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Inflation in the Euro Area?**

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Does money Granger cause inflation in the euro area?^(a)

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Abstract

In this paper we re-evaluate the empirical evidence on money-inflation Granger causality for the euro area and, in contrast to Trecroci and Vega (2000), conclude that money does in fact Granger cause inflation. We also show that it takes about a year and a half for changes in money growth to start passing on to inflation and five years for the whole adjustment to take place.

1. Introduction

Under the first pillar of the monetary policy strategy, the ECB attributed a prominent role to money in explaining future price developments. To signal this prominent role of money, the ECB Council has announced, since December 1998, a reference value for the rate of growth of the M3 aggregate. As the ECB puts it “the announcement of the reference value represents a visible public commitment on the part of the Governing Council to analyse and explain monetary developments and their implications for the risks to price stability in detail”¹.

An important characteristic of the reference value is that it must be consistent with the price stability hypothesis (i.e., less than 2 per cent annual inflation) and with the long run relation between money, prices and output, for a given annual growth rate of potential output. Therefore, the announcement of a reference value for money growth should imply that two fundamental requests are met: first, the existence of a stable money demand (both in the long run, to determine the reference value, and in the short run, to enable the ECB to explain deviations from that value)

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¹ ECB, Monthly Bulletin, November 2000.

and second, that money has leading indicator properties for future inflation. This paper addresses this second request.

So far the studies conducted by the ECB concerning the leading indicator properties of money addressed this issue either by performing Granger causality tests or by evaluating the information content of money using the Stock and Watson (1998,1999) approach. The conclusions from these two alternative approaches are apparently in contradiction. While Trecroci and Vega (2000) (henceforth TV (2000)) conclude that the null of Granger non-causality cannot be rejected for the relevant monetary aggregate M3, Altimari (2001) concludes that in fact the monetary aggregates contain substantial information for forecasting price developments in the euro area.

This paper adds to this literature by re-evaluating the empirical evidence on money-inflation Granger causality tests and by identifying the relevant transmission lags. In what regards the Granger causality tests, methodologically the paper does not significantly depart from TV (2000) but uses a different data set. The main conclusion is that money does in fact Granger cause inflation in the euro area for most of the alternative estimated VAR models. This new empirical evidence reinforces the role of money as a leading indicator of inflation. It is also shown that it takes about a year and a half for changes in money growth to start passing on to inflation and that the adjustment is completed by the end of the fifth year. This evidence should be taken into account when evaluating the short to medium term consequences of monetary developments.

The remaining part of this paper is organised as follows. Section 2 reviews the empirical evidence on the leading indicator properties of money so far produced by the ECB. Section 3 describes the data used in the empirical section. Section 4 presents the statistical methodology used in the paper. Section 5 reports the new empirical evidence on money-inflation Granger causality tests for the euro area. Section 6 identifies the relevant transmission lags in simple money-prices dynamic models defined in the year-on-year growth rates, and section 7 concludes.

2. Overview of the empirical literature for the euro area

As abovementioned the ECB has so far produced two main papers on the leading indicator properties of money: TV (2000) and Altimari (2001). TV (2000) evaluate the money-prices causality tests within the model identified in Coenen and Vega (1999) [CV(1999)], which is a VAR defined in the 5 variables: real money stock ($m-p$) (where m is the natural log of the

monetary aggregate M3 and p the GDP deflator), inflation (Δp), real GDP (y), short (s) and long-term (l) interest rates. The causality tests are performed on the equation for Δp using the tests proposed in Toda and Phillips (1993, 1994) [henceforth TP(1993,1994)] and in Toda and Yamamoto (1995) [henceforth TY(1995)]. Besides the general VAR defined in the 5 variables above, TV (2000) also test for Granger causality in almost all the subsystems that can be obtained as special cases by dropping some of the variables. Specifically, they carried out causality tests by applying the TP (1994) and TY (1995) approaches on a large number of VAR models defined by taking in turn the following sets of variables: $(m-p, \Delta p, y, l, s)$, $(m-p, \Delta p, y, l-s)$, $(m-p, \Delta p, y, s)$, $(m-p, \Delta p, y)$ and $(m-p, \Delta p)$. Besides these models TV (2000) also applied the TY(1995) approach on the unrestricted versions of these models, i.e., the VAR models obtained without imposing long run homogeneity between money and prices (and/or interest rates): (m, p, y, l, s) , $(m, p, y, l-s)$, (m, p, y, s) , (m, p, y) , (m, p) . In addition, under the TY (1995) approach the hypothesis of prices and money being I(2) or I(1) were alternatively considered. For none of these 10 models could the null of money-prices Granger non-causality be rejected. Thus, the authors conclude: “there appears to be little empirical support for rejecting at standard confidence levels Granger non-causality of m on p within this information set”.

TV (2000) then proceed by analysing an extension of a p-star model of inflation closely following Gerlach and Svensson (2000) [GS(2000)] and conclude that the “real money gap” (i.e., the gap between current real balances and long run equilibrium real balances) has substantial predictive power for future inflation².

More recently, Altamari (2001) investigated the properties of monetary and credit aggregates as indicators of future price developments in the euro area. This author followed Stock and Watson’s methodology (1998 and 1999), by comparing the out-of-sample forecasting performance of a large number of models based on monetary and non-monetary indicators (forecasts are made using only the information available prior to the forecasting period, with the forecasting horizons varying from one quarter to three years ahead). The author considered several monetary indicators, including the various concepts of money (M1, M2 and M3), credit and money based indicators, such as the “real money gap” and the “monetary overhang”, and

² GS (2000) conclude for the superiority of the real money gap compared to the output gap and the nominal money gap (which they call the Eurosystem’s money-growth indicator), as an information variable for future inflation. However in a more recent version of the paper (see Gerlach and Svensson (2002)) the authors change the conclusion in that now both the output gap and the real money gap appear as containing considerable information regarding future inflation (the nominal money gap still appears as containing little information about future inflation).

performed the exercise for different measures of inflation, different sample periods and different information sets (bivariate and multivariate). Altimari (2001) emphasises three conclusions: 1) monetary and credit aggregates contain substantial information for forecasting future price developments in the euro area, being comparatively advantageous in relation to other non-monetary indicators for longer forecast horizons; 2) indicators derived within the P-star framework, including the “real money gap” and the “monetary overhang” appear to perform well, but, contrasting with the findings by TV (2000) and Gerlach and Svensson (2000), the “real money gap” should not be the preferable focus of policy; 3) money contains additional, independent information, beyond the information contained in the usual determinants considered in money demand relationships.

To sum up, the evidence so far found by the ECB suggests that the existence of leading indicator properties of money to prices is still an open issue. While there seems to exist unequivocal empirical evidence that for long horizons money helps to predict prices (in line with theory), it is still highly debatable that for the horizon relevant for monetary policy (around two years) money and money-based indicators provide useful information.

This paper adds to this evidence by computing Granger causality tests in line with TV (2000) but using a different data set, which is discussed in the next section.

3. The data

TV (2000) carried out the causality tests with the data set previously used in CV (1999), which differs from the one used later in Brand and Cassola (2000) [henceforth BC (2000)]. In the CV(1999) data set, the series for the nominal stock of money, real GDP, GDP deflator and interest rates for the euro-area were constructed using fixed weights based on 1995 GDP at PPP rates. In turn, the series for the nominal stock of money, real and nominal GDP used in BC (2000) were constructed using the irrevocable fixed exchange rates of 31 December 1998 and the GDP deflator computed as the implicit GDP deflator obtained as the ratio between nominal and real GDP.

Figure 3.1 depicts the two series for the inflation rate, computed as the first difference of logged prices. It is readily seen that the inflation rate in the CV data set is systematically above the

