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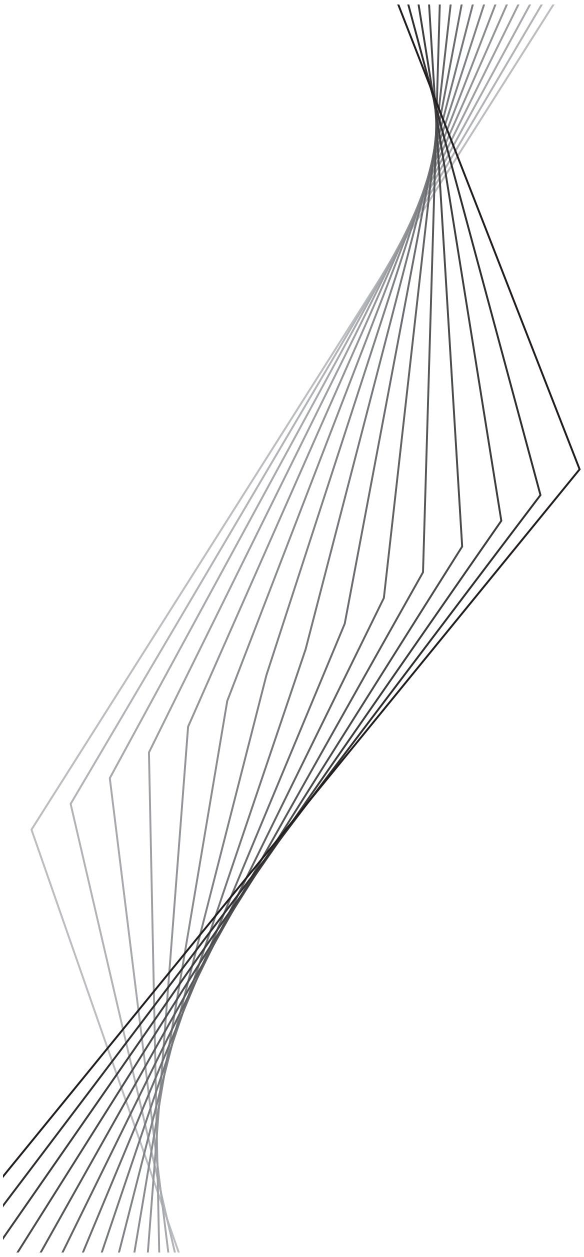

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* *The authors are grateful to the many European Central Bank colleagues who have made helpful comments and suggestions. The authors also thank D. Gerdesmeier and P. Vlaar for insightful discussions. In addition, the paper has greatly benefited from comments from two anonymous referees and from staff in the research departments of a number of national central banks. Views expressed represent exclusively the opinion of the authors and do not necessarily reflect those of the European Central Bank. Any remaining errors are of course the sole responsibility of the authors.*

Address	**Kaiserstrasse 29**
	D-60311 Frankfurt am Main
	GERMANY
Postal address	**Postfach 16 03 19**
	D-60066 Frankfurt am Main
	Germany
Telephone	**+49 69 1344 0**
Internet	**http://www.ecb.int**
Fax	**+49 69 1344 6000**
Telex	**411 144 ecb d**

ISSN 1561-0810

Abstract

In this paper, an empirically stable money demand model for M3 in the euro area is constructed. Starting with a multivariate system, three cointegrating relationships with economic content are found: (i) the spread between the long- and the short-term nominal interest rates, (ii) the long-term real interest rate, and (iii) a long-run demand for broad money M3. There is evidence that the determinants of M3 money demand are weakly exogenous with respect to the long-run parameters. Hence, following a general-to-specific modelling approach, a parsimonious conditional error-correction model for M3 money demand is derived which can be interpreted economically. For the conditional model, long- and short-run parameter stability is extensively tested and not rejected. Insights into the dynamics of money demand are gained by means of SVAR techniques exploring the impulse response functions of the cointegrated multivariate system.

JEL Classification System: C22, C32, E41

Keywords: Money demand, euro area, cointegration, error-correction model, impulse response analysis

I Introduction

In October and December 1998, the Governing Council of the European Central Bank announced the key elements of the monetary policy strategy of the Eurosystem. These comprise a quantitative definition of the primary objective, namely price stability, and the "two pillars" used for achieving this objective: a prominent role for money, as signalled by the publication of a reference value for the growth rate of broad money M3, and a broadly based assessment of the outlook for, and risks to, price stability in the euro area.[1]

Whilst the role assigned to money in the strategy is primarily based on theoretical grounds (namely that inflation is ultimately a monetary phenomenon), empirical evidence, together with conceptual considerations, may indeed play an important role in selecting the particular monetary aggregate that best serves the purposes at hand. The existence of a stable and predictable relationship between the demand for a given monetary aggregate and its macroeconomic determinants has traditionally been considered a key element in this respect. Further considerations refer to leading indicator and controllability properties.[2]

Against this background, this paper presents some results of recent empirical research carried out at the European Central Bank on the demand for broad money M3 in the euro area. The paper is organised as follows. Section 2 briefly discusses our benchmark long-run specification for the estimation of the demand for broad money M3 in the euro area and the data underlying the empirical analysis. The data set is reproduced in Annex 1. Section 3 investigates the cointegration properties of the data by means of the application of the Johansen procedure to a set of variables consisting of real holdings of M3 (m-p), real GDP (y), short- (s) and long-term (l) interest rates and the inflation rate as measured by the annualised quarterly changes in the log of the GDP deflator ($\pi/4= \Delta p$). In the light of the results, Section 4 develops a conditional model for M3 money demand in the euro area. We proceed in two steps and follow a general-to-specific modelling approach. Firstly, an unrestricted autoregressive distributed lag (ADL) model in m-p, y, s, l and π is estimated and its long-run solution computed. And secondly, the results obtained in the first step are then used for deriving a parsimonious, economically interpretable, conditional error-correction model for $\Delta(m$-$p)$.[3]

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1 See ECB (1999a). Angeloni et al. (1999) discusses the analytical foundations of the monetary policy strategy of the ECB.

