relation between Δp and Δulc in the long-run will hold in all standard models (see e.g. Sbordone, 2002), and that taking (3) and (4) together would imply feedback G-causation between unit labor costs and inflation, i.e. something like a wage-price spiral.

Next we consider the Fisher relation, where the long-term (nominal) interest rate i consists of an equilibrium real rate rr and fluctuations determined by inflation expectations:

$$i = rr + \Delta p^{e} \tag{5}$$

Long-term interest rates i thus should be predictive for realized inflation.

In addition to long-term interest rates there could also be a role for short-term rates by considering a standard monetary-policy reaction function:

$$i_{short} = \beta \Delta p^{(e)} + \gamma \Delta y \tag{6}$$

Here we have formulated the rule in terms of output growth, another standard approach would be to include some measure of the output gap. Since our information set already includes the long rates i, we use the spread $i - i_{short}$ as an equivalent substitute for the short rates in the empirical analysis.

Finally, we account for the open-economy dimension by considering the following two relations: First we employ a differenced form of the real-exchange rate identity:

$$\Delta p = \Delta p^* + \Delta E_{real} - \Delta E, \tag{7}$$

where Δp^* is world inflation and ΔE and ΔE_{real} are the changes of nominal and real exchange rates, respectively. The direction of causality here is mainly an empirical issue. Since relationship (7) is an identity, the difference $\Delta E_{real} - \Delta E$ already captures the informational content and we drop foreign inflation to avoid perfect collinearity. The second potential channel affecting inflation in an open economy is given by the hypothesis of Romer (1993) where an economy's degree of openness dampens the inflationary bias of monetary policy:

$$\Delta p = f(\underset{\ominus}{open}) \tag{8}$$

For a discussion and evidence of this connection see IMF (2006); Pain et al. (2008).

This concludes our fairly comprehensive tour of potentially relevant economic theories for the low-frequency movements of inflation. Altogether, the set of variables that are included in the analysis is therefore given by:

$$\Delta p, \Delta m, \Delta ulc, ulc - p, urate, i, spread, \Delta y, open, \Delta E_{real} - \Delta E, \Delta h$$

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 measured as the ratio of imports and exports over output.

3 Frequency-domain analysis[5]

For long-term interest rates we use 10-year government bond yields, for money we use 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 turmoil of the financial and economic crisis, because we believe that this episode represents a different regime. 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.

3.1 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

$$f_{xx}(\omega) = \frac{1}{2\pi}\sum_{\tau=-\infty}^{\infty} R_{xx}(\tau)\exp(-j\omega\tau), \qquad -\pi \leq \omega \leq \pi \tag{9}$$

[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.