Figure 3.1- Inflation in the euro area

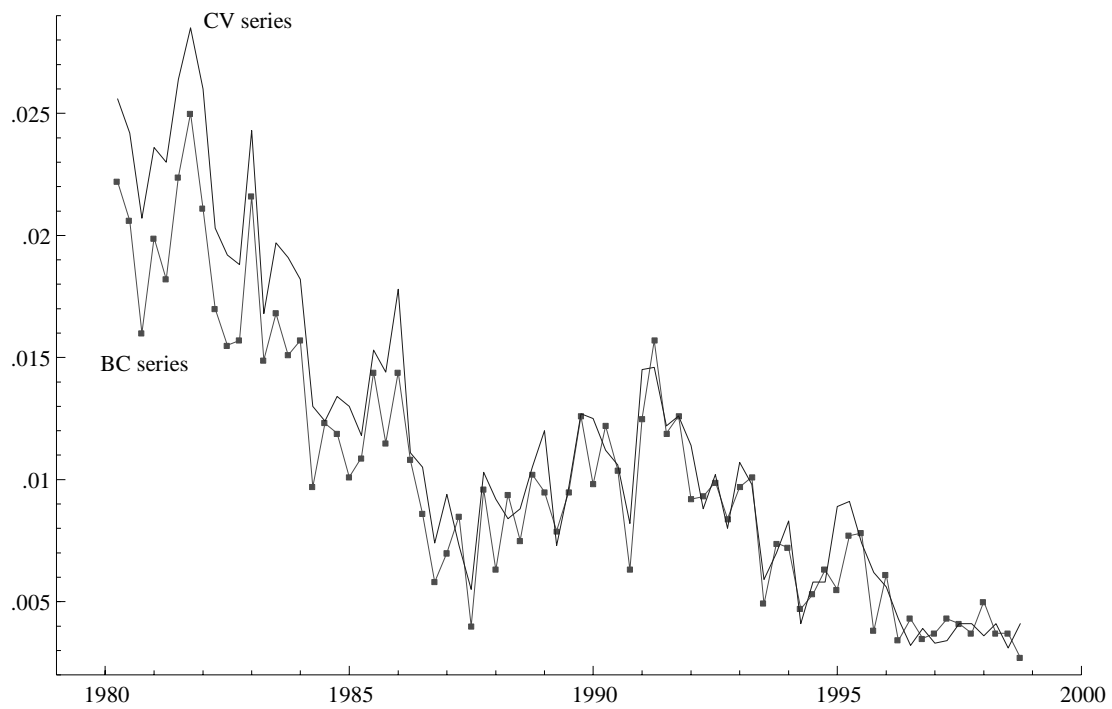


Figure 3.2- Money growth

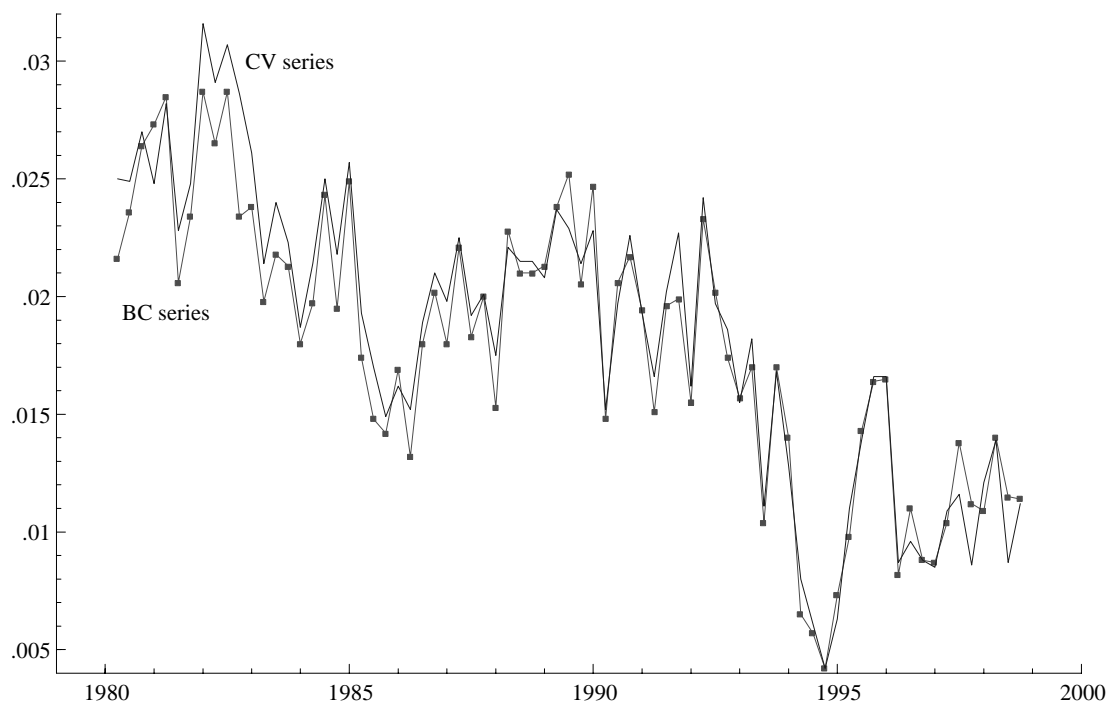
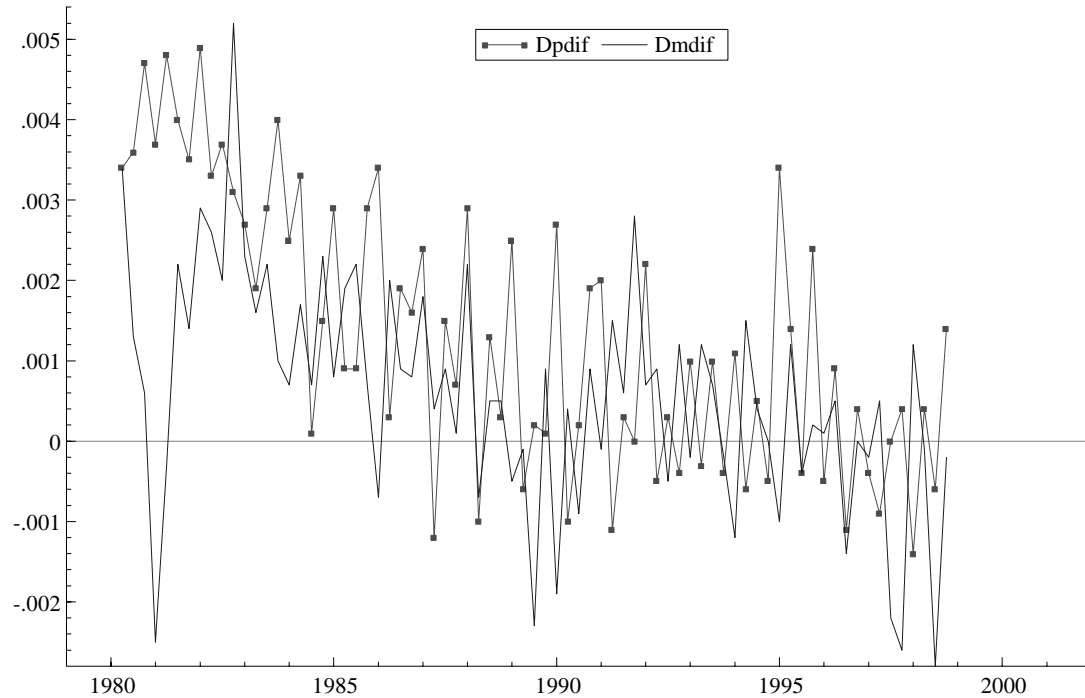


Figure 3.3- Difference in the CV series vis-à-vis the BC series



inflation rate in BC data set from the beginning of the sample until 1988. Figure 3.2 depicts the two series for money growth, computed as the first difference of logged M3. Also in this case, the CV series is most of the time above the BC series. However the difference is neither as large nor as systematic as for the inflation series (Figure 3.3).

In this paper we use an updated version of the BC (2000) data set, as this has become of widespread use in empirical studies for the euro area. We note that, in particular, the monetary aggregate M3 corresponds to the official data released by the ECB and was downloaded from the ECB's web page (long historical series seasonally adjusted).

We restricted the sample to the period 1980:1 to 2000:4 to avoid statistical problems with the entry of Greece and for comparability with TV (2000).

4. Causality tests: Toda and Phillips, Toda and Yamamoto and Phillips procedures.

This section briefly reviews the causality tests whose results are presented in the empirical section. These include the tests suggested in TP (1993, 1994), TY (1995) and Phillips (1995).

In order to introduce some notation consider the n -vector of $I(1)$ time series y_t generated by the k -th order VAR model:

$$y_t = \sum_{i=1}^k A_i y_{t-i} + \varepsilon_t \quad (4.1)$$

Under the assumption of r cointegrating vectors (4.1) may be reparameterised as a vector error correction model (VECM) given by:

$$\Delta y_t = \sum_{i=1}^{k-1} A_i^* \Delta y_{t-i} + \alpha \beta' y_{t-1} + \varepsilon_t \quad (4.2)$$

where $A_i^* = -\sum_{r=i+1}^k A_r$ and $\alpha \beta' = -(I - \sum_{i=1}^k A_i)$ with α and β (n, r) full-rank matrices of the loading factors and cointegrating vectors, respectively,

Now suppose that we want to test if there are causal effects from the last n_3 elements of y_t to the first n_1 elements of this vector, and accordingly partition y_t into three sub-vectors: $y_t = (y_{1t}, y_{2t}, y_{3t})$ such that y_{1t} , y_{2t} and y_{3t} are $(n_1, 1)$, $(n_2, 1)$ and $(n_3, 1)$ vectors, respectively, with $n_1 + n_2 + n_3 = n$. For ease of presentation let us further partition the A_i^* , α and β matrices conformably to y_t :

$$y_t = \begin{bmatrix} y_{1t} \\ y_{2t} \\ y_{3t} \end{bmatrix} \quad A_i^* = \begin{bmatrix} A_{i,11}^* & A_{i,12}^* & A_{i,13}^* \\ A_{i,21}^* & A_{i,22}^* & A_{i,23}^* \\ A_{i,31}^* & A_{i,32}^* & A_{i,33}^* \end{bmatrix} \quad \alpha = \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \end{bmatrix} \quad \beta = \begin{bmatrix} \beta_1 \\ \beta_2 \\ \beta_3 \end{bmatrix} \quad (4.3)$$

where α_1 denotes the first n_1 rows of the loading coefficient matrix α and β_3 the last n_3 rows of the matrix of the cointegrating vectors β .

The null hypothesis of non-causality based on model (4.2) can be formulated as:

$$H^*: A_{1,13}^* = A_{2,13}^* = \dots = A_{k-1,13}^* = 0 \quad \text{and} \quad \alpha_1 \beta_3' = 0 \quad (4.4)$$

Following TP (1993 and 1994), we will refer to the first part of the hypothesis (4.4) as “short run non-causality” and the second part as “long run non-causality”. To carry out the test in (4.4) these authors suggest a sequential procedure, based on three alternative strategies. Let us define the following tests:

$$H_{\perp}^*: A_{1,13}^* = A_{2,13}^* = \dots = A_{k-1,13}^* = 0 \quad (\text{short - run noncausality}) \quad (4.5)$$

$$H_1^*: \alpha_1 = 0 \quad (\text{weak - exogeneity}) \quad (4.6)$$

$$H_3^*: \beta_3 = 0 \quad (\text{long - run exclusion}) \quad (4.7)$$

$$H_{13}^*: \alpha_1 \beta_3' = 0 \quad (\text{long - run noncausality}) \quad (4.8)$$

Based on these tests, TP (1994) recommend the following three alternative strategies denoted (P1), (P2) and (P3):

$$(P1): \text{Test } H_1^* \cdot \begin{cases} \text{If } H_1^* \text{ is rejected, test } H^* \text{ using a } \chi_{n_3k}^2 \text{ critical value} \\ \text{Otherwise, test } H_{\perp}^* \end{cases}$$

$$(P2): \text{Test } H_3^* \cdot \begin{cases} \text{If } H_3^* \text{ is rejected, test } H^* \text{ using a } \chi_{n_3k}^2 \text{ critical value} \\ \text{Otherwise, test } H_{\perp}^* \end{cases}$$

$$(P3): \text{Test } H_{\perp}^* \cdot \begin{cases} \text{If } H_{\perp}^* \text{ is rejected, reject the null hypothesis of non - causality} \\ \text{Otherwise, test } H_1^* \text{ and } H_3^* \cdot \begin{cases} \text{If both are rejected test } H_{31}^* \text{ if } r > 1 \\ \text{or reject the null if } r = 1. \\ \text{Otherwise, accept the null of noncausality} \end{cases} \end{cases}$$

To better understand the rational of the proposed strategies it is perhaps useful to note that, as demonstrated in TP (1993), in order to apply conventional asymptotic chi-square tests using the Wald type statistics proposed by the authors we need one of the two following conditions to be

met under the null hypothesis: $rank(\alpha_1) = n_1$ or $rank(\beta_3) = n_3$ ³. That is why the strategies (P1) and (P2) start with the tests H_1^* or H_3^* , as only if one of these tests is rejected does the standard χ^2 limit distribution apply for the test H^* .