2 See ECB (1999b).

3 Computations in Sections 3 and 4 were made using PcGive and CATS.

Some insight into the dynamics of money demand can be gained by looking at the lag weights of the dynamic single-equation model estimated in Section 4. It is noted, however, that the simulation experiment involves a rather unrealistic assumption, namely that the variables on the r.h.s. of the money demand equation are orthogonal. Against this background, Section 5 makes an attempt to overcome this drawback by conducting a more realistic impulse response analysis within a multivariate framework which allows for the interplay of all variables within the system.[4] This section heavily builds on results in Vlaar (1998) and Vlaar and Schuberth (1999). A brief review of the methodology is given in Annex 2. Finally, Section 6 draws the main conclusions from the analysis.

2 The economic model and the data

2.1 The model

Whilst money is held for a number of purposes[5], most theories of money demand lead, as argued in Ericsson (1999), to a long-run specification of the form:

$$M^d / P = g(Y, \tilde{R}) \qquad (1.a)$$

where M^d , P, Y and $\tilde{R}$ stand for nominal money, the price level, a scale variable and a vector of returns on various assets. In applied work, a (semi-) log-linear form is often found to be an acceptable empirical approximation to equation (1), namely:

$$m_t^d - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2 R_t^{own} + \gamma_3 R_t^{out} + \gamma_4 \pi_t \qquad (1.b)$$

where variables in lower case indicate logs, $\pi/4=\Delta p$, and R^{own} and R^{out} stand, respectively, for the nominal rates of return on financial assets included in and excluded from the definition of the monetary aggregate.

In (1.b) above, γ_1 measures the long-run elasticity of money demand with respect to the scale variable, whilst γ_2, γ_3 and γ_4 are, in turn, the long-run semi-elasticities with respect to the own and alternative rates of money and with respect to the inflation rate. Expected signs for the parameters in (1.b) are $\gamma_1>0$, $\gamma_2>0$, $\gamma_3<0$, $\gamma_4<0$ and, possibly, $\gamma_2=-\gamma_3$. In the latter case, long-run money demand can be expressed as a function of the spread R^{out}-R^{own}, which in turn is interpretable as the opportunity cost of holding money.

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4 The impulse response analysis was conducted using Malcolm and MATLAB.

5 Traditionally, a number of distinct motives for holding money are pointed out in the literature, giving rise to a transactions demand, a precautionary demand and a speculative demand for money. See Goldfeld and Sichel (1990) and Laidler (1993).

Long-run price homogeneity of money demand has been assumed in (1.b), as predicted by most theories, but this can be empirically tested. Some theories also predict particular values for γ_1. For instance, γ_1=0.5 in the Baumol-Tobin model or γ_1=1.0 under some formulations of the quantity theory of money. Values γ_1>1.0 are also found with some frequency in the empirical literature for broad definitions of money, which in turn is customarily interpreted as proxying omitted wealth effects in equation (1.b). Extension of equation (1.b) to include wealth can be justified under a standard portfolio approach to asset demand theory. This is not pursued here, however, due to the lack of reliable wealth data for the euro area.

The inclusion of the inflation rate in (1.b) is the subject of some ongoing controversy in the literature. However, separate consideration of the inflation rate in dynamic models of money demand may be of particular interest for a number of reasons. Firstly, it permits a reparameterisation of the models in terms of real money holdings and the inflation rate. Such reparameterisation allows for the theoretically plausible hypothesis of long-run price homogeneity of money demand but does not impose any untested (and frequently empirically rejected) common factor restriction of short-run price homogeneity. In the context of cointegrated systems, it may also permit some convenient simplifications when the money stock and the price level are found to be CI(2,1), i.e. *m* and *p* are I(2) but *m-p* is I(1), such that the I(2) system can be mapped into an I(1) system.[6]

Secondly, numerous authors have forcefully argued for the inclusion of the inflation rate as an important determinant of constant-parameter empirical models of money demand.[7] This is customarily justified on the basis that it represents the opportunity cost of holding money rather than real assets.[8] On different grounds, within a cost-minimising framework similar to that in Hendry and von Ungern-Sternberg (1981), Wolters and Lütkepohl (1997) show that, in the presence of a short-run nominal adjustment mechanism and under inflation persistence, the inflation rate may enter the long-run relation even if it does not appear in the desired long-run demand for money function.

And thirdly, it could be argued that the inclusion or exclusion of inflation in models of real money demand is an issue of dynamic specification to be settled at the empirical level. In this sense, whilst some ambiguity would necessarily remain on the interpretation of the role of the inflation rate, the consideration of inflation as one of the variables entering the long-

6 See, for instance, Johansen (1992).

7 For instance, seven out of the fourteen articles included in the book *Money Demand in Europe* recently edited by H. Lütkepohl and J. Wolters include inflation as a determinant of the long-run demand for money.

8 See, for instance, Ericsson (1999).

run demand for money or, alternatively, affecting only the process of dynamic adjustment to the long-run equilibrium would have little empirical content, since both interpretations lead to observationally equivalent empirical models.[9]

2.2 The data

The long-run specification given by equation (1.b) is our maintained hypothesis for estimation of the demand for broad money M3 in the euro area. Following earlier work carried out at the European Monetary Institute in the context of preparatory work for Stage Three[10], the following empirical counterparts proxying the variables in the r.h.s. of (1.b) were chosen: real GDP (y) for the scale variable, the GDP deflator (p) for the price level, the short-term money market rate (s) for the return on assets included in the definition of M3, and the long-term bond yield (l) for the return on assets excluded from the monetary aggregate. The choice of real GDP and the GDP deflator as the scale and price variables in the money demand function is standard in existing empirical work, though alternative measures such as total final expenditure, consumption or wealth are also frequently found. The choices of the short- and long-term interest rates could be justified on the basis of the broadness of the M3 aggregate, which includes assets that are remunerated at or close to market rates, though alternative measures of the own rate deserve being explored as longer area-wide time series on deposit rates become available.