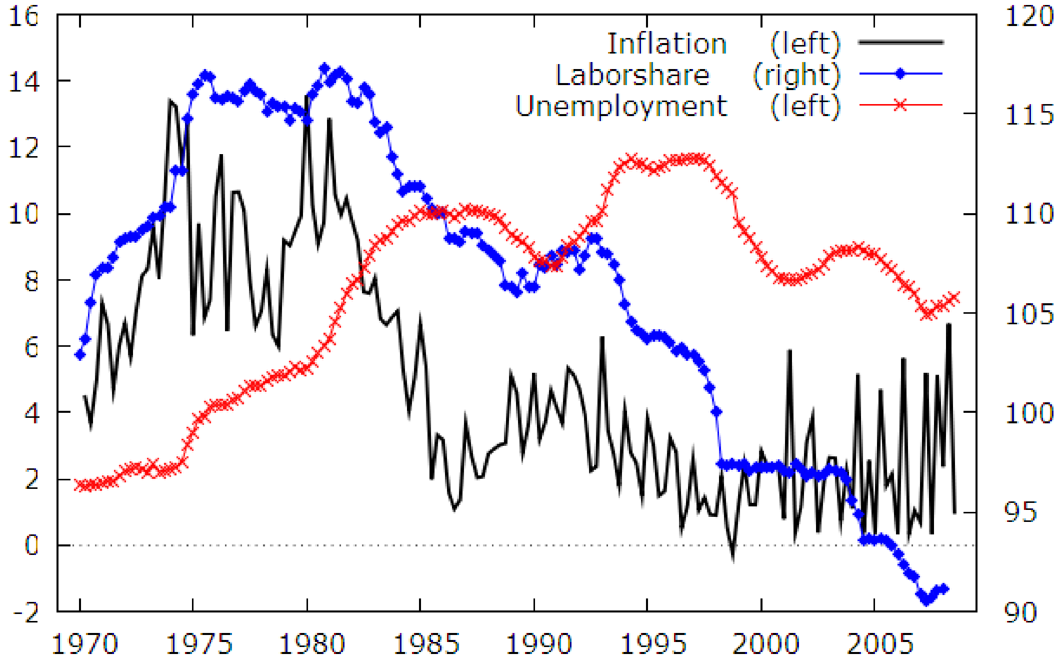

Figure 2: Inflation with labor share and unemployment

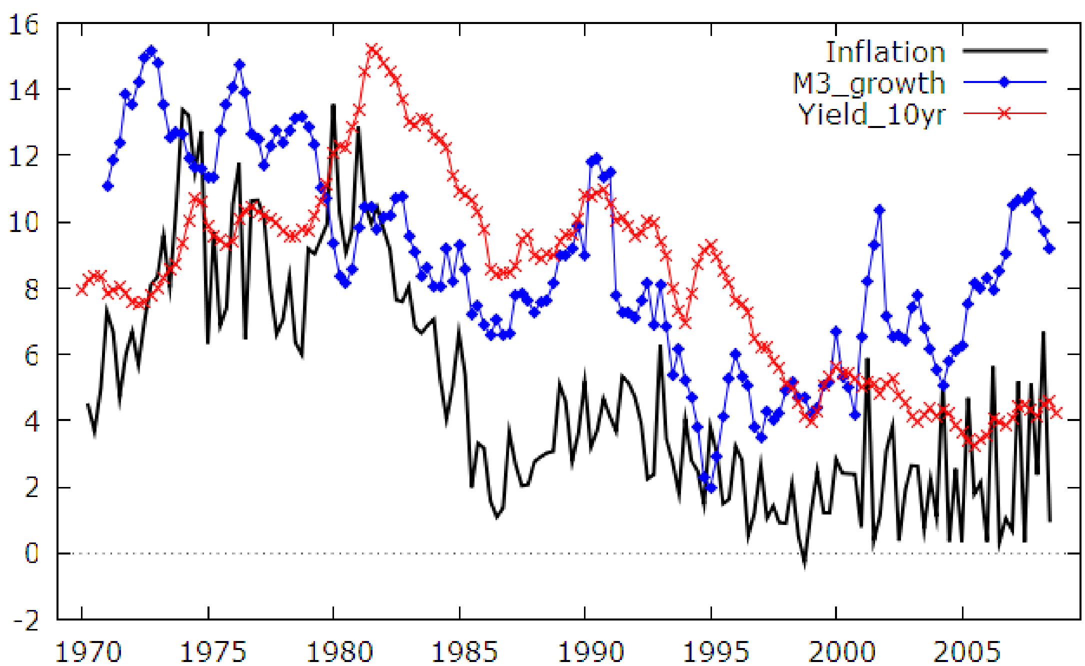

Figure 3: Inflation with money growth and long-term interest rates

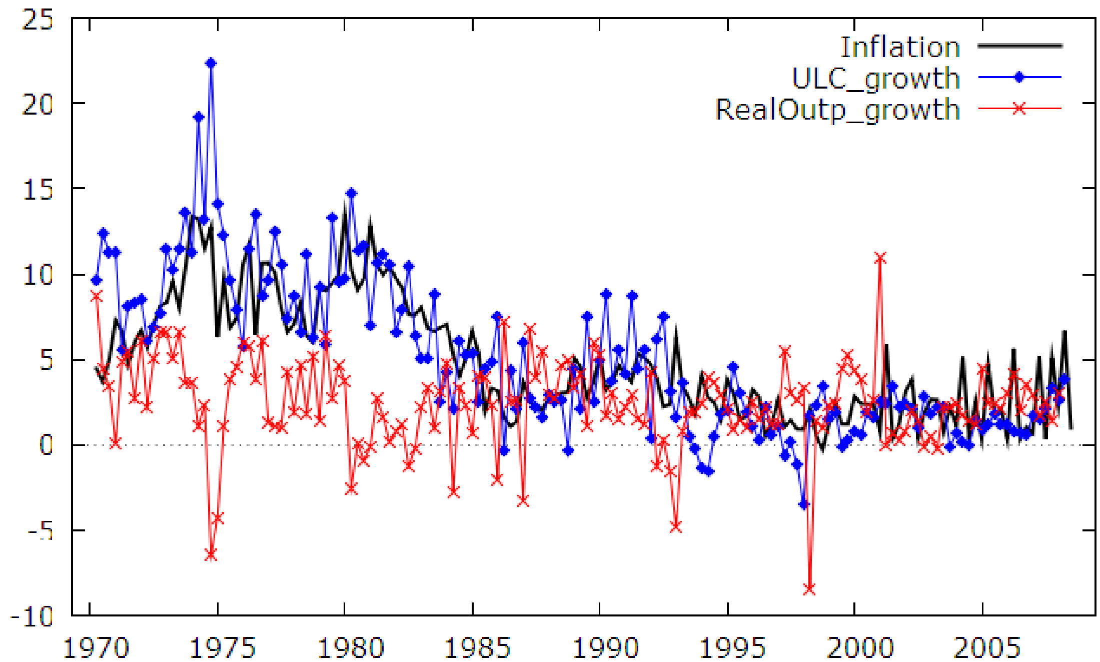

Figure 4: Inflation with unit labor cost growth, and real output growth