The strategies (P1) and (P2) are applicable when $n_1=1$ and $n_3>1$, but (P3) is applicable only when $n_1=n_3=1$. In fact, this strategy takes advantage of the fact that $n_1=n_3=1$, (i.e., both H_1^* and H_3^* are tested in the second step)⁴. Within this strategy we start by testing for “short-run non-causality”, and if this is not rejected we proceed by testing whether $\beta_3=0$ and whether $\alpha_1=0$. If both are rejected and $r=1$, we reject the null of non-causality. Otherwise we still need to test whether $\alpha_1\beta_3'=0$. Note however that in (P3) it does not make a difference whether we start by testing H_\perp^* or by testing H_1^* and H_3^* (and H_{13}^* if $r>1$), i.e., the results should be unchanged even though we alter the order of testing.

In our case, we just want to test for Granger non-causality from money to prices and vice-versa and thus we would always have $n_1=n_3=1$. Furthermore, as we shall see below, for the VARs defined in inflation and money growth (regardless whether we also introduce GDP growth as an exogenous regressor) we have $r=1$ (a single cointegrating vector) so that, in our case, the strategy (P3) with $r=1$ appears as particularly convenient. For presentation purposes the testing procedure suggested in TP(1993,1994) will be denoted the ECM approach as it resorts to the Johansen estimators⁵.

A useful alternative Granger causality test was suggested by TY (1995). This test directly applies to VARs in levels. Resorting to our VAR model (4.1) with variables in levels the non-causality null hypothesis can be formulated as

$$H_0^*: A_{1,13} = A_{2,13} = \dots = A_{k,13} = 0 \quad (4.9)$$

³ For instance, when $n_1 = n_3 = r = 1$, TP (1993) demonstrate that if $\alpha_1 = \beta_3 = 0$ then the Wald statistic for the non-causality hypothesis that $\alpha_1\beta_3 = 0$ has a limit distribution that differs from the χ^2 distribution, which is the limit distribution that we would obtain if either α_1 or β_3 are nonzero. Thus, before we apply conventional asymptotic chi-square tests to non-causality hypothesis, we would have to test empirically whether $rank(\alpha_1) = n_1$ or $rank(\beta_3) = n_3$.

⁴ Notice that if $n_1 = 1$ the condition $rank(\alpha_1) = n_1$ is equivalent to $\alpha_1 \neq 0$ and if $n_3 = 1$ the condition $rank(\beta_3) = n_3$ is equivalent to the condition $\beta_3 \neq 0$.

⁵ We note that the statistics used in the tests (4.4)-(4.8) assume that the cointegrating vectors are normalised according to the suggestion in Johansen (1988).

where $A_{i,13}$ is the $n_1 \times n_3$ upper-right submatrix of A_i in (4.1). As we have in our case $n_1 = n_3 = 1$, the $A_{i,13}$ submatrices are just scalars given by the upper-right entry in the A_i matrices.

It is well known that the conventional Wald tests of restrictions on the coefficients of VARs in levels with I(1) and cointegrated variables generally have non-standard asymptotic properties. This precludes the use of the Wald tests directly on the null hypothesis given by (4.9) for the VAR (4.1) estimated in levels by OLS. TP (1993) have shown that the two noteworthy exceptions occur when i) there is cointegration and $rank(\beta_3) = n_3$, in which case the Wald statistic would have an asymptotic χ^2 distribution and ii) when there is no cointegration, in which case the limit distribution is non-standard, but free of nuisance parameters⁶.

However, TY (1995) have developed a simple device that allows testing for Granger non-causality in levels VARs estimated by OLS with integrated variables. In fact they have shown that if the true model for the (possibly) nonstationary vector y_t is a VAR(k) and we instead fit a VAR(k+d) by OLS, where d is the maximal order of integration that we suspect might occur in the y_t process, then the usual Wald statistic for Granger non-causality based on levels estimation has an asymptotic chi-square distribution. Thus, with this device the tests may be performed directly on the least squares estimators of the coefficients of the VAR process specified in the levels of the variables⁷.

The approach suggested in TY (1995) is applicable whether the VAR's are stationary (around a deterministic trend), integrated of an arbitrary order, or cointegrated of an arbitrary order. For presentation purposes will be denoted the lag augmented Var approach, or simply the LA-VAR approach.

The main advantage of this approach is that it does not require any pretesting and so allows one to pay little attention to the integration and cointegration properties of the time series data in hand. In fact, in most applications it is not known *a priori* whether the variables are (trend) stationary, integrated or cointegrated and so, within the approach suggested in TP (1993, 1994), pretests for

⁶ In this case the critical values for the causality tests in levels VARs can be tabulated conveniently. But, of course, if it is known that the system is I(1) with no cointegration, causality tests based on differences VARs are also valid, and in this tests the usual chi-square critical values are employed. Furthermore, these tests in differences VARs are likely to have higher power in finite samples, as the non-causality hypothesis in levels (4.9) contain redundant parameter restrictions. For further details see TP (1993).

⁷ On this issue see also Dolado and Lutkepohl (1996).

unit root(s) and cointegration are usually required before estimating the VAR model in which statistical inferences are conducted (VECM). Now if some of these tests are not reliable in small samples, the non-causality tests conditioned on unit roots and cointegration tests may suffer from important pretest biases.

Of course, the TY (1995) approach is also not free from power problems. For instance, suppose that there is uncertainty whether the variables are I(1) or I(0). In order to avoid the pretest we must fit the model with an extra lag. In this case there would be a loss of power. If the variables are I(0) we are in fact introducing an irrelevant additional lag and if the variables are I(1) we are in fact throwing away both the possibility of no cointegration (in which case the use of a model in differences would deliver tests with higher power) or the possibility of cointegration (in which case some coefficients or linear combinations of them may be estimated more efficiently with larger rate of convergence and some redundant restrictions may be avoided).

Finally, as a third Granger causality test we also performed the test suggested in Phillips (1995). This procedure, denoted below as the FM-VAR approach, uses the fully-modified OLS estimator for VAR models. Let us assume again our equation (4.2). If we write $\pi_{13} = \alpha_1 \beta'_3$ the non-causality hypothesis (4.4) reads as

$$H^*: A_{1,13}^* = A_{2,13}^* = \dots = A_{k-1,13}^* = 0 \quad \text{and} \quad \pi_{13} = 0 \quad (4.10)$$

The FM-VAR approach tests (4.10) through a Wald statistic developed in Phillips (1995). The test is valid regardless of whether the variables in the VAR are stationary or integrated (of order one) and in this latter case regardless whether they are cointegrated or not. However, as the asymptotic distributions of the Wald statistic varies according to each case, we just carried out a test with a conservative size (asymptotically) by taking a critical value from the Chi-square distribution which is an upper bound for the true values of the limit distributions of the Wald statistic that apply in each specific situation⁸.

Recently Yamada and Toda (1998) carried out a very informative exercise, which compares the power and size of the three alternative approaches to causality testing: the ECM, the LA-VAR and the FM-VAR approach. The authors conclude that none of the three approaches emerges as

⁸ One important assumption of the test concerns the bandwidth parameter used in the Kernel estimators of the long run covariance matrices used in the computations of the FM-OLS estimators. For the results in Phillips (1995) to apply this parameter must be between $T^{1/4}$ and $T^{2/3}$. With a sample of 84 observations this rule implies a minimum of 3 and a maximum of 19.

clearly superior to the other two. In what concerns the size of the tests, the LA-VAR approach exhibits a stable size for sample sizes that are typical for economic time series and in this respect this approach excels the other two. The ECM approach performs satisfactorily when the cointegrating rank of the system is accurately detected (because this test is conditional on the choice of the cointegrating rank), but the FM-VAR approach is very sensitive to the values of certain parameters and in some cases a large size distortion may occur. In what concerns the power of the three procedures, the authors conclude that the FM-VAR is always as powerful as or more powerful than the other two. In turn the ECM is more powerful than the LA-VAR except for some combinations of the parameter values.

All in all, the above discussion suggests that we should look at more than the results of a single approach. Thus, in the following section we report the outcome of the Granger causality tests using the three testing strategies discussed above: the ECM, the LA-VAR and the FM-VAR approaches.

5. Empirical findings

This section discusses the VAR systems used in testing for Granger-causality between money and prices and the key results obtained, with a special reference to previous evidence. A robustness analysis is also provided.

In what follows m_t stands for the natural log of the nominal M3 money stock, p_t for the natural log of the GDP deflator, y_t for the natural log of real GDP, l and s for the long and short run interest rates respectively.

5.1 Discussion of the set-up

The estimated VARs are basically of two types. The first type is composed of VAR models where all the variables are integrated of order one. It includes the $\text{VAR}(m-p, \Delta p, y, l, s)$ as the general case and the $\text{VAR}(m-p, \Delta p, y)$ and $\text{VAR}(m-p, \Delta p)$, as special cases. The second type is composed of VAR models where some variables are assumed as potentially integrated of order two (m and p) and the others as integrated of order one (y, l, s). It includes the $\text{VAR}(m, p, y, l, s)$ as the general case and the $\text{VAR}(m, p, y)$ and the $\text{VAR}(m, p)$, as special cases.

To the first set of VARs all three available testing approaches are applicable, while for the second only the TY (1994) approach can be performed, as the TP (1994) and Phillips (1995) approach were designed for I(1) systems only.

The choice of using both sets of VARs is motivated by the comparability to TV (2000). We discuss the implications for the results of their homogeneity assumption and the possible sensitivity of the results to the small sample performance of the testing approaches.