The data on y, p, s and l are taken from the ECB area-wide model database and their construction can be briefly described as follows. Euro area real GDP and the GDP deflator are seasonally adjusted and obtained from EUROSTAT for the period in which these time series are available; for earlier periods, they are calculated from different national sources and are then aggregated using fixed weights based on 1995 GDP at PPP rates. The short- and long-term interest rates are obtained from the BIS database as weighted averages of national rates using also fixed weights based on 1995 GDP at PPP rates.

As defined in December 1998 by the Governing Council of the ECB, M3 consists of holdings by euro area residents of currency in circulation plus certain liabilities issued by Monetary Financial Institutions (MFIs) and, in the case of deposits, liabilities issued by some institutions which are part of the central government (such as Post Offices and Treasuries). These include: overnight deposits, deposits with agreed maturity up to 2 years,

9 See Goldfeld and Sichel (1987).
10 See Fagan and Henry (1999).

deposits redeemable at notice up to three months, repurchase agreements, money market fund shares, money market paper, and debt securities with maturity up to two years.[11]

As from January 1999, the ECB publishes regularly data on euro area monetary aggregates (M1, M2 and M3) denominated in euro. The monetary aggregates are compiled on the basis of the consolidated balance sheet of the MFI sector from data collected under the new harmonised system of money and banking statistics within the framework of ECB Regulation of 1 December 1998 (ECB/1998/16). This consolidated balance sheet in euro is, in turn, available with the same degree of detail only back to September 1997. For periods prior to September 1997, the ECB is not in a position to produce historical data on euro area monetary aggregates according to its regular compilation procedures and, therefore, longer time series can only be constructed on the basis of the aggregation of estimated national contributions to the euro area aggregates compiled from a number of not fully harmonised national statistical sources[12], including — as far as information is available — cross-border positions of MFIs within the euro area. In its February 1999 Monthly Bulletin the ECB released some historical estimates back to 1980. The time series were expressed in euro, with the estimated historical national contributions aggregated by their conversion into the single currency using the irrevocable conversion rates fixed on 31 December 1998.

It is fair to say that there is no uncontroversial aggregation method for linking euro area pre- and post-1999 data, reflecting the fundamental problem that it is only from 1999 onwards that a single currency is in place.[13] The existing empirical literature on area-wide money demand has indeed addressed this issue and a number of methods have been employed in this respect. Thus, in Monticelli and Strauss-Kahn (1992) area-wide aggregates are computed in a single currency using current exchange rates vis-à-vis the ECU. In Fase and Winder (1999) fixed exchange rates are, instead, employed. Fagan and Henry (1999) propose to weight national aggregates according to a fixed weighting scheme based on GDP at PPP. Fase and Winder (1997) discussed the effects of alternative aggregation methods. The use of a consistent aggregation method for monetary aggregates, on the one hand, and for the r.h.s. variables in (1.b), on the other, is often stressed in the literature as an important requirement to bear in mind when developing empirical models of money demand in the euro area.

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[11] A detailed description of euro area monetary aggregates can be found in ECB (1999b).

[12] National contributions' to euro area monetary aggregates need to be distinguished from old national monetary aggregates. In particular, the euro area definition of M3 does not coincide (in terms of asset coverage and sector/instrument/maturity classification) with national definitions of broad money in place during Stage Two. These national contributions have been estimated on the basis of national data following a set of across-countries common specifications compatible with the new system of money and banking statistics. A description of the construction of historical estimates can be found in the statistical annex to ECB (1999b).

[13] See Winder (1997) for a discussion on aggregation methods.

Against the background of the choice of r.h.s. variables explained above, we proceed as follows throughout the paper. From September 1997 onwards, M3 as regularly published by the ECB is always employed. For the period prior to September 1997, the fixed-weights method proposed in Fagan and Henry (1999) is used to aggregate the historical national contributions and to produce back-estimates of area-wide M3 starting from September 1997 levels. This method of aggregation ensures full compliance with the conceptual consistency requirement highlighted above and, therefore, it is used throughout the paper, bar Section 4.3.[14] This way of proceeding, however, has one drawback, namely that it departs from the compilation procedures for M3 based on the use of fixed conversion rates which are in place since the start of Stage Three. Since this may potentially have some implications if the estimated models were to be used out of the estimation sample, we explore in Section 4.3 the effects that the aggregation of the estimated national contributions using fixed conversion rates (as published in ECB 1999b) would have on our estimates when no change is introduced in the way in which the r.h.s. variables in (1.b) are computed.

Finally, it needs to be born in mind that the use of euro area historical data for monetary aggregates compiled using a set of across-countries common specifications as compatible as possible with the sector/instrument/maturity classification laid down in the new system of euro area money and banking statistics, marks a significant departure from the data used in previous empirical research on area-wide money demand. Regardless of the aggregation issue, the new data set minimises the risks that the move to the new system of statistics introduced a sizeable break in the estimates, rendering empirical evidence useless in the new context. On the other hand, it makes it difficult to compare the results with previous analyses and, accordingly, this avenue is not pursued in this paper.

Figures 1.a and 1.b plot the time series used in the empirical analysis:[15] real holdings of M3 and real GDP (Figure 1.a), the short- and long-term interest rates and the inflation rate measured by the annualised quarterly changes in the GDP deflator (Figure 1.b). From visual inspection, the trending behaviour which often characterises non-stationary series is apparent in the plots of m-p and y. Furthermore, s, l and π appear to share a common, possibly stochastic, nominal trend during the sample under investigation. The latter

[14] Alternatively, historical series for euro area nominal and real GDP at fixed conversion rates could be computed and the corresponding implicit GDP deflator be calculated. Whilst ensuring consistent aggregation, this procedure cannot be applied to interest rates. Moreover, area-wide GDP inflation so defined would, in general, be different from the weighted average of countries' developments, which is the concept most used in economic analysis.