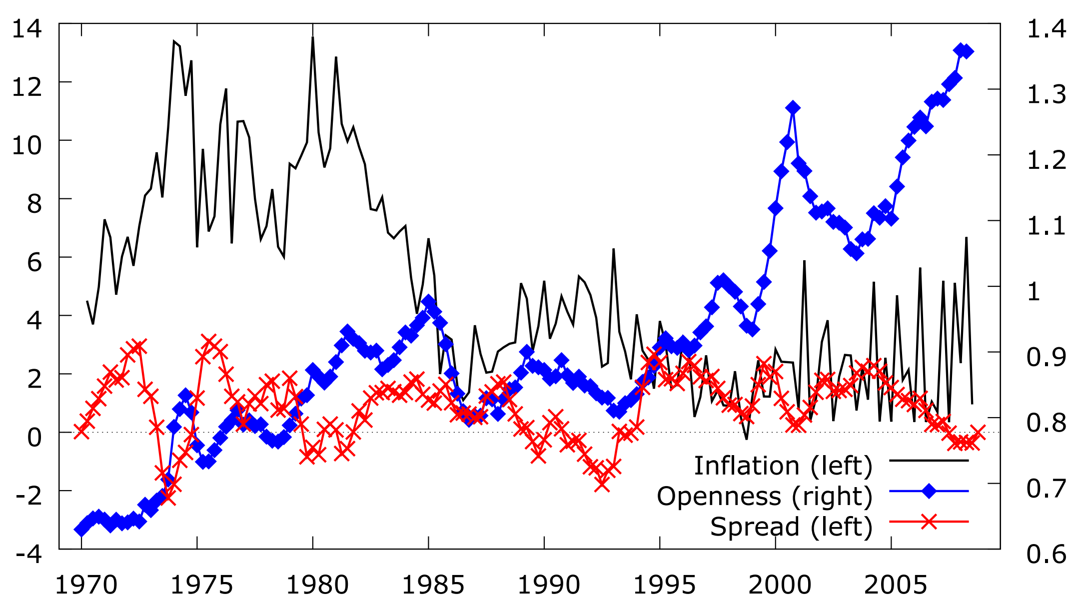

Figure 5: Inflation with the openness measure and the spread

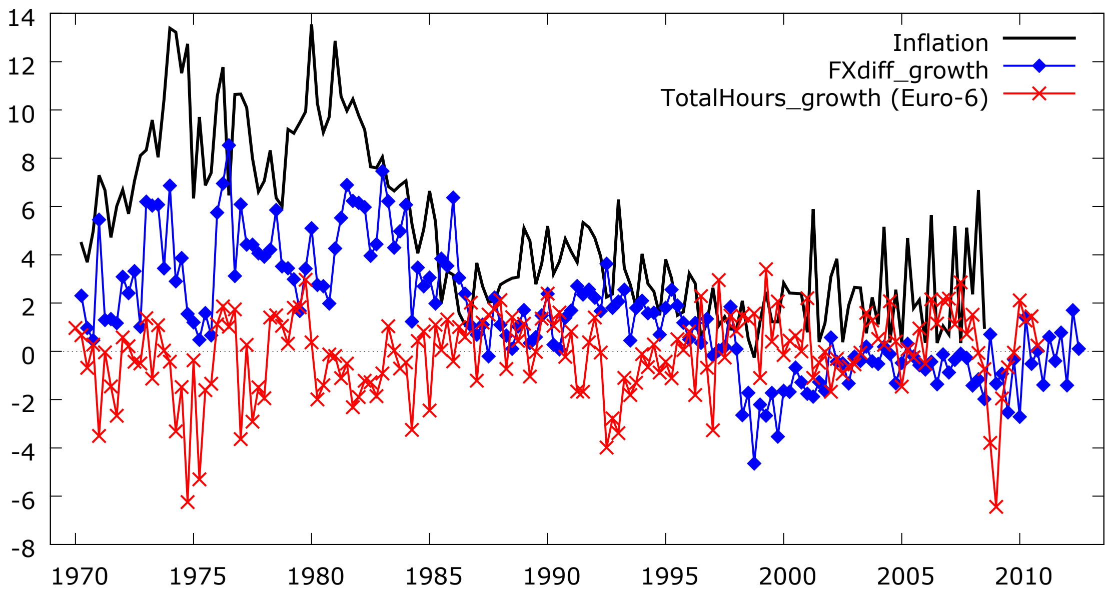

Figure 6: Inflation with the difference of nominal and real exchange rate changes (FXdiff_growth) and growth of total hours