Besides the abovementioned two sets of VAR models we also pay attention to the simpler VAR($\Delta m, \Delta p$) and VAR($\Delta m, \Delta p; \Delta y$) models. These two models can be seen as special cases of the VAR(m, p, y, l, s) type of models. But in these cases all the three testing strategies are applicable as they are I(1) systems.

TV (2000), even though having also estimated small VAR models, appear to attribute special importance to the outcome of the Granger causality tests for the general VAR ($m - p, \Delta p, y, l, s$) model. The VAR ($m - p, \Delta p, y, l, s$) model is probably a suitable instrument when the purpose of the analysis is the identification and estimation of a stable money demand function. This is clearly the case of the VAR models used in CV (1999) and BC (2000). However, if the single purpose of the analysis is the study of money-prices causality the use of simpler models clearly offers parsimony. Actually the use of large VAR systems, containing all the potentially relevant variables, may imply less informative conclusions, particularly, in small samples, due to the increased variance around the estimates. In our case it may be argued that interest rates are not expected to have significant direct effects on prices, as the main transmission channels from interest rates to prices include either activity or money as the relevant intermediate variables. If this is the case we would expect the simpler VAR ($m - p, \Delta p, y, l$) models to allow more robust conclusions on money Granger causality tests⁹.

On the other hand, the approach followed in TV (2000), while statistically sophisticated, seems to have overlooked a point in the estimated VAR models that may influence the results in some dimension. To see how the argument follows let us take the simplest model estimated in TV (2000). The authors claim that the VAR (m, p) reduces to the VAR ($m - p, \Delta p$) if long run homogeneity holds. We argue that the equivalence between the two models holds for the long run

⁹ Of course against this view it may be argued that interest rates are expected to affect prices through changes in marginal costs. To test for this possibility we would need to estimate a model in which a proxy for such a channel is introduced.

but not for the short run so that working with VAR models defined in the variables $(m-p, \Delta p)$ may have damaging consequences for Granger causality tests. To see this point in some detail let us take the VAR (m, p) model with four lags:

$$\begin{aligned} m_t &= a_{11}m_{t-1} + \dots + a_{14}m_{t-4} + b_{11}p_{t-1} + \dots + b_{14}p_{t-4} + \varepsilon_{1t} \\ p_t &= a_{21}m_{t-1} + \dots + a_{24}m_{t-4} + b_{21}p_{t-1} + \dots + b_{24}p_{t-4} + \varepsilon_{2t} \end{aligned} \quad (5.1)$$

Assuming, as in CV (1999) and TV (2000), that prices (p) and nominal money (m) are I(2), and that long run homogeneity holds (i.e., that $(m-p)$ is I(1)) it is straightforward to show that model (5.1) may be reparameterised in the error correction form as:

$$\begin{aligned} \Delta^2 m_t &= \alpha_{11}\Delta^2 m_{t-1} + \alpha_{12}\Delta^2 m_{t-2} + \beta_{11}\Delta^2 p_{t-1} + \beta_{12}\Delta^2 p_{t-2} \\ &\quad + \gamma_{11}\Delta m_{t-1} + \gamma_{12}\Delta p_{t-1} + \mu_1(m_{t-1} - p_{t-1}) + \varepsilon_{1t} \end{aligned} \quad (5.2)$$

$$\begin{aligned} \Delta^2 p_t &= \alpha_{21}\Delta^2 m_{t-1} + \alpha_{22}\Delta^2 m_{t-2} + \beta_{21}\Delta^2 p_{t-1} + \beta_{22}\Delta^2 p_{t-2} \\ &\quad + \gamma_{21}\Delta m_{t-1} + \gamma_{22}\Delta p_{t-1} + \mu_2(m_{t-1} - p_{t-1}) + \varepsilon_{2t} \end{aligned}$$

It can readily be seen that besides the so called “short term dynamics” accounted for by the variables in second differences the model includes the usual levels variables in the error correction term and an error correction term with variables in first differences, whose coefficients are, in principle, free from any *a priori* restrictions. Assuming that $(m-p)$ is I(1), cointegration requires the I(1) variables $\Delta p_t, \Delta m_t$ and $(m-p)_t$ to be cointegrated.

The relevant question is whether model (5.1)-(5.2) can be replaced by the VAR $(m-p, \Delta p)$ ¹⁰. The VAR (m, p) reduces to the VAR $(m-p, \Delta p)$ in what concerns the long run analysis. However, an exact equivalence between the two models does not seem to hold in what concerns the short-term dynamics. To see that notice that the VAR $(m-p, \Delta p)$ with three lags,

$$\begin{aligned} (m-p)_t &= a_{11}(m-p)_{t-1} + \dots + a_{13}(m-p)_{t-3} + b_{11}\Delta p_{t-1} + \dots + b_{13}\Delta p_{t-3} + \varepsilon_{1t} \\ \Delta p_t &= a_{21}(m-p)_{t-1} + \dots + a_{23}(m-p)_{t-3} + b_{21}\Delta p_{t-1} + \dots + b_{23}\Delta p_{t-3} + \varepsilon_{2t} \end{aligned} \quad (5.3)$$

may be reparameterised in the error correction form as

¹⁰ This issue is raised, for instance, in Johansen (1992).

$$\begin{aligned}\Delta(m-p)_t &= \alpha_{11}\Delta(m-p)_{t-1} + \alpha_{12}\Delta(m-p)_{t-2} + \beta_{11}\Delta^2 p_{t-1} + \beta_{12}\Delta^2 p_{t-2} \\ &\quad + \gamma_{12}\Delta p_{t-1} + \mu_1(m_{t-1} - p_{t-1}) + \varepsilon_{1t}\end{aligned}\tag{5.4}$$

$$\begin{aligned}\Delta^2 p_t &= \alpha_{21}\Delta(m-p)_{t-1} + \alpha_{22}\Delta(m-p)_{t-2} + \beta_{21}\Delta^2 p_{t-1} + \beta_{22}\Delta^2 p_{t-2} \\ &\quad + \gamma_{22}\Delta p_{t-1} + \mu_2(m_{t-1} - p_{t-1}) + \varepsilon_{2t}\end{aligned}$$

and further as

$$\begin{aligned}\Delta^2 m_t &= \alpha_{11}^*\Delta^2 m_{t-1} + \beta_{11}^*\Delta^2 p_{t-1} + \beta_{12}^*\Delta^2 p_{t-2} \\ &\quad + \gamma_{11}^*\Delta m_{t-1} + \gamma_{12}^*\Delta p_{t-1} + \mu_2^*(m_{t-1} - p_{t-1}) + \varepsilon_{1t}\end{aligned}\tag{5.5}$$

$$\begin{aligned}\Delta^2 p_t &= \alpha_{21}^*\Delta^2 m_{t-1} + \beta_{21}^*\Delta^2 p_{t-1} + \beta_{22}^*\Delta^2 p_{t-2} \\ &\quad + \gamma_{21}^*\Delta m_{t-1} + \gamma_{22}^*\Delta p_{t-1} + \mu_2^*(m_{t-1} - p_{t-1}) + \varepsilon_{2t}\end{aligned}$$

It is readily seen that while model (5.2) includes two lags of $\Delta^2 m_t$ model (5.5) includes only one. In fact model (5.5) is a special case of model (5.2) that one obtains imposing the restrictions $\alpha_{12} = \alpha_{22} = 0$ ¹¹. Thus by working with the VAR $(m-p, \Delta p)$ we are introducing untested restrictions on the short-term dynamics vis-à-vis the general VAR (m, p) with homogeneity. Apparently the two models do not differ in what concerns the long run properties. So, if the purpose of the analysis is the assessment of the long run relationship between the money stock and the price level that restriction may not be an important issue. However, if we aim at testing Granger causality following the TP(1994) approach, which tests separately for “short run” and “long run” causality, this untested restriction may condition the outcome of the tests. In any case, the point estimates for the α_{ij}^* , β_{ij}^* , γ_{ij}^* and μ_j^* coefficients in (5.5) may significantly differ from their counterparts in (5.2) if the imposed restrictions on the short-term dynamics are not data consistent.

We also notice that model (5.1)-(5.2) encompasses the VAR($\Delta m, \Delta p$) as a special case if we impose the restrictions $\mu_1 = \mu_2 = 0$. Working with model (5.1)-(5.2) unrestrictedly raises some practical difficulties as it requires an I(2) cointegration analysis. To circumvent the problem TV (2000) applied only the TY (1995) approach to the VAR (m, p) , which does not require any

¹¹ Of course if we start with the VAR $(m-p, \Delta p)$ with four lags we will end up in (5.5) with an equation for $\Delta^2 p_t$ with two lags of $\Delta^2 m_t$ but three lags of $\Delta^2 p_t$. The point is that in fact there is no exact equivalence between model (5.1)-(5.2) and model (5.3)-(5.5).

cointegration analysis. But in order to carry out the TP (1994) approach we need to restrict the general model so that it may be analysed within the I(1) apparatus developed by these authors. Estimating both the VAR $(m-p, \Delta p)$ and the VAR $(\Delta m, \Delta p)$ appears a natural solution to this problem. TV (2000) estimated the VAR $(m-p, \Delta p)$ but not the VAR $(\Delta m, \Delta p)$. However, the VAR $(\Delta m, \Delta p)$ arises quite naturally as the simplest VAR model for testing money-prices Granger causality. In fact, in many situations we want to know whether just by looking at money growth we are able to infer something about the likely future path of inflation. In other words, this simple VAR allows answering the very often-repeated question of whether money growth (not money stock) Granger causes inflation (not prices). This constitutes an important reason why we pay special attention to this simple model below¹². And, because it might be interesting to test whether the conclusions with this simplest model still hold should we account for GDP growth we also estimate the VAR $(\Delta m, \Delta p; \Delta y)$ model where GDP growth enters as an exogenous regressor.

5.2 General results and robustness

The results on money-prices Granger causality tests for all the systems are summarised in the Table 5.1, where we bring together the findings in TV (2000).

The reader interested in the full discussion of the output for each estimated VAR model is referred to Appendix 1.