[15] The lack of reliable non-seasonally adjusted series for area-wide real GDP and the GDP deflator limits our choice between seasonally adjusted and unadjusted data. Thus, quarterly averages of seasonally adjusted M3 monthly data are used in the analysis. Seasonal adjustment has been made using SEATS.

observation is more visible in Figures 1.c, 1.d and 1.e, which depict time series for the real long-term interest rate (Figure 1.c), the real short-term rate (Figure 1.d) and the spread between the long- and the short-term rates (Figure 1.e). All of them look much more stationary than their individual counterparts in Figure 1.b and the application of the Johansen procedure below will confirm that they indeed constitute simple cointegrating relationships within our set of variables. Finally, Figure 1.f plots the income velocity of M3, showing the downward trending behaviour which has been reported elsewhere. Income velocity of M3 has declined by a cumulated 13% since the early 80s, representing a -0.7% annual decline per year. It has tended to stabilise, however, in the most recent years. This pattern in velocity parallels to some extent the developments in area-wide inflation shown in Figure 1.b.

ADF tests for unit roots support the view that m-p, y, s, and l are I(1) for the sample under investigation.[16] The same outcome applies to income velocity y+p-m. Less clear-cut results are obtained, however, for m and p. On the one hand, ADF tests tend to reject a second unit root for the full sample under consideration when the alternative contains a deterministic trend, i.e. Δm and Δp could be trend-stationary. On the other hand, recursive computation of the corresponding ADF tests indicates that this finding is not robust to the choice of sample and provides support for assuming that both variables can be better described as I(2). Additionally, since univariate tests are known to have low power against some stationary alternatives, multivariate tests for starionarity will be conducted as well within the application of the Johansen procedure in Section 3. The results therein provide further empirical support for the hypothesis that the inflation rate is not stationary during the period analysed.

3 The cointegration analysis

In this section the cointegration properties of the data are investigated by means of the application of the Johansen[17] procedure to the set of variables $z=(m\text{-}p, y, s, l, \pi)'$ consisting of real holdings of M3 (m-p), real GDP (y), short- (s) and long-term (l) interest rates and the inflation rate as measured by the annualised quarterly changes in the log of the GDP deflator (π). An unrestricted constant, allowing for a linear trend in the variables but not in the cointegrating relationships and a dummy variable (*DUM86*)[18] are also included in the

16 Detailed results are available from the authors upon request.

17 See Johansen (1995).

18 *DUM86* equals: 0.5 in the first, third and fourth quarters of 1986, 1 in the second quarter of 1986, and 0 elsewhere. This dummy variable is needed for taking account of special developments in German data around the period in which debt securities were subjected to reserve requirements, leading to substantial differences in the annual growth rates of German broad money with and without debt securities.

system. The data are quarterly and the estimation sample spans the period 1980:Q4 to 1997:Q2. The monetary aggregate, the GDP deflator and real GDP are seasonally adjusted. The interest rates are measured as percent per annum, expressed as fractions.

The results from the application of the Johansen procedure are summarised in Table 1. The top panel of the table reports the Schwarz (SC) and Hannan-Quinn (HQ) information criteria for the selection of lag-length (k) and various diagnostic statistics on the system residuals: Lagrange-multiplier tests for first- (LM(1)), fourth- (LM(4)) and up to eighth-order (LM(1-8)) residual autocorrelation, Doornik and Hansen (1994) multivariate test for normality (NORM), and a White-type test for heteroskedasticity (HET).

According to the test statistics above, the VAR with k=2 appears reasonably well specified over the estimation sample, although some residual non-normality is revealed by the Doornik and Hansen test. Excess kurtosis in the residuals of the equations for the short-term rate and, to a lesser extent, for real income, due to the presence of some outliers at the beginning of the estimation period, appears primarily responsible for this finding.

As regards cointegration, Table 1 shows the results for the trace test, since this is considered to be robust to the non-normality encountered in the data.[19] Furthermore, as critical values are affected by the inclusion of dummy variables, the rank test statistics reported in Table 1 refer to the system both with and without *DUM86* (the latter in square brackets). After small sample corrections are made as suggested in Cheung and Lai (1993), the results point to the existence of three cointegrating vectors at the 90% confidence level. Conditional on the choice of cointegration rank r=3, tests for long-run exclusion, stationarity and weak exogeneity of each variable as well as tests for structural hypothesis on β are then reported in the table. The tests conducted do not allow to reject at standard confidence levels the stationarity of both the spread between the long- and the short-term interest rate (H_0^1), consistent with the term structure of interest rates, and the real long-term interest rate (H_0^2), consistent with the Fisher parity. The stationarity of the real short-term interest rate is in turn implied by the non-rejection of the joint hypothesis $H_0^1 \cap H_0^2$. On the contrary, long-run homogeneity of real money and real income is rejected at standard confidence levels, both when the hypothesis is tested in isolation and when it is tested jointly with $H_0^1 \cap H_0^2$. The tests yield $\chi_3^2 = 9.90\,[.020]$ and $\chi_5^2 = 10.38\,[.065]$, respectively.

[19] See Cheung and Lai (1993).

The estimated cointegrating vectors are reported in Table 1 in the form of irreducible cointegrating relations (IC), following Davidson (1998).[20] Besides the two over-identified relationships, namely the spread *(l-s)* [β_2'=(0, 0, -1, 1, 0)] and the real long-term rate *(l-π)* [β_3'=(0, 0, 0, 1, -1)], an additional just-identified cointegrating vector [β_1'=(1, -δ, -η, 0, 0)] expressing real holdings of M3 as a function of the real and nominal stochastic trends driving the system is reported in the table:

$$\underset{}{(m-p)_t} = \underset{(.034)}{1.17}\ y_t - \underset{(.17)}{1.26}\ s_t + \hat{v}_t \tag{2}$$

Just-identification of β_1 has been obtained by setting arbitrarily to zero the coefficients of the long-term interest rate and the inflation rate. It should be noticed that, whilst this normalisation has been chosen as to make (2) look like a traditional textbook money demand function, nothing so far guarantees its structural interpretation.[21] In particular, any linear combination $\beta = \sigma_1\beta_1 + \sigma_2\beta_2 + \sigma_3\beta_3$ is also a cointegrating relationship and, for some parameter values, would be a plausible candidate for constituting our relationship of interest. Normalising in real money holdings (σ_1=1), $\beta' z_t$ can be written as follows:

$$(m-p)_t = k + \delta y_t + \eta s_t - \sigma_2(l_t - s_t) - \sigma_3(l_t - \pi_t) \tag{3.a}$$
$$= \gamma_0 + \gamma_1 y_t + \gamma_2 s_t + \gamma_3 l_t + \gamma_4 \pi_t \tag{3.b}$$

where economic theory suggests that money demand would correspond to equation (3.b) with: γ_1>0; γ_2>0 (if the short-term rate proxies the own rate of money), γ_3<0 (if the long-term rate measures the return of financial assets alternative to those included in the monetary aggregate) and, possibly, $\gamma_2 = -\gamma_3$ (if the spread is the relevant opportunity cost of holding money relative to financial assets not included in the definition of money), and γ_4<0 (if money is a substitute for real assets). Equation (3.b) corresponds to our benchmark long-run specification for the estimation of the demand for broad money M3 in the euro area and the parameters γ in (3.b) constitute our long-run parameters of interest.