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

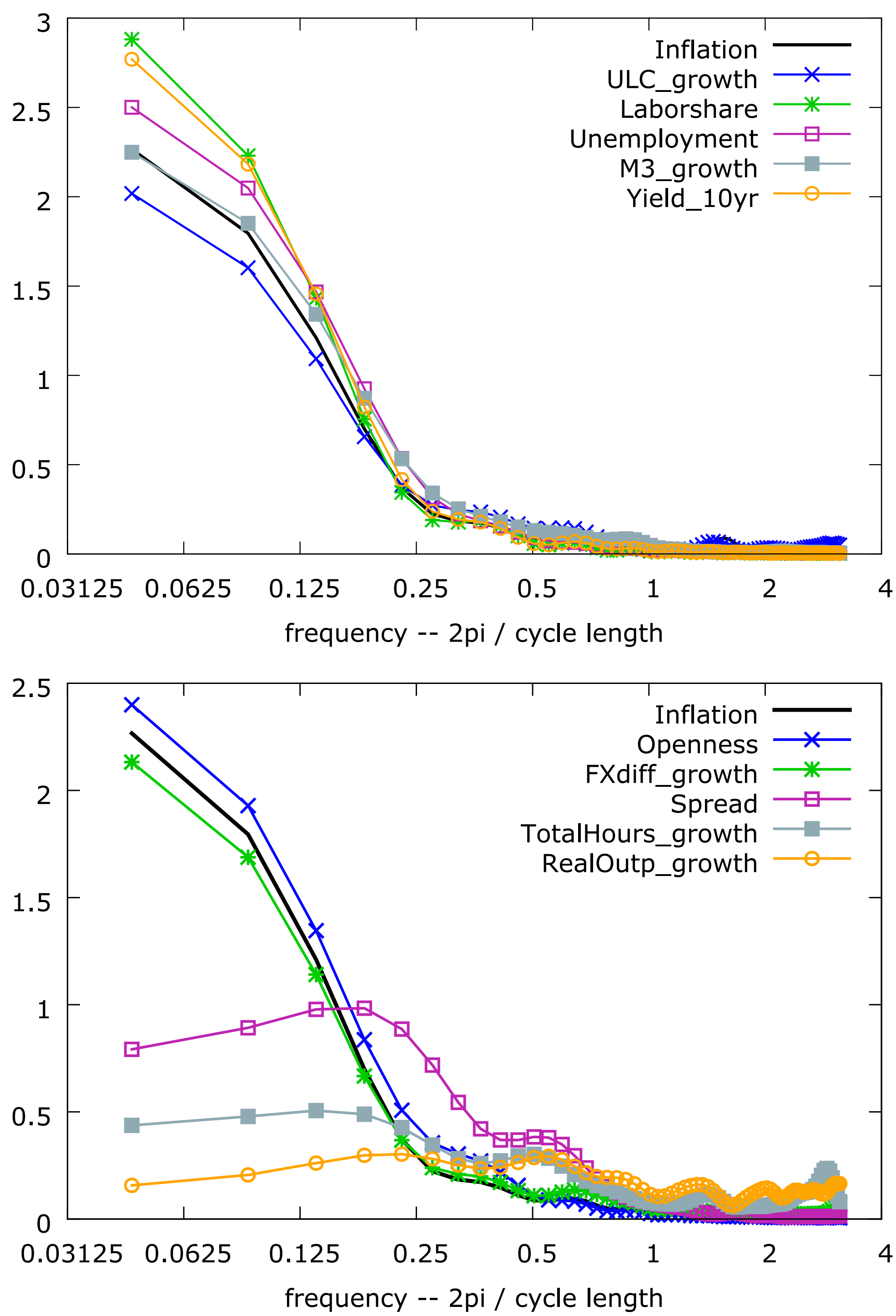

Figure 7: Spectra of the variables. The frequency axis is given in logarithmic scale to emphasize the low-frequency portion. Variables were normalized to have unit variance.

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:

$$\Psi(L) = \begin{pmatrix} \Psi_{11}(L) & \Psi_{12}(L) \\ \Psi_{21}(L) & \Psi_{22}(L) \end{pmatrix} \tag{10}$$

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

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

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-

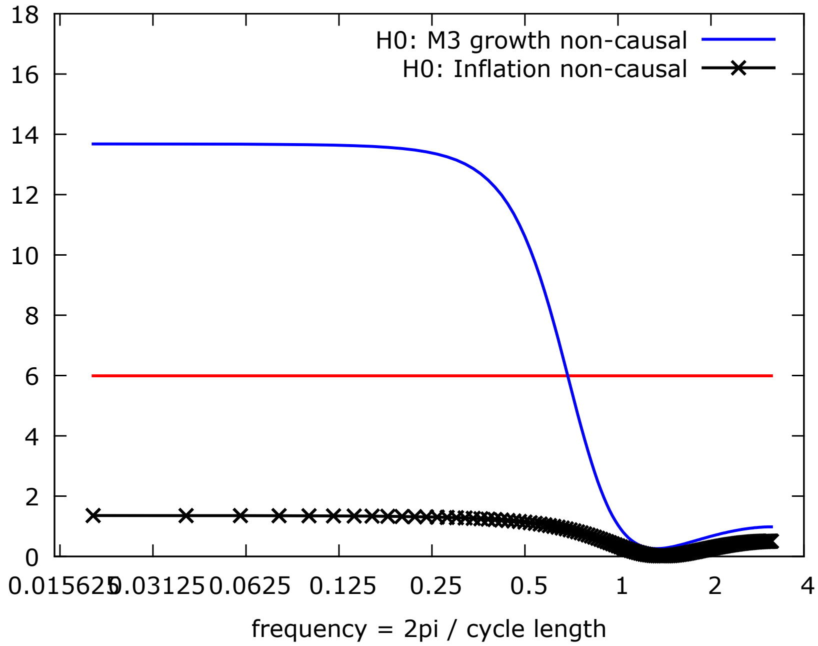

Figure 8: B&C test results, bivariate system of inflation and M3 growth replicating AW&G's results. The critical value is represented by the horizontal line.