First, we notice that the overall picture is now very different from the one obtained in TV (2000). In fact, for most of the studied cases the null of Granger non-causality of m on p or Δm on Δp is rejected. Only for the general VAR with interest rates it is not possible to reject the null of Granger non-causality. But, as previously argued, this might not be the best model to address Granger causality tests as it appears to be too large a system and thus with accrued power problems.

Obviously this different conclusion stems basically from the fact that we have worked with a different data set. However, should Trecroci and Vega estimate the VAR models in the growth

¹² As we shall see below it turns out that the VAR $(\Delta m, \Delta p)$ exhibits cointegration, but the VAR $(m-p, \Delta p)$ clearly does not. This outcome is easily understood as the simple fact that the real money stock $(m-p)$ is I(1) implies that $(\Delta m - \Delta p)$ is I(0) or in other words that Δm_t and Δp_t are cointegrated with a unit coefficient. But we do not expect the real money stock $(m-p)$ and the inflation rate Δp_t to cointegrate because they exhibit quite different time paths.

rates of the variables $(\Delta m, \Delta p)$ and $(\Delta m, \Delta p; \Delta y)$ and their picture on money-prices Granger causality tests might have been less negative.

Table 5.1-Summary of money-prices causality tests
Comparison with TV(2000)^(a)

Our results		Trecroci and Vega (2000)	
ECM approach	LA-VAR approach	ECM approach	LA-VAR approach
VAR($\Delta m, \Delta p$)			
YES	YES	YES ^(b)	NO ^(b)
VAR($\Delta m, \Delta p; \Delta y$)			
YES	YES	YES ^(b)	NO ^(b)
VAR($m - p, \Delta p$)			
---	YES	---	NO
VAR($m - p, \Delta p, y$)			
YES/NO	YES	NO	NO
VAR($m - p, \Delta p, y, l, s$)			
NO	YES	NO	NO
VAR(m, p)			
---	YES	---	NO
VAR(m, p, y)			
---	YES	----	NO
VAR(m, p, y, l, s)			
---	NO	---	NO

(a) The usable sample period actually taken for the computations is 1980/4-2000/4, except otherwise stated.

(b) Our computations for the period 1980/4-1998/4 using CV(1999) data set.

As a second point we notice that, as can be seen in Appendix 1, Δm and Δp are cointegrated but $(m - p)$ and Δp are not. This suggests that the VAR($\Delta m, \Delta p$) is more appropriate than the VAR($m - p, \Delta p$) to analyse the Granger non-causality hypothesis. The money-prices Granger non-causality hypothesis is strongly rejected in the VAR($\Delta m, \Delta p$) and VAR($\Delta m, \Delta p; \Delta y$) models. However, the evidence against Granger non-causality clearly weakens as we move to the

$\text{VAR}(m - p, \Delta p, y, l, s)$ type of models. As discussed above it may be the case that these type of VAR models are imposing important restrictions on the short run dynamics or lacking parsimony or both, with important consequences for the Granger causality tests.

Finally, the null of Granger non-causality from prices to money is basically rejected in the same models in which money to prices non-causality is. We note, however, that it is mainly the long run causality that appears as significant. Again, a full discussion of the results can be found in Appendix 1.

Robustness analysis

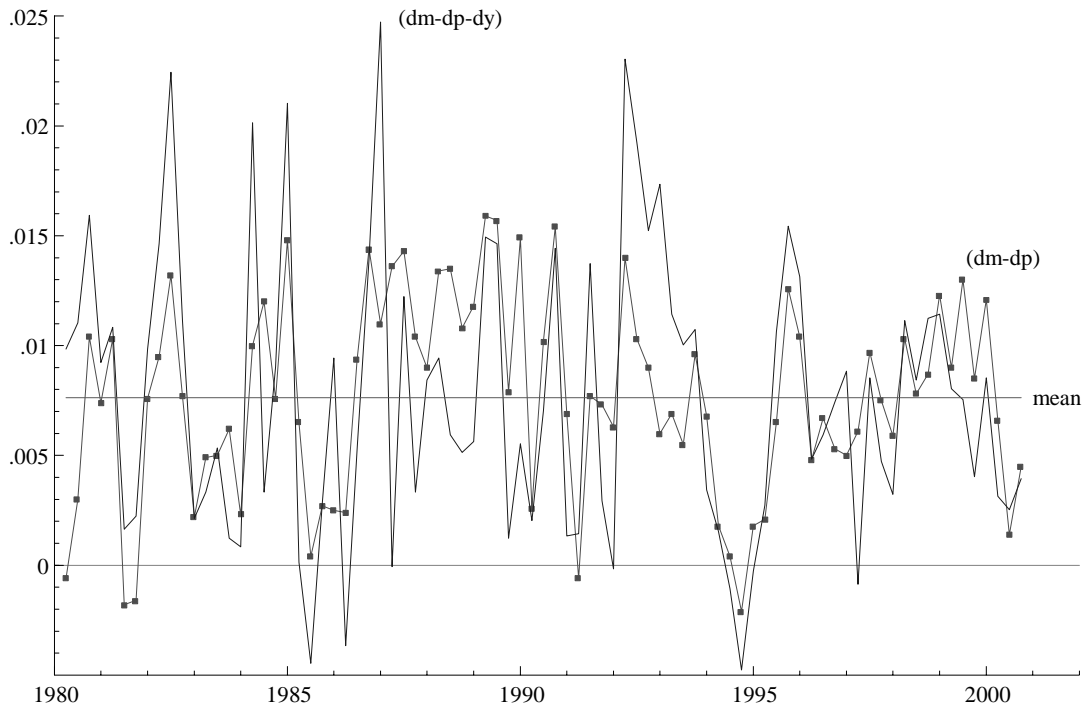
In order to get an idea on the robustness of the conclusions presented above we repeated the analysis for the 1985-2000 period for the models $\text{VAR}(\Delta m, \Delta p)$, $\text{VAR}(\Delta m, \Delta p; \Delta y)$ and $\text{VAR}(m - p, \Delta p, y)$. For none of the estimated models do the main conclusions change. Money to prices Granger non-causality is strongly rejected according to the ECM approach and rejected at around 10% according to the LA-VAR approach. Prices to money Granger non-causality is also rejected according to the ECM approach for the three estimated models, and to the LA-VAR approach for the two first models. See Table A2.1 in Appendix 2.

6. How shall we read the money-prices causality evidence?

At this stage a natural question arises: How useful is the evidence on money to prices Granger causality? With no doubt a major conclusion in this paper is that $(\Delta m - \Delta p)$ is clearly a stationary variable and that money growth Δm_t , to some extent, leads inflation. However, if we look at Figure 6.1 we realise that $(\Delta m - \Delta p)$ may stay above (or below) the mean for quite a large period of time. Particularly, this was so, between 1986/3 and 1990/1 (15 consecutive quarters!). If instead we look at the real money growth adjusted for GDP growth, $(\Delta m - \Delta p - \Delta y)$, we realise that the mean reversion have increased (specially in the second half of the eighties), but on the other hand the volatility has also increased (throughout the whole sample period) due to the volatility of the Δy_t series. This suggests that looking to $(\Delta m - \Delta p)$ or to $(\Delta m - \Delta p - \Delta y)$ is probably not the best way to draw interesting conclusions, in what concerns the money-prices relationship. It may also be argued that probably it is more useful to look at the year-on-year growth rates instead, as economic agents or at least economic analysts in their price and monetary

assessments look at these growth rates more, than they do at the monthly or quarterly growth rates.

Figure 6.1
 $(\Delta m - \Delta p)$ and $(\Delta m - \Delta p - \Delta y)$
 (Mean adjusted figures)

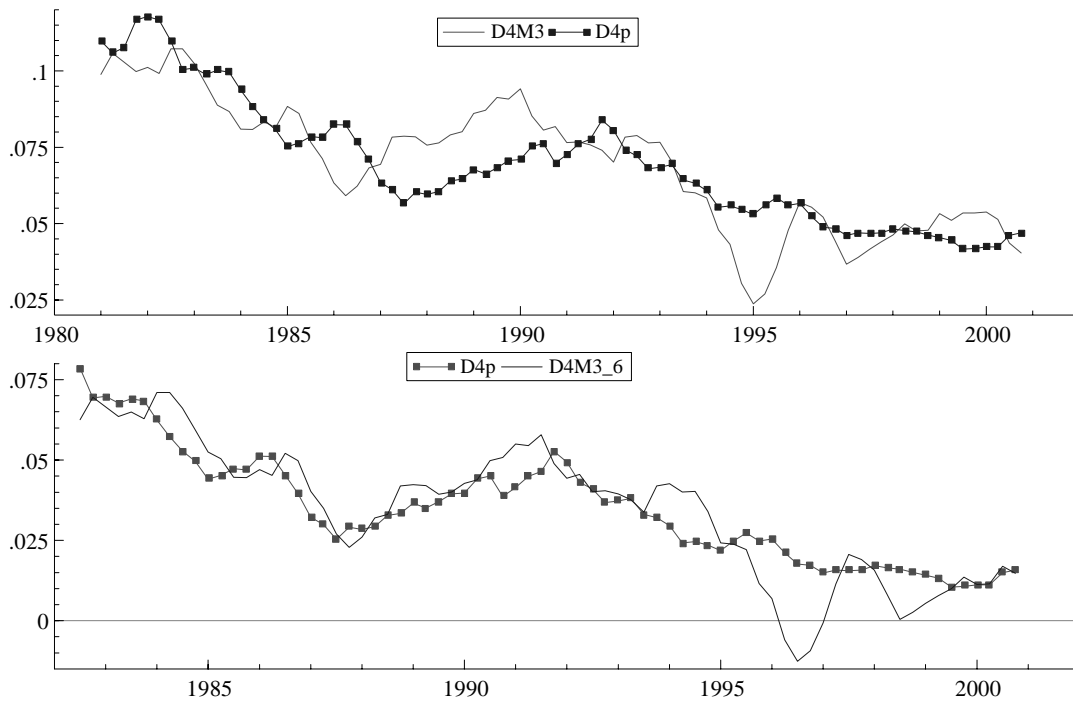


This section tries to answer the following question: if we look at the year-on-year money and GDP growth rates, what should we expect the reaction of the inflation rate to be? How many lags, if any, should we expect it to take for money growth to pass on to inflation? What role does GDP growth plays in such a relation?