It is also straightforward to show that the parameters in (3.a) and (3.b) are related as follows:

$$\gamma_1 = \delta \tag{3.c}$$
$$\gamma_2 = \eta + \sigma_2 \tag{3.d}$$
$$\gamma_3 = -(\sigma_2 + \sigma_3) \tag{3.e}$$
$$\gamma_4 = \sigma_3 \tag{3.f}$$

[20] A set of I(1) variables is called irreducibly cointegrated (IC) if they are cointegrated, but dropping any of the variables leaves a set that is not cointegrated.

[21] As opposed to a solved form, following the terminology in Davidson (1998).

From (3.c) to (3.f) some results of interest concerning the parameters δ and η in the first cointegrating vector β_1 follow. Firstly, the parameter δ identifies unambiguously the long-run income elasticity of money demand. Secondly, η identifies unambiguously the effect that the common nominal trend has on M3 real balances, i.e. $\eta = \gamma_2 + \gamma_3 + \gamma_4$. It does not identify the parameter measuring the semi-elasticity of money demand with respect to the short-term rate (γ_2) unless $\gamma_3 + \gamma_4 = 0$, which runs against economic intuition since both γ_3 and γ_4 are expected to be of negative sign. Thirdly, if the two interest rates entered the long- run money demand as a spread ($\gamma_2 = -\gamma_3$), η would correspond to the long-run semi-elasticity of money demand with respect to inflation ($\eta = \gamma_4$). And finally, if the inflation rate does not enter the long-run demand for money function ($\gamma_4 = 0$), the results reported in Table 1 (i.e. the significance of η) would rule out the spread formulation for long-run money demand.

Interestingly enough, weak exogeneity of real income, the short- and long-term interest rates and the inflation rate when the parameters of interest are those of the first cointegrating vector is not rejected by the tests reported in Table 1 (H_0^3).[22] One implication of this result is that, given the structure of the estimated cointegration space, weak exogeneity is not rejected either for the parameters of any linear combination involving β_1, on the one hand, and the two other over-identified cointegrating vectors β_2 and β_3, on the other. This holds in particular for the parameters γ of the long-run demand for money function which, as derived above, can be expressed as such a linear combination. That in turn indicates that, as far as the parameters of the long-run money demand are concerned, nothing can be learnt from the equations in the system other than the equation for real balances. Accordingly, efficient inference on the parameters of long-run money demand will be made in the next section on the basis of a parsimonious, conditional, single-equation model for area-wide money demand.[23]

4 A dynamic model of money demand in the euro area

In view of the results on cointegration and weak exogeneity obtained from the application of the Johansen procedure in Section 3, a conditional model for M3 money demand in the euro area is developed in this section. We proceed in two steps and follow a general-to-specific

[22] Estimation of δ and η in β_1 under $H_0^1 \cap H_0^2 \cap H_0^3$ provides the following results (standard errors in parentheses): δ=1.136 (.027) and η=-1.458 (.120)

[23] The results from the tests for weak exogeneity, however, need to be interpreted with caution since the 'weak exogeneity' status of a given variable may change when the system is augmented with additional variables. By contrast, cointegration is a property which is invariant to expansions of the system.

modelling approach.[24] Firstly, given the choice of lag length in the VAR, a second-order unrestricted autoregressive distributed lag (ADL) model in *m-p*, *y*, *s*, *l* and π is estimated and its long-run solution computed. And secondly, the results obtained in the first step are then used for deriving a parsimonious, economically interpretable, conditional error-correction model for $\Delta(m\text{-}p)$.

The estimation period spans the sample 1980:Q4 to 1997:Q2, leaving six quarters of 'fresh data' (1997:Q3-1998:Q4) for analysis of out-of-sample forecasting performance and stability. Since it has been argued that the introduction of the euro marks a significant regime shift which could affect empirical relationships estimated on the basis of past data[25], analysis of money demand stability during the period immediately preceding the introduction of the single monetary policy should provide evidence on the empirical relevance of such risks by detecting any anticipated effect.

4.1 Long-run money demand

The long-run solution to the estimated ADL(2) model is given by:

$$(m-p)_t = \underset{(.058)}{1.140}\, y_t - \underset{(.352)}{0.820}\,(l_t - s_t) - \underset{(.321)}{1.462}\,\pi_t + \hat{u}_t \qquad (4)$$

where the spread restriction turned out to be statistically acceptable [F(1,51)=.385 (.37)], whilst the exclusion of the inflation rate and the unit income elasticity were rejected [F(1,52)=20.73 (.00) and F(1,52)=5.86 (.019), respectively].

Equation (4) closely resembles our benchmark long-run specification and, following Boswijk (1995)[26], it is a natural candidate for structural interpretation as a money demand cointegrating relation. It shows a simple specification for the long-run demand for M3 whereby holdings of M3 real balances are determined by a measure of the volume of transactions (as proxied by real income), the opportunity cost of holding money relative to financial assets not included in the definition of the monetary aggregate (as proxied by the spread between the long- and short-term interest rates) and the inflation rate (proxied by changes in the GDP deflator).

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24 See Hendry (1995) and the references therein.
25 See, for instance, ECB (1999a).
26 See also Bårdsen and Fisher (1995). A careful discussion of Boswijk (1995a) can be found in Ericsson (1995) and a reply in Boswijk (1995b).