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. 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.

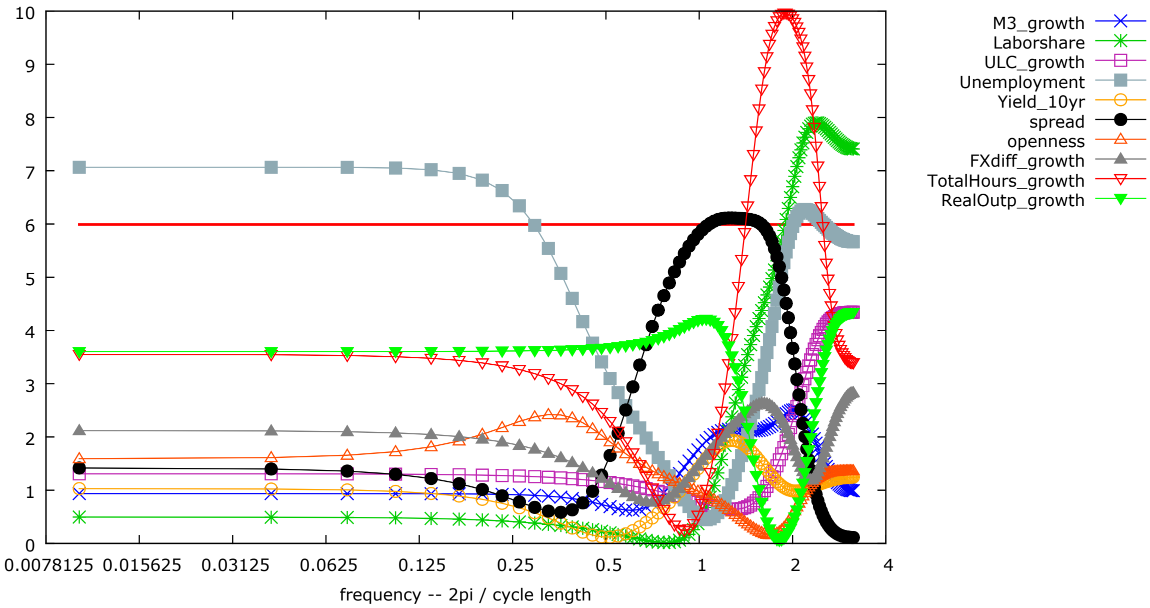

Figure 9: Detailed B&C test results, full variable set, target variable is the inflation rate. The critical value is represented by the horizontal line.

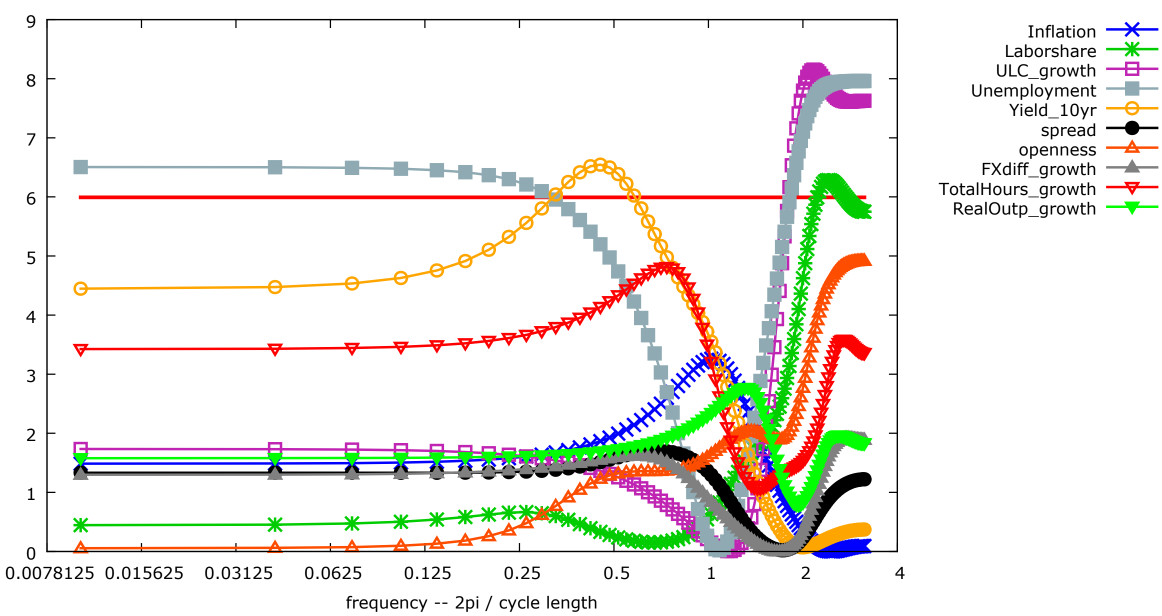

Figure 10: Detailed B&C test results, full variable set, target variable is money growth. The critical value is represented by the horizontal line.