In order to try to answer the questions above we start by looking at the static relation between the year-on-year price changes measured as $\Delta_4 p_t$ and the year-on-year money growth rate measured as $\Delta_4 m_t$. After some experimentation it was possible to conclude that the maximum correlation between inflation and money growth occurs at lag six. As can be seen from Figure 6.2 $\Delta_4 p_t$ and $\Delta_4 m_{t-6}$ (mean adjusted) almost coincide. So, the empirical evidence accords with economic

theory, which suggests that the effects from money to prices take time to materialise and are expected to occur with a lag between one year and two years (four to eight quarters). This first result already tells us that in order to identify the relevant lags we must specify a model with a large number of lags¹³.

Figure 6.2
Prices and money growth rates (mean adjusted)



We think that a conventional VAR model in the variables $\Delta_4 m_t$ and $\Delta_4 p_t$ is probably not the most adequate framework to investigate the relevant lags in the money-prices relationship, especially if, as it is the case, we expect the relevant lags for the price and money equations not to coincide. In fact if the relevant lags in the two equations are not the same it is not possible to identify the correct model from the conventional unrestricted VAR model. Of course, different zero restrictions may be imposed in each equation, but this will imply the use of special estimation techniques, probably making more difficult the identification of the relevant lags. For

¹³ This of course does not necessarily call into question the VAR models estimated above to test for Granger causality, as we are now looking at year-on-year and not quarterly growth rates, and this may significantly change the relevant lags, in the identified VAR models.

these reasons, below we rather analyse the dynamics of the relationship between money and prices resorting to a single equation error correction model for $\Delta_4 p_t$.

In order to be sure that we would be able to correctly identify the relevant lags in the money-prices relationship we start with a very general over-parameterised ADL model with 13 lags in each variable. In a first step, the so-called “interim reparameterisation”¹⁴ was imposed in order to find the first relevant lag. In a second step, an ECM reparameterisation was estimated taking the levels variables in the first relevant lag identified in the first step. Following the well-known general-to-specific methodology we end up with the following parsimonious error correction specification:

$$\begin{aligned}
\Delta_4 p_t = & -0.005 - 0.592\Delta_4 p_{t-4} + 0.247\Delta_4 p_{t-6} - 0.248\Delta_4 p_{t-8} \\
& (-4.30) \quad (-6.14) \quad (2.77) \quad (-2.92) \\
& + 0.180\Delta_4 p_{t-10} - 0.143\Delta_4 m_{t-8} + 0.128\Delta_4 y_{t-10} \\
& (2.13) \quad (-2.41) \quad (3.16) \\
& - 0.082\Delta_4 y_{t-12} - 0.223\Delta_4 p_{t-4} + 0.136\Delta_4 m_{t-6} + 0.081\Delta_4 y_{t-4} \\
& (-2.14) \quad (-6.47) \quad (4.88) \quad (3.36)
\end{aligned} \tag{6.1}$$

Period: 1984/2-2000/4; $R^2 = 0.662$; $\sigma = 0.19\%$; $AR(4)=1.12(0.36)$; $AR(12)=0.74(0.70)$;

A major conclusion that may be drawn from equation (6.1) is that changes in money growth take six quarters before starting to pass on to inflation (4 lags in case of GDP growth).

Table 6.1 depicts the accumulated effects on $\Delta_4 p_t$ of a change in the money growth rate (permanent increase of 1 p.p. in $\Delta_4 m$). As expected given equation (6.1) it takes six quarters before the effect on inflation of a permanent increase in money growth becomes significant. The accumulated effect after two years (8 quarters) is only 43.53% but it is completed by the end of the fifth year.

It seems useful to note that this outcome on the relevant lags involved, basically accords with the empirical evidence obtained for other countries. For instance, thirty years ago, Friedman (1972), even though following a different statistical approach, found that the “the highest correlation for consumer prices was for money leading twenty months for M1, and twenty-three months for

¹⁴ This reparameterisation differs from the so-called error correction reparameterisation in that the variables in levels are those of the longest level rather than those of period (t-1). In the interim reparameterisation the coefficients of the short-term dynamics are “almost” the interim multipliers.

M2". More recently Bernanke et al.(1999) describe a two-year lag between policy actions and their effect on inflation as a “common estimate”(pp. 315-320).

Table 6.1

Effect on $\Delta_4 p_t$ of a 1 p.p. increase in $\Delta_4 m$

Number of quarters	Accumulated effect Percentage of total effect
6	22.30
8	43.53
12	87.87
16	90.52
20	99.89

7. Conclusions

This paper re-evaluates the evidence on money-prices Granger causality for the euro area. At the theoretical level the paper does not significantly depart from Trecroci and Vega (2000) but at the empirical level the data set of Brand and Cassola (2000) is used instead, as it has become of widespread use in empirical studies for the euro area.

In contrast to TV (2000), who could not reject the money to prices Granger non-causality null hypothesis for any of the estimated VAR models, we were able to reject this same null hypothesis for most of the estimated VAR specifications. This different new conclusion mainly stems from the fact that we use a different data set, but to some extent, also from the fact that we estimate variant VAR models not considered in Trecroci and Vega (2000), which do not impose restrictions on the short run dynamics and/or correspond to more parsimonious specifications.

This outcome obviously constitutes new interesting evidence as it reinforces the role of money as leading indicator of inflation, detected in Altamari (2000).

Finally, it is seen that it takes about a year and a half for changes in money growth to start passing on to inflation and that the adjustment is completed by the end of the fifth year. This outcome is in line with the empirical evidence for other countries.

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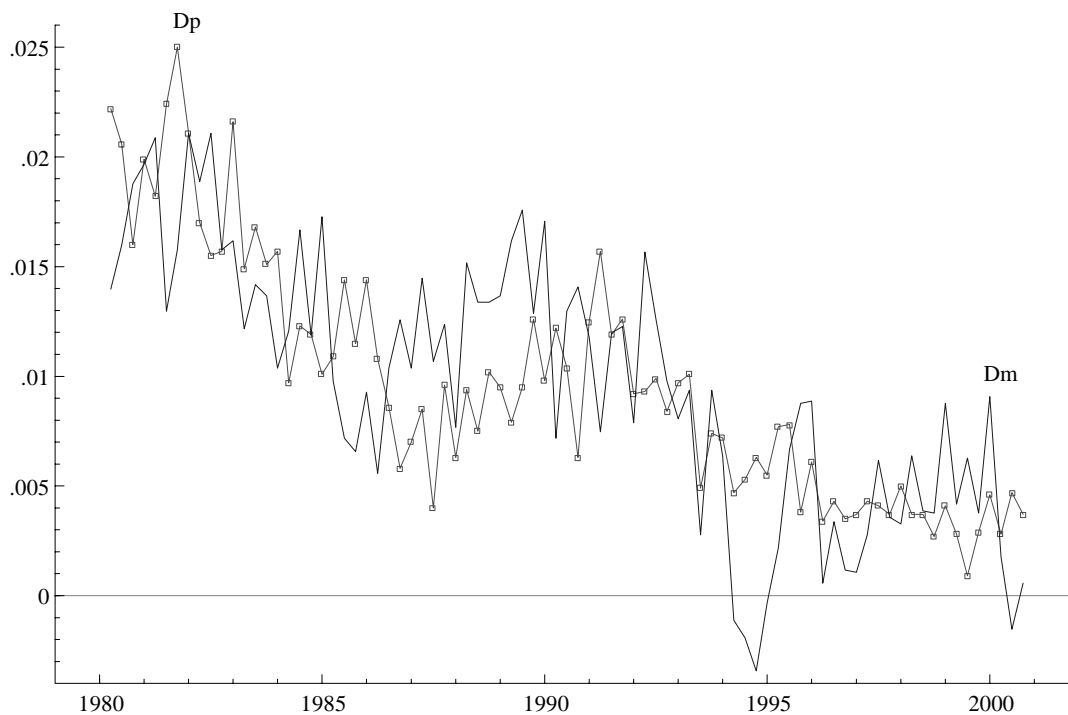
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Appendix 1: detailed results and full discussion of Granger non-causality tests

Evidence within the bivariate VAR ($\Delta m, \Delta p$) model

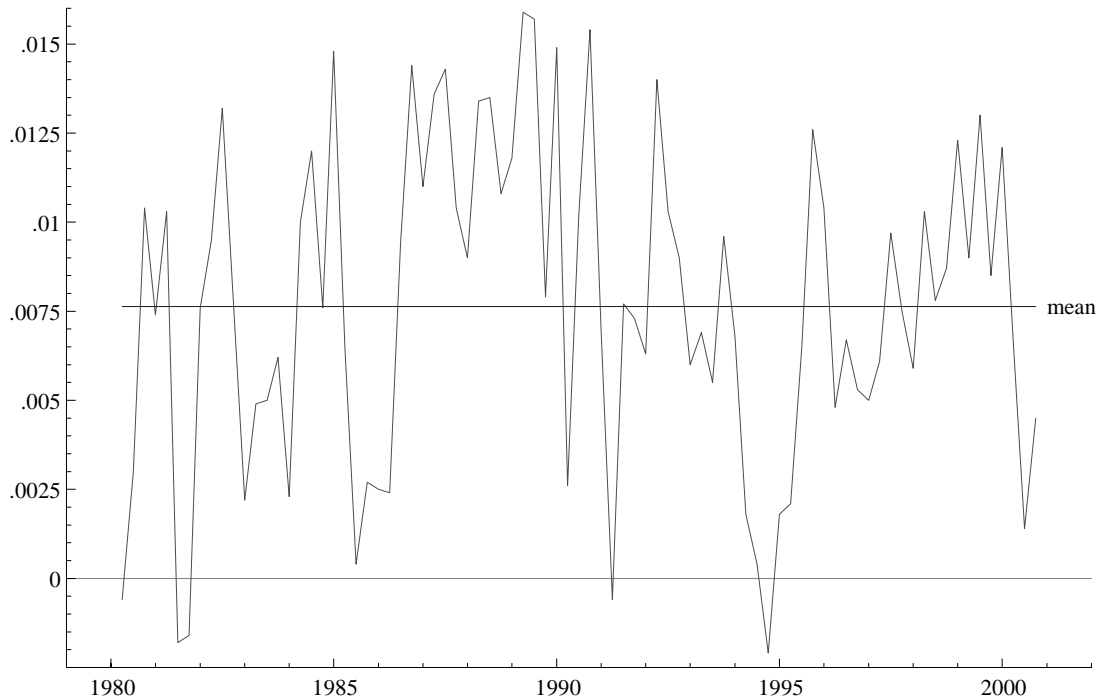
In order to allow a better understanding of the outcome of the tests within the simple VAR ($\Delta m, \Delta p$) we depict the two series in Figure A1.1 (after adjustment for the sample mean). Simple graphical inspection shows that both exhibit a clear non-stationary behaviour during the sample period. Thus, in accordance with previous evidence in CV (1999) and BC (2000), we also assume that they can be characterised as I(1) variables.