The signs and magnitudes of the estimated long-run coefficients also appear quite plausible on theoretical grounds. The long-run income elasticity is estimated significantly above one (1.14), a finding which is often interpreted as proxying omitted wealth effects in the demand for money function.[27] The estimated long-run semi-elasticities with respect to the spread and the inflation rate are, respectively, -0.82 and -1.46 (though in both cases less precisely estimated than the income coefficient).[28] Equation 4 provides an explanation for the downward trending behaviour of income velocity of M3 depicted in Figure 1.f in terms of two main factors: firstly, the income elasiticity of money demand higher than one; and secondly, the significant fall in the inflation rate witnessed by the euro area during the sample period. The spread between the long- and the short-term interest rates, which shows no particular trending behaviour when the entire sample is considered, does not appear to contribute significantly to this long-term pattern of velocity. However, the low speed of adjustment of this variable to its estimated long-run equilibrium indicates that its contribution cannot be disregarded at medium-term horizons.

The income elasticity and the inflation semi-elasticity do not differ significantly from the estimates obtained within the system approach in Section 3, providing further evidence on the validity of weak exogeneity when the parameters of interest are the long-run parameters of money demand. In this latter respect, formal statistical tests do not reject at standard confidence levels that the long-run relationship given by (4) lies in the cointegration space estimated within the system approach used in Section 3. The tests yield χ_2^2 =.43 [.43] (if the cointegration space is estimated unrestrictedly), χ_6^2 =5.60 [.47] (if the cointegration space is estimated under $H_0^1 \cap H_0^2$) and χ_{10}^2 =9.71 [.47] (if the cointegration space is estimated under $H_0^1 \cap H_0^2 \cap H_0^3$).

The parameters of equation (4) above turn out to be pretty stable in recent times, both within and out of the estimation sample. In this respect, Figure 2 shows their recursive estimates from 1993:Q4 onwards when the equation is estimated over the extended period up to 1998:Q4. More precisely, the figure plots the deviations of the long-run income, spread and inflation parameters with respect to the values reported in (4) above. Similar recursive estimates for the deviations from the hypothesis of long-run price homogeneity are also shown in the graph. In this latter respect, the recursive estimates suggest that long-run

[27] Some evidence at the MU level on the relevance of wealth in the long-run demand for money function can be found in Fase and Winder (1997,1999).

[28] The relative size of these two coefficients is somewhat counter-intuitive if the latter were to be interpreted as the sensibility of money demand to non-financial assets. As argued in Section 2.1, however, this is not the only possible interpretation for the inclusion of the inflation rate in dynamic models of money demand.

price homogeneity, which was assumed in Section 2 and which is maintained throughout the paper, is an acceptable characterisation for our estimates of long-run money demand.

4.2 A dynamic model of money demand

Whilst the unrestricted ADL model is valuable for computing the long-run equation (4), it is clearly over-parameterised for many other purposes. Therefore, we proceeded in a second step to estimate a parsimonious dynamic model for money demand which were both economically interpretable and statistically acceptable. The estimation results for the period 1980:Q4-1997:Q2 are summarised in equation (5) below (t-ratios in parentheses):

$$\begin{aligned}\Delta(m-p)_t = & -\underset{(11.07)}{.739} + \underset{(1.86)}{.075}\ \Delta^2 y_t + \underset{(2.63)}{.194}\ \overline{\Delta s_t} - \underset{(4.43)}{.359}\ \Delta l_{t-1} - \underset{(10.73)}{.526}\ \overline{\Delta\pi_t} \\ & - \underset{(11.17)}{.136}\ [(m-p) - 1.140y + .820(l-s) + 1.462\pi]_{t-2} - \underset{(4.92)}{.009}\, DUM86_t + \hat{\omega}_{1t}\end{aligned} \tag{5}$$

T=67 (1980:Q4-1997:Q2)	R^2=0.80	σ=0.231%	DW=2.17
LM(1)= .478 [.49]	*LM(4)*=.137 [.71]	*LM(1,4)*=.787 [.13]	
ARCH(4)=.286 [.89]	*HET*=.650 [.79]	*NORM*=.824 [.66]	
RESET=.164 [.69]	*RED*=.456 [.81]	$HANS^1$=.060	
$HANS^2$=.589	*FOR(6)*=6.28 [.39]	*CHOW(6)*=1.02 [.42]	

where $\overline{\Delta s_t} = \dfrac{\Delta s_t + \Delta s_{t-1}}{2}$ and $\overline{\Delta\pi_t} = \dfrac{\Delta\pi_t + \Delta\pi_{t-1}}{2}$, and where *LM(i)* and *LM(1,i)* stand for the Lagrange multiplier F-tests for residual autocorrelation of order i^{th} and up to the i^{th} order, respectively,[29] *ARCH* is the Engle (1982) F-test for autoregressive conditional heteroskedasticity, *HET* is the White (1980) F-test for heteroskedasticity, *NORM* is the Doornik and Hansen (1994) χ^2-test for normality, *RESET* is the regression specification F-test due to Ramsey (1969), *RED* tests whether model (5) parsimoniously encompasses the unrestricted ADL(2) model, $HANS^1$ and $HANS^2$ are tests for variance and parameter within-sample stability, following Hansen (1992), *FOR* and *CHOW* are the out-of-sample forecast test and the Chow test for parameter stability over the period 1997:Q3-1998:Q4. Figure 3 records the time series of fitted and actual values for *Δ(m-p)*, the scaled residuals

[29] See Harvey (1990) for a description.

from the model, the residual correlogram and the sequence of 1-step forecasts of $\Delta(m\text{-}p)$ for the period 1997:Q3-1998:Q4.

From a statistical point of view, the estimated model appears well specified, with tests showing no signs of residual autocorrelation, heteroskedasticity or non-normality. The reductions involved in moving from the unrestricted ADL(2) to the parsimonious model given by equation (5) are not rejected either (*RED*) and, therefore, equation (5) parsimoniously encompasses the ADL(2). The coefficient of the equilibrium correction term is highly significant, indicating that a long-run relationship exists between real holdings of M3, on the one hand, and real income, the spread and the inflation rate, on the other. The size of this coefficient, however, indicates that disequilibria are corrected only slowly. This is in line with empirical evidence for EU countries, as surveyed in Browne et al. (1997), and suggests that the costs of being out of the equilibrium (or alternatively the benefits of being in equilibrium) are small. The model parameters appear reasonably stable within the estimation sample, as indicated by Hansen's tests, and no major problems are detected when the equation is used for producing one-step ahead forecasts over the six quarters not included in the estimation sample.