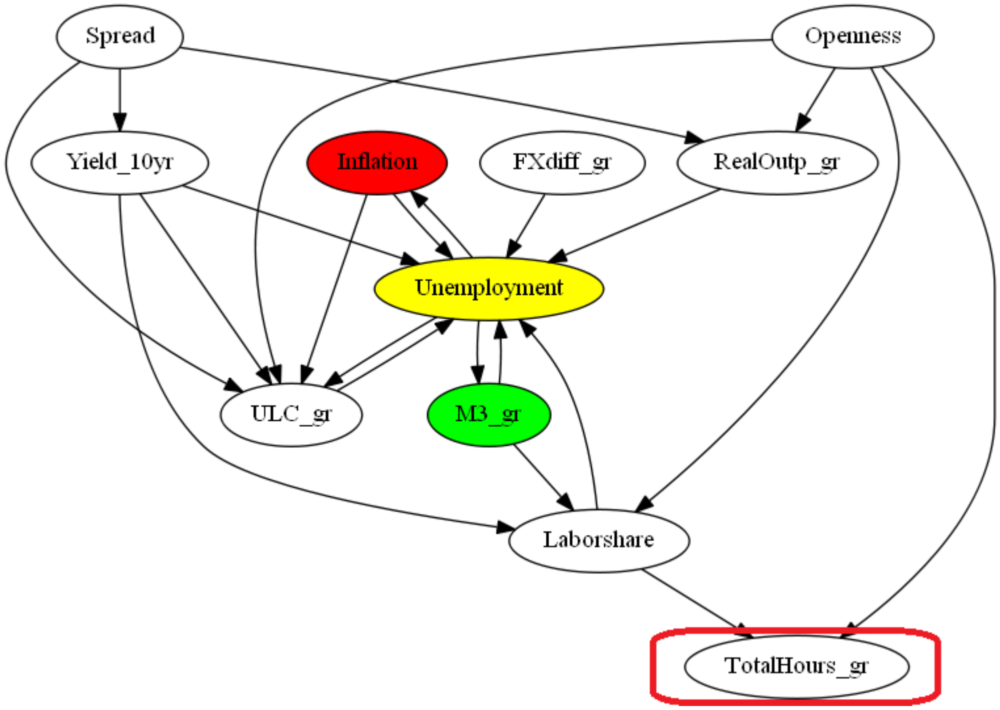

Figure 11: Low-frequency G-causality graph, full variable set. Drawn edges with arrows indicate significant low-frequency Granger causality (G-causality).

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.