Figure A1.1
Inflation and money growth



Resorting to the usual lag selection criteria we conclude that 2 is the optimum number of lags, and according to the Johansen tests we conclude that the two variables are cointegrated¹⁵. The estimated cointegrating vector is $(\beta_{11}, \beta_{12}) = (1, -0.936)$ and the null of $\beta_{12} = -1$ is clearly not rejected by the data (the P-value of the test is 0.673). So we conclude that Δp_t and Δm_t are cointegrated with cointegrating vector $(1, -1)$, which means that $(\Delta m_t - \Delta p_t)$ is a stationary variable. This undoubtedly is a very important result as it implies that in the long run money growth and inflation evolve in line with each other¹⁶. Figure A1.2 depicts the $(\Delta m_t - \Delta p_t)$ variable, which exhibits a clear stationary behaviour around the sample mean.

Figure A1.2
 $(\Delta m - \Delta p)$



The estimated VAR in the ECM format reads as¹⁷:

¹⁵ For the trace test we obtain LR1=26.88 and for the maximum eigenvalue test we get LR2=22.81. The null of non-cointegration is rejected even at 1% tests (the critical values for a 1% test are 24.6 and 20.20 for the trace and maximum eigenvalue test, respectively).

¹⁶ This result is also a simple statistical implication of the fact that real money stock $(m - p)$ is I(1).

¹⁷ The estimates in equations (A1.1) and (A1.2) were obtained with EVIEWS for the period 1980/4-2000/4. t-statistics are shown in parenthesis.

$$\Delta^2 p_t = -0.286\Delta^2 p_{t-1} - 0.251\Delta^2 m_{t-1} - 0.255[\Delta p_{t-1} - 0.936\Delta m_{t-1} + 0.007] \quad (\text{A1.1})$$

(-2.91) (-3.49) (-3.70) (-6.99) (2.82)

$$\Delta^2 m_t = -0.384\Delta^2 p_{t-1} - 0.247\Delta^2 m_{t-1} + 0.342[\Delta p_{t-1} - 0.936\Delta m_{t-1} + 0.007] \quad (\text{A1.2})$$

(-2.65) (-2.32) (3.36) (-6.99) (2.82)

At first sight, informally, we note that in the two equations both the short-run as well as the long run coefficients appear to be significantly different from zero (the t statistics seem high enough). But of course, for a safe conclusion we need a formal statistical analysis. The results of the causality tests are depicted in Table A1.1.

Table A1.1 – VAR ($\Delta m, \Delta p$)

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^* = 17.76$ (0.000)	$\chi^2 = 8.73$ (0.013)	$H^* = 15.21$ (0.000)	$\chi^2 = 5.27$ (P=0.072)
$H_{\perp}^* = 12.84$ (0.000)		$H_{\perp}^* = 7.96$ (0.005)	
$H_1^* = 11.62$ (0.001)	Phillips	$H_1^* = 10.85$ (0.001)	Phillips
$H_3^* = 18.26$ (0.000)	χ^2 (19)=15.64(0.000)	$H_3^* = 18.95$ (0.000)	χ^2 (19)=12.88(0.002)
$H_{13}^* = 12.64$ (0.000)	χ^2 (3)=7.57 (0.023)	$H_{13}^* = 11.81$ (0.001)	χ^2 (3)=5.22 (0.074)

Note: The Toda and Phillips (ECM) and Phillips (FM-VAR) tests were performed using a Gauss program kindly provided by H. Toda and H. Yamada, with the exception of the H_1^* (weak exogeneity) and H_3^* (long run exclusion) statistics, which were obtained using the Johansen approach within PcFiml (the corresponding Toda and Phillips statistics suggested in TP(1994) displayed very strange large results). The χ^2 (19) and χ^2 (3) statistics in the Phillips approach stand for the Phillips statistic computed using 19 lags (approximately equal to $T^{2/3}$) and 3 lags (approximately equal to $T^{1/4}$) respectively in the kernel estimator as explained in section 4. Figures between parentheses are the marginal significance levels (p-values).

In what regards the money-prices causality, all the tests: Toda and Phillips (ECM), Toda and Yamamoto (LA-VAR) and Phillips (FM-VAR) reject the null of Granger non-causality. In fact, for the LA-VAR and FM-VAR approaches the conclusion obtains immediately as the null of non-causality is rejected a 5% level (the p-value is always less than 0.05). For the ECM approach we note that the null of Granger non-causality is rejected according to the three alternative strategies

(P1), (P2) and (P3) suggested by Toda and Phillips. We note that, according to strategy (P3), the fact that we reject H_{\perp}^* (the null of short-run non-causality, in our case, boils down to test the coefficient of $\Delta^2 m_{t-1}$ in equation (A1.1)) allows one to immediately reject the general non-causality hypothesis. However, in our case, we also reject the long-run non-causality hypothesis as both H_1^* (weak-exogeneity) and H_3^* (long-run exclusion) are rejected. We thus conclude that money growth does in fact Granger cause inflation, “both in short as well as in the long run”, in the context of the VAR ($\Delta m, \Delta p$) .

As to the prices-money causality we also conclude for the existence of Granger causality from inflation to money growth. However, the evidence is not as strong as in the money-inflation case. In fact the ECM approach also clearly rejects the null of non-causality, but the LA-VAR approach does not reject the null of non-causality for a 5% test even though it rejects the null for a 10% test.

Evidence within the “extended” bivariate VAR ($\Delta m, \Delta p; \Delta y$) model

Within the simple VAR ($\Delta m, \Delta p$) framework it might be interesting to test whether the above conclusions still hold should we account for GDP growth. Under the assumption that the GDP growth rate is I(0) we do not expect this new model to change our conclusions on the Granger long-run causality tests obtained under the ECM approach. However, the conclusions about short run causality may change. Also under the LA-VAR approach the conclusions may change if it is the case that the GDP growth rate is of the same type as the other two variables (i.e., all the three variables are integrated of order one, which seems not a very reasonable assumption or all the three variables are stationary which some authors may find a reasonable hypothesis). The selected VAR model in the ECM format for the period 1980/4-2000/4 now reads:

$$\begin{aligned} \Delta^2 p_t = & -0.259\Delta^2 p_{t-1} - 0.253\Delta^2 m_{t-1} - 0.077\Delta y_{t-1} + 0.027\Delta y_{t-2} - 0.266[\Delta p_{t-1} - 0.909\Delta m_{t-1} + 0.006] \\ & (-2.58) \quad (-3.47) \quad (-1.72) \quad (0.61) \quad (-3.72) \quad (-6.84) \quad (2.08) \end{aligned} \quad (A1.3)$$

$$\begin{aligned} \Delta^2 m_t = & -0.346\Delta^2 p_{t-1} - 0.232\Delta^2 m_{t-1} - 0.123\Delta y_{t-1} + 0.146\Delta y_{t-2} + 0.352[\Delta p_{t-1} - 0.909\Delta m_{t-1} + 0.006] \\ & (-2.40) \quad (-2.22) \quad (-1.91) \quad (2.27) \quad (3.44) \quad (-6.84) \quad (2.08) \end{aligned} \quad (A1.4)$$

The first important point to note is that, as expected, the cointegration results still apply to the extended model¹⁸. The estimate for the long run coefficient even though somewhat smaller (it is now equal to 0.909) is still statistically not different from one (the p-value for this null hypothesis is 0.538). As to the GDP growth variable it appears not relevant in the inflation equation, but significant in the money growth equation. The remaining coefficients relevant for the Granger causality tests all remain significant (in the sense that the t-statistics are not significantly reduced vis-à-vis the simple $(\Delta m, \Delta p)$ bivariate model).

The results of the tests are displayed in Table A1.2. As expected the null of money-inflation non-causality is clearly rejected according to the Toda and Phillips ($P1$), ($P2$) and ($P3$) strategies as well as to the Toda and Yamamoto approach.

Table A1.2 – “Extended” Bivariate VAR $(\Delta m, \Delta p, \Delta y)$

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^* = 17.360$ (0.000)	$\chi^2 = 9.125$ (P=0.010)	$H^* = 15.196$ (0.001)	$\chi^2 = 5.238$ (P=0.073)
$H_{\perp}^* = 12.171$ (0.000)		$H_{\perp}^* = 7.021$ (0.008)	
$H_1^* = 12.234$ (0.001)	Phillips	$H_1^* = 12.143$ (0.001)	Phillips
$H_3^* = 19.113$ (0.000)		$H_3^* = 19.694$ (0.000)	
$H_{13}^* = 12.943$ (0.000)		$H_{13}^* = 12.828$ (0.001)	

Note: See table A1.1. We did not compute the Phillips tests as the econometric routine was not designed to deal with the Fully Modified estimator when the VAR model includes I(0) exogenous variables.

As to the inflation-money growth causality we still conclude for the existence of Granger causality according to the Toda and Phillips three strategies, as well as to the Toda and Yamamoto LA-VAR approach, even though, for this latter test with a marginal significance level slightly above 5%.

¹⁸ For the trace statistic we get LR1=26.93 (the 1% critical value is 24.6) and for the maximum eigenvalue test we get LR2=23.11 (the 1% critical value is 20.2).