Following Urbain (1992), Wu-Haussman tests for weak exogeneity of real income (τ_1), the short-term rate (τ_2), the long-term rate (τ_3), and the inflation rate (τ_4) with respect to the short-run parameters of money demand were also conducted. That involves testing for the significance of the residuals from the marginal unrestricted reduced-form models for *y, s, l* and π in the money demand equation. None of the tests rejects at standard confidence levels the null of weak exogeneity[30], validating the single-equation approach adopted herein for estimation of money demand.

With a view to gaining some insight into the dynamic properties of the estimated money demand equation, Figure 4 shows the cumulative normalised lag weights of prices, real income and the short- and the long-term interest rates obtained from equation (5). The mean lags for *p*, *y*, *s* and *l* are 5.8, 7.4, 5.6 and 4.1 quarters, respectively. The median lags are somewhat shorter: 4.1, 6.2, 4.5 and 2.5 quarters, respectively. The short-run elasticities of nominal money with respect to prices are close to zero, in contrast with the estimated long-run elasticities. This result, which is often found in the empirical literature on money demand, is consistent — as argued in Ericsson (1999) — with Ss-models of money demand, with short-run factors determining the changes in money holdings given desired

[30] Results were as follows: τ_1= .083 (.77), τ_2=.555 (.46), τ_3= .178 (.67) and τ_4= .399 (.53). The joint test yields: τ=.415 (.797).

upper and lower bands and longer-run factors determining the bands themselves. Money demand is estimated to respond more rapidly to changes in the alternative rate than to changes in the own rate. That feature is of particular interest if the central bank were to exert some short- to medium-term control on the money stock.[31] It should be born in mind, however, that the experiment conducted in computing the lag weights involves a rather unrealistic assumption, namely that the r.h.s. variables are orthogonal. That precludes policy conclusions. In Section 5 a more realistic simulation is made within a multivariate framework which allows for the interplay of the different variables in the system.

Finally, as regards parameter stability, equation (5) was estimated recursively over the extended sample 1980:Q4-1998:Q4. A graphical summary of the results from 1990:Q4 to 1998:Q4 is shown in Figure 5. The figure records the recursively estimated coefficients plus/minus twice their standard errors together with other relevant output from the recursive estimation. The latter includes: residual sum of squares as the sample increases, standardised innovations, 1-step forecast errors plus/minus twice the recursively estimated equation standard error, 1-step ahead Chow tests, and breakpoint Chow tests scaled by their 5% significance values. The estimated parameters appear constant and significant over most of the sample. The Chow tests do not reveal any major non-constancy either. Estimation of equations (4) and (5) over the full sample 1980:Q4-1998:Q4 provides the following results:

$$(m-p)_t = \underset{(.058)}{1.125}\, y_t - \underset{(.359)}{0.865}\,(l_t - s_t) - \underset{(.329)}{1.512}\,\pi_t + \hat{u}_t \qquad (4.a)$$

$$\begin{aligned}\Delta(m-p)_t = & -\underset{(10.99)}{.690} + \underset{(1.78)}{.071}\,\Delta^2 y_t + \underset{(2.65)}{.194}\,\overline{\Delta s_t} - \underset{(4.47)}{.353}\,\Delta l_{t-1} - \underset{(10.79)}{.526}\,\overline{\Delta\pi_t} \\ & - \underset{(11.10)}{.132}\,[(m-p) - 1.125y + .865(l-s) + 1.512\pi]_{t-2} - \underset{(5.00)}{.0095}\,DUM86_t + \hat{\omega}_{1t}\end{aligned} \qquad (5.a)$$

T=73 (1980:Q4-1998:Q4)	R^2=0.80	σ=0.230%	DW=2.18
$LM(1)$= .649 [.42]	$LM(4)$=.319 [.57]	$LM(1,4)$=.734 [.57]	
$ARCH(4)$=.532 [.71]	HET=.58 [.85]	$NORM$=1.53 [.47]	
$RESET$=.158 [.69]	RED=.44 [.82]	$HANS^1$=.041	$HANS^2$=.622

31 See Cabrero et al. (1998).

4.3 Some results on aggregation

As pointed out in Section 2.2, a number of methods for the calculation of historical aggregated series for the euro area have been employed in existing empirical work. It is also fair to say that any choice between them is somewhat arbitrary, reflecting the fundamental problem that it is only from 1999 onwards that a single currency is in place. In particular, the time series for area-wide M3 employed above is calculated backwards from September 1997 M3 levels in euro, using an output-weighted average of participating countries' developments. As discussed, this method of aggregation provides for a coherent historical calculation of euro area monetary aggregates, on the one hand, and the r.h.s. variables in the money demand equation, on the other. Of particular interest is that the resulting growth rates for M3 are consistent with the concept of area-wide inflation most widely used in economic analysis, i.e. a weighted average of countries' inflation rates.

One drawback, however, is that the aggregation method departs from the compilation procedures for euro area monetary which are in place since the start of Stage Three, whereby national data are aggregated in a common currency, the euro, using the irrevocable conversion rates announced on 31 December 1998. Since this may potentially have implications if the estimated models were to be used in the new context, we explore below the effect that the use of euro area M3 compiled on the basis of fixed conversion rates (as published in ECB, 1999b) may have on our estimates whilst noting that a discrepancy with respect to the way in which the remaining variables in our analysis are calculated is introduced.

On a conceptual basis, M3 growth figures resulting from both aggregation procedures can be interpreted as weighted averages of the estimated historical national contributions with weights depending on the method employed. Under the first aggregation method, weights are fixed and are calculated as countries' shares in area-wide GDP evaluated at PPP rates. Under the second, weights are time varying and depend on countries' shares on area-wide money M3 evaluated at fixed conversion rates. On average, the second method gives more weight to countries with higher liquidity ratios.