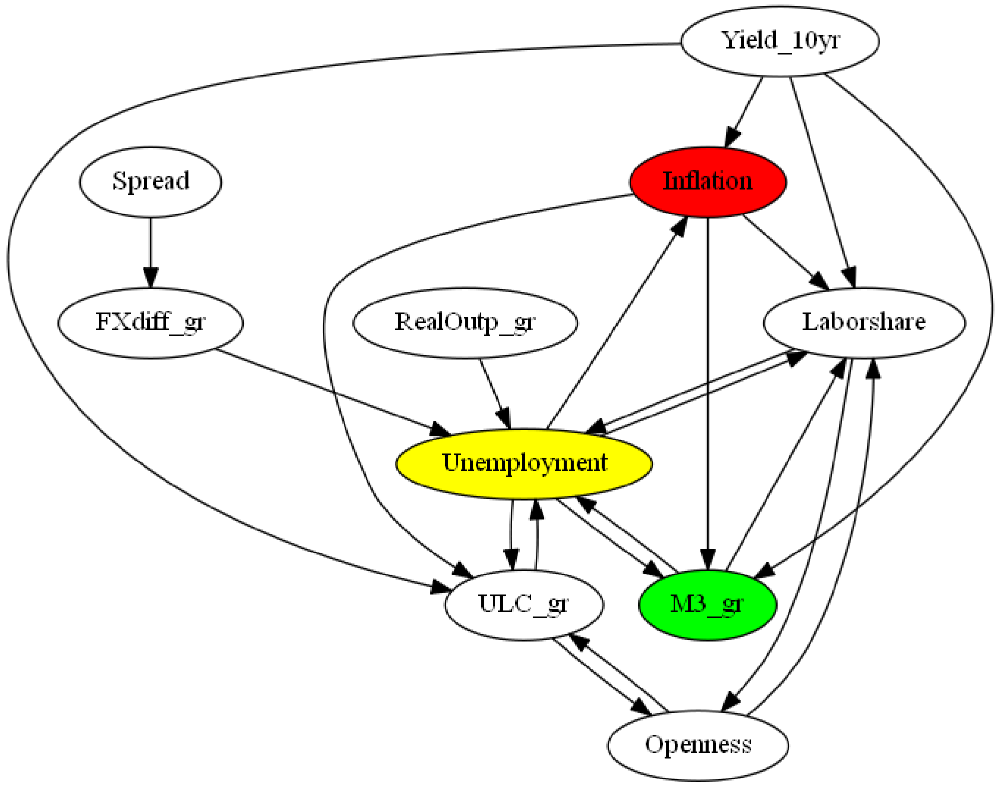

Figure 12: Low-frequency G-causality graph, 2nd iteration

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

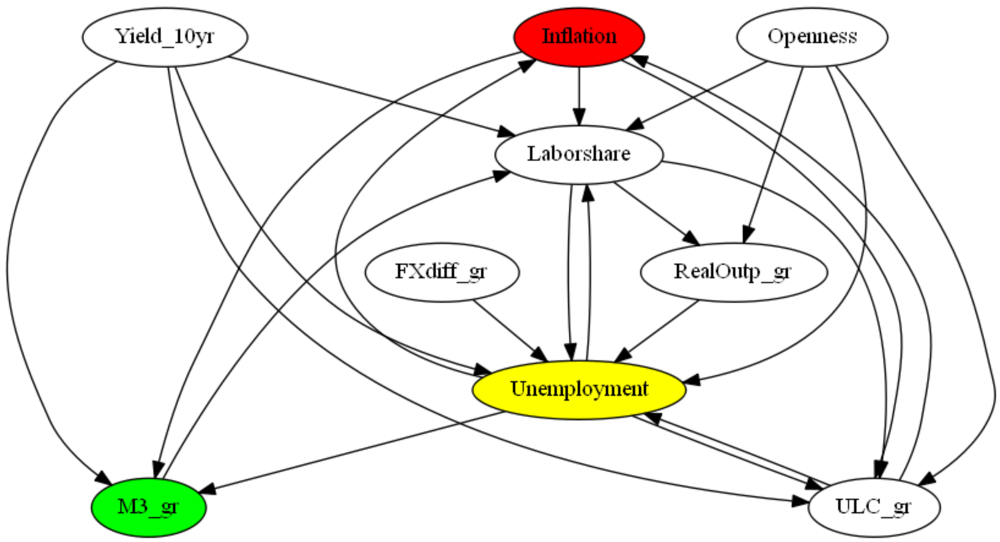

Figure 13: Low-frequency G-causality graph, 3rd iteration

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

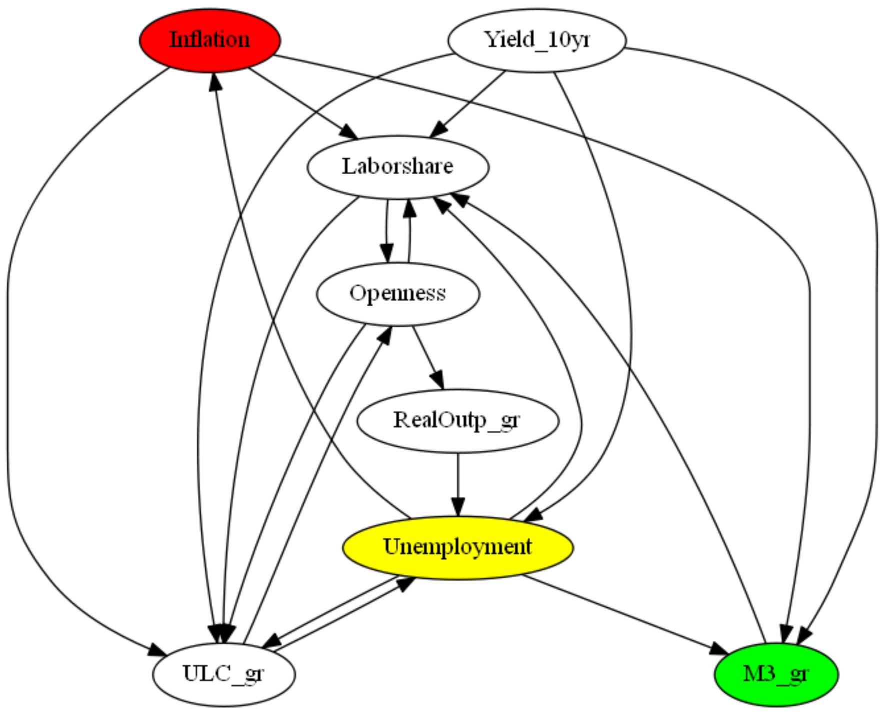

Figure 14: Low-frequency G-causality graph, 4th iteration

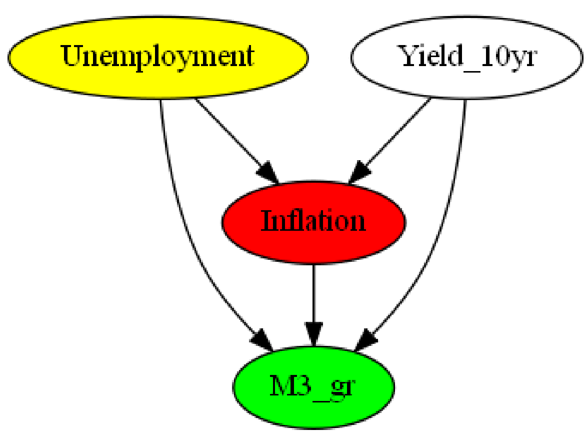

Figure 15: Low-frequency G-causality graph, final

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):

$$\Delta x_t = \alpha \beta^{*'} (x'_{t-1}, 1)' + \sum_{k=1}^{K-1} \Gamma_k \Delta x_{t-k} + \varepsilon_t \tag{12}$$

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.

[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.

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).