Evidence within the bivariate VAR ($m-p, \Delta p$) model

As expected this simple model does not exhibit cointegration. The trace and maximum eigenvalue statistics are 9.54 and 9.33 respectively, which are not significant even for a 10% test¹⁹. The ECM approach is not applicable, but the LA-VAR rejects the null of Granger non-causality ($\chi^2=9.423(0.009)$). The evidence from the FM-VAR approach is mixed: the $\chi^2(19)=7.223(0.027)$ test rejects the null, but $\chi^2(3)=1.83(0.401)$ does not.

Evidence within the trivariate VAR ($m-p, \Delta p, y$) model

We note that this VAR obtains from the general VAR used in TV(2000) by dropping the interest rates. Using again a VAR with 2 lags we conclude for the existence of one cointegrating vector, which reads as (1, 5.334, -1.248)²⁰. The outcome of the tests is in Table A1.3.

Table A1.3 – VAR ($m-p, \Delta p, y$)

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^*=6.260(0.044)$	$\chi^2=8.442(0.015)$	$H^*=9.028(0.011)$	$\chi^2=0.972(0.615)$
$H_{\perp}^*=2.083(0.149)$		$H_{\perp}^*=0.536(0.464)$	
$H_1^*=1.487(0.223)$	Phillips	$H_1^*=6.830(0.009)$	Phillips
$H_3^*=10.565(0.001)$	$\chi^2(19)=3.473(0.176)$	$H_3^*=9.717(0.002)$	$\chi^2(19)=1.403(0.496)$
$H_{13}^*=2.021(0.155)$	$\chi^2(3)=1.225(0.542)$	$H_{13}^*=8.556(0.003)$	$\chi^2(3)=0.654(0.721)$

Note: See table A1.1

The evidence on money-prices causality is now mixing. If anything the introduction of GDP in the model has weakened the causality relation between the “real money stock” and inflation.

We recall that even though there is some correspondence between the “free” VAR (m, p) with long run homogeneity and the VAR ($m-p, \Delta p$) type models, the latter imposes some restrictions on the short-term dynamics, which as we have seen may have strong consequences for the Granger causality tests. According to the LA-VAR approach the null of Granger non-causality is

¹⁹ These results are for a VAR with 2 lags, but the conclusions do not change for a VAR with 3 lags.

²⁰ For the trace statistics we get LR1=36.14 and for the maximum eigenvalue LR2=24.64. Both are significant at 5%.

strongly rejected (p-value of 0.015). The same conclusion is obtained from the ECM approach for the strategy (P2). However, according to strategies (P1) and (P3) the null of Granger non-causality cannot be rejected for standard significance levels. According to the Phillips FM-VAR approach the null of non-causality is also not rejected.

For the prices to money causality tests the conclusion is basically the same. The null of non-causality is rejected according to the ECM approach (for all the three testing strategies). We note however that in this case only “long run causality” is significant. On the other hand, the null of non-causality is not rejected according to both the LA-VAR and FM-VAR approaches.

Evidence within the general $(m-p, \Delta p, y, s, l)$ model

As in TV(2000) there is evidence of three cointegrating vectors in this general VAR model. In this case testing for causality becomes somewhat cumbersome because to compute the H_1^* and H_3^* statistics within the Johansen approach requires previous identification of the three cointegrating vectors. Furthermore the identification must be such that it is not lost under the null hypothesis of long run non-causality.

In order to compute the tests H_3^* and H_1^* , in testing money-prices causality, we assumed the following exactly identified cointegrating vectors: $\beta_1 = (\beta_{11}, \beta_{12}, 1, 0, 0)$, $\beta_2 = (\beta_{21}, \beta_{22}, 0, 1, 0)$ and $\beta_3 = (\beta_{31}, \beta_{32}, 0, 0, 1)$, so that we have $\beta_{11} = \beta_{21} = \beta_{31} = 0$ as the null under H_3^* .

Looking at table A1.4 the conclusion is that the evidence in favour of Granger causality has weakened even further vis-à-vis the simpler VAR($m-p, \Delta p, y$) model. Now the null of Granger non-causality is still rejected under the LA-VAR approach. But, under the ECM-VAR approach this null is not rejected for any of the three alternative testing strategies (P1), (P2) and (P3). For the FM-VAR strategy the evidence is now mixing: the null of Granger causality being rejected by the χ^2 (19) but not by the χ^2 (3) statistic.

In order to compute the tests H_3^* and H_1^* , in testing for prices-money causality, we assumed the following exactly identified cointegrating vectors: $\beta_1 = (\beta_{11}, \beta_{12}, 1, 0, 0)$, $\beta_2 = (0, \beta_{22}, 0, 1, \beta_{25})$ and $\beta_3 = (\beta_{31}, \beta_{32}, 0, 0, 1)$, so that we have $\beta_{12} = \beta_{22} = \beta_{32} = 0$ as the null under H_3^* . In this case the evidence against Granger causality is overwhelming: for none of the three testing strategies is the null of Granger causality rejected.

Table A1.4 – VAR ($m-p, \Delta p, y, l, s$)

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^* = 1.179$ (0.555)	$\chi^2 = 5.37$ (0.068)	$H^* = 0.119$ (0.942)	$\chi^2 = 0.094$ (0.954)
$H_{\perp}^* = 0.170$ (0.680)		$H_{\perp}^* = 0.119$ (0.730)	
$H_1^* = 15.36$ (0.002)	Phillips	$H_1^* = 10.79$ (0.013)	Phillips
$H_3^* = 20.43$ (0.000)	χ^2 (19) = 9.436 (0.009)	$H_3^* = 14.31$ (0.003)	χ^2 (19) = 2.603 (0.272)
$H_{13}^* = 0.969$ (0.325)	χ^2 (3) = 2.063 (0.357)	$H_{13}^* = 0.037$ (0.848)	χ^2 (3) = 0.101 (0.951)

Note: See table A1.1.

For this large VAR model, we note that our conclusions on money-prices causality tests completely agree with the ones drawn in TV(2000).

Evidence from VAR (m, p, y, l, s) type models

Table A1.5 shows the results of the LA-VAR approach for the VAR (m, p, y, l, s) type models²¹. Even though the table considers the two cases of I(2) and I(1) variables, we attach more importance to the I(2) case, as the hypothesis of p_t and m_t being I(1) for the sample period under analysis does not seem very realistic.

In what regards money-prices causality, we conclude from Table A1.5 that the null of Granger non-causality is rejected for the models VAR(m, p) and VAR(m, p, y), but not for the model with interest rates VAR(m, p, y, l, s).

As to prices-money causality tests, the general conclusion is that the null of Granger non-causality cannot be rejected for any of the three estimated models.

²¹ We recall that neither the Toda and Phillips nor the Phillips approaches are applicable for models with I(2) variables.

Table A1.5 - VAR (m, p, y, l, s) type models

Money-prices causality		Prices-money causality	
Toda and Yamamoto LA-VAR approach			
I(2) model	I(1) model	I(2) model	I(1) model
VAR (m, p)			
χ^2 =6.353 (0.042)	χ^2 =6.84 (0.033)	χ^2 =3.729 (0.155)	χ^2 =3.719 (0.156)
VAR (m, p, y)			
χ^2 =6.093 (0.048)	χ^2 =6.352 (0.042)	χ^2 =3.365 (0.186)	χ^2 =3.451 (0.178)
VAR (m, p, y, l, s)			
χ^2 =1.488 (0.475)	χ^2 =2.876 (0.237)	χ^2 =2.183 (0.336)	χ^2 =1.884 (0.390)

Note: all the models in the table were estimated for the period 1980/4-2000/4 with two lags (plus two lags for the I(2) models and plus one lag for the I(1) models).

Appendix 2: robustness analysis with results for the Granger causality tests for the period 1985/1-2000/4²²

Table A2.1 – VAR ($\Delta m, \Delta p$)

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^* = 10.04$ (0.007)	$\chi^2 = 4.65$ (0.098)	$H^* = 8.71$ (0.013)	$\chi^2 = 4.92$ (P=0.085)
$H_{\perp}^* = 5.55$ (0.019)		$H_{\perp}^* = 6.17$ (0.013)	
$H_1^* = 8.24$ (0.004)	Phillips	$H_1^* = 5.38$ (0.020)	Phillips
$H_3^* = 10.76$ (0.001)	χ^2 (19)=4.72 (0.095)	$H_3^* = 12.20$ (0.001)	χ^2 (19)=7.49 (0.024)
$H_{13}^* = 8.29$ (0.004)	χ^2 (3)=3.68 (0.159)	$H_{13}^* = 5.72$ (0.017)	χ^2 (3)=4.00 (0.135)

²² The usable sample period is, actually, 1985/4-2000/4.

Table A2.2 – “Extended” Bivariate VAR ($\Delta m, \Delta p, \Delta y$)

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^* = 10.397$ (0.006)	$\chi^2 = 4.556$ (P=0.102)	$H^* = 9.024$ (0.011)	$\chi^2 = 5.199$ (P=0.074)
$H_{\perp}^* = 5.036$ (0.025)		$H_{\perp}^* = 6.078$ (0.014)	
$H_1^* = 8.854$ (0.003)	Phillips	$H_1^* = 5.682$ (0.017)	Phillips
$H_3^* = 10.204$ (0.001)		$H_3^* = 12.122$ (0.001)	
$H_{13}^* = 9.037$ (0.003)		$H_{13}^* = 6.326$ (0.012)	

Table A2.3 – VAR ($m - p, \Delta p, y$)

Money-prices causality tests		Prices-money causality tests	
Toda and Phillips	Toda and Yamamoto	Toda and Phillips	Toda and Yamamoto
$H^* = 6.783$ (0.034)	$\chi^2 = 4.398$ (0.111)	$H^* = 5.713$ (0.057)	$\chi^2 = 1.068$ (0.586)
$H_{\perp}^* = 1.048$ (0.306)		$H_{\perp}^* = 0.401$ (0.527)	
$H_1^* = 3.019$ (0.082)	Phillips	$H_1^* = 4.670$ (0.031)	Phillips
$H_3^* = 10.956$ (0.001)	χ^2 (19)=3.674(0.159)	$H_3^* = 8.961$ (0.003)	χ^2 (19)=0.715(0.699)
$H_{13}^* = 3.902$ (0.048)	χ^2 (3)=3.053 (0.217)	$H_{13}^* = 5.560$ (0.018)	χ^2 (3)=0.199 (0.905)

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