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, see the upper panel of table 1. Furthermore, the lower panel reports that it is statistically acceptable to restrict the corresponding cointegration vector to a 1:1 relationship. The super-consistent coefficients of this irreducible cointegration vector 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 between inflation and money growth:

$$\Delta m = 3.71 + \Delta p$$

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

[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.

(apart from the normalization of the inflation coefficient):

$$\Delta p = -8.45 + 0.43\, i - 0.90\, urate$$

These estimated coefficients have plausible signs: higher unemployment tends to lead to lower inflation, and a higher nominal yield tends to be reflected in higher inflation, ceteris paribus. The latter result is in line with the Fisher effect motivation, but of course a strict bivariate Fisher interpretation is difficult to reconcile with this trivariate longer-run relation, apart from the fact that the interest rate coefficient is quite far from unity.

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,

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).

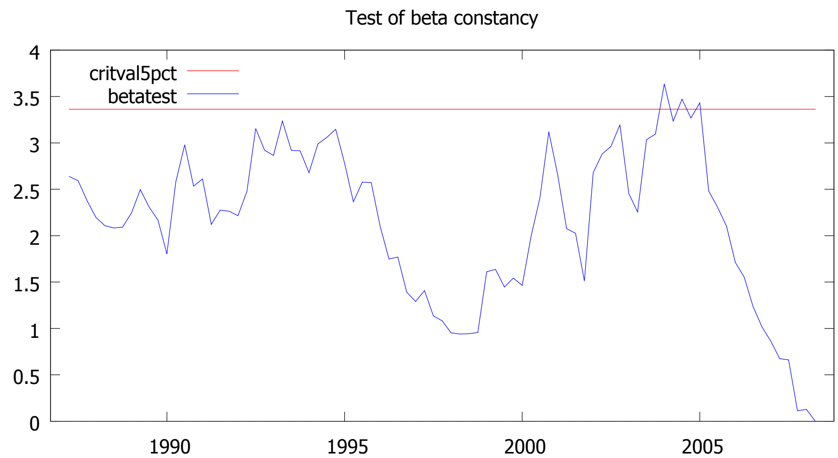

Figure 16: Stability test of the cointegration coefficients (four-dimensional system as in table 2, but based on unrestricted estimates)

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 may have been structural breaks in the dynamic adjustment behavior of the system, for example perhaps due to German unification in 1990/91. Such a break could have had 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.

We model and test for a structural break as follows: The timing of the potential break is fixed exogenously at the middle of the sample, i.e. in the year 1991, which in this sample also roughly coincides with German unification.[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 sepa-

[9]Of course, this date also coincides roughly with the end of the disinflation process in the euro area. We do not choose the potential break period based on the inflation data, however, because that would give rise to pretesting issues.

rate long-run relationships reliably. We thus take the two-dimensional error-correction terms $ect_t = \hat{\beta}^{*\prime}(x_t', 1)'$ 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:

$$\Delta x_t = \alpha_1\, ect_{t-1} + \alpha_2 s_{1991q1,t} ect_{t-1} + \sum_{k=1}^{K-1} \Gamma_{break,k} \Delta x_{t-k} + \varepsilon_{break,t} \tag{13}$$

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 13 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 13. 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 cointegration 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.

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).

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 we find evidence for unemployment and long-term interest rates as predictors –and thus Granger-causal variables– of low-frequency movements of inflation in the euro area. None of the other investigated variables turned out to be relevant long-run drivers once unemployment and long-term interest rates were taken into account, including money growth.[10] We therefore conclude that some recent studies in the literature that attribute an empirically important role to money growth for inflation at low frequencies may suffer from omitted-variable bias. Our findings therefore do not support a prominent role for monetary stance analysis even for the longer run. Instead, for longer-term inflation assessments the ECB as well as the general public would have to focus on the unemployment rate and long-term interest rates.[11]

It is intuitively plausible that unemployment as the main indicator of labor market tight-

[10] The list of variables included growth of unit labor costs, hours worked, real output, and of the gap between nominal and real exchange rates; and the spread, a trade openness indicator, and the labor share. In a reduced bivariate dataset we could also replicate the result by Assenmacher-Wesche and Gerlach (2007, 2008a) that the Granger causality at low frequencies appears to run from money to inflation.

[11] With respect to interest rates one might argue that they are themselves determined by the central bank. However, the central bank only controls the short-term policy rates and the link from short-term to long-term rates (i.e., the yield curve) is not constant. In our analysis the term spread also turned out as irrelevant for low-frequency inflation predictions, given the other variables. In any case, only the policy makers themselves know for sure to what extent changes in their policy rate are reactions to changed long-run inflation expectations. They are free to discard the information contained in long-term rates. But for private agents the signals emitted by movements of long-term interest rates are clearly valuable to assess the long-run inflation outlook.

ness signals future inflation changes (for example, rising unemployment tends to dampen inflation in the long run). Equally plausible is the positive low-frequency effect of long-term interest rates on inflation, because they likely signal movements of long-run inflation expectations which later materialize in observed inflation rates. In addition, we can 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.

Checking for structural shifts in these patterns, we only 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 monetary aggregates would 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 return to standard monetary policy operations our findings would again be directly relevant.

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