At the empirical level, a first conclusion that can be drawn from visual inspection of Figure 6 is that the difference between the two aggregation procedures cannot be overstated. The figure plots the levels (Figure 6a) and the quarterly growth rates (Figure 6b) of M3 resulting from the application of the two aggregation methods. Both series in levels move closely together for most of the sample period with differences mainly concentrated at the beginning of the eighties. The differences in the quarterly growth rates are of moderate size in statistical terms when compared, for instance, with the standard error reported in equation (5) above. They are quite persistent, however, implying long-lasting effects when the levels of the series are considered. Thus, during the first half of the eighties the output-weighted M3 grew around 5% more in cumulative terms than the corresponding aggregate computed using the fixed conversion rates.

Formal analysis using the Johansen procedure confirms that both time series do cointegrate with cointegrating vector (1, -1). That in turn suggests that the cointegration results reported in Section 3 would be little affected by the change in the aggregation method for M3. Table 2 confirms this intuition by reproducing the analysis in Section 3 and Table 1 for the monetary aggregate M3 computed using the fixed conversion rates. None of the conclusions in Section 3 are changed by the evidence presented in the table.

As for the short-run dynamics, we proceed by estimating equations (4) and (5) over the full sample 1980:Q4-1998:Q4 using M3 computed using the fixed conversion rates. The results are summarised in equations (4.b) and (5.b) below and in Figure 7, which show some graphical evaluation statistics:

$$(\tilde{m}-p)_t = \underset{(.059)}{1.111}\, y_t - \underset{(.391)}{0.710}\,(l_t - s_t) - \underset{(.370)}{1.201}\,\pi_t + \hat{u}_t \qquad (4.b)$$

$$\begin{aligned}\Delta(\tilde{m}-p)_t = & -\underset{(11.69)}{.635} + \underset{(1.82)}{.074}\,\Delta^2 y_t + \underset{(3.07)}{.114}\,\overline{\Delta s_t} - \underset{(3.95)}{.320}\,\Delta l_{t-1} - \underset{(9.71)}{.478}\,\overline{\Delta\pi_t} \\ & - \underset{(11.81)}{.126}\,[\,(\tilde{m}-p) - 1.111y + .710(l-s) + 1.201\pi\,]_{t-2} - \underset{(4.94)}{.0096}\,DUM86_t + \hat{\omega}_{1t}\end{aligned} \qquad (5.b)$$

T=73 (1980:Q4-1998:Q4)	R^2=0.80	σ=0.235%	DW=2.03
LM(1)= .022[.88]	*LM(4)*=.904 [.35]	*LM(1,4)*=1.40 [.24]	
ARCH(4)=.624 [.65]	*HET*=.319 [.98]	*NORM*=.234 [.89]	
RESET=.319 [.98]	*RED*=.58 [.72]	$HANS^1$=.105	$HANS^2$=.679

where $\tilde{m}$ denotes that the historical series for M3 is calculated using the fixed conversion rates. As before, equation (4.b) is obtained as the long-run solution of an ADL(2) model and equation (5.b) parsimoniously encompasses the unrestricted autoregressive distributed lag model.

Comparison of (4.b) and (5.b) with (4.a) and (5.a) indicates that the change of aggregation method for M3 does not have any noticeable impact either on the long- or the short-run parameters of money demand, with differences in point estimates always well within one standard error. That in turn suggests that it is unlikely that the shift to the calculation of euro area M3 using fixed conversion rates may in itself render our results useless in the context of Stage Three.

Although (based on the respective estimated standard errors) there appears to be little to choose between both aggregation procedures, encompassing tests were conducted with a view to discriminating between the two competing models (5.a) and (5.b). In order to deal with different dependent variables, we follow the approach that Ericsson, Hendry and Tran (1994) take for testing models with seasonal adjusted data versus models with non-seasonally adjusted data. In this context, the hypothesis that (5.a) encompasses (5.b) [H_0^a] can be tested by adding $\Delta(m-\tilde{m})$ and the error correction term in (5.b) to equation (5.a) and testing for their joint significance. Conversely, the null that (5b) encompasses model (5a) [H_0^b] can be tested by adding $\Delta(m-\tilde{m})$ and the error correction term in (5.a) to equation (5.b). Likelihood ratio tests for these hypotheses provide the following results (p-values in parentheses): $\chi^2(2)$= 1.29 (.524) and $\chi^2(2)$= 4.25 (.120) for the nulls H_0^a and H_0^b, respectively. Therefore, although neither H_0^a nor H_0^b can be formally rejected at standard confidence levels, it appears that the evidence for H_0^b is less compelling, somewhat favouring (5.a) over (5.b). Recursive computation of the encompassing tests tends to confirm this conclusion.

5 An impulse response analysis of the multivariate money demand system

In order to overcome the rather unrealistic assumptions which underlie the analysis of the dynamic properties of money demand in Section 4.2, we now step back to the multivariate money demand system and investigate the dynamic interactions of all its variables by means of the system's impulse response functions. The computation of the impulse response functions starts from the system incorporating the long-run money demand relationship (4) within the cointegration space estimated under $H_0^1 \cap H_0^2 \cap H_0^3$. The employed methods which are extensions of structural VAR methods to cointegrated systems have recently been developed in Vlaar (1998) and are briefly reviewed in Annex 2. In Vlaar and Schuberth (1999) these methods are applied to address the issue of controllability of a broad monetary aggregate which was computed from old harmonised M3 data for the EU countries without Luxembourg.

5.1 The identification of the structural innovations

Given the outcome of our empirical analysis of the 5 - dimensional money demand system with 3 cointegrating vectors, we explore the system's dynamic properties by means of its impulse responses, $\{B_i\}_{i=0}^{\infty}$, to a set of underlying structural innovations. The identification of these structural innovations will be achieved by separating them along two dimensions. First, we separate 2 structural innovations having permanent effects from 3 innovations with merely transitory effects. Within each of the two subsets, we then identify innovations which are assumed to be related to monetary policy. In view of this objective, let us tentatively label the innovations — in order of their occurrence within the system — as an aggregate supply shock, a change in the monetary policy objective, an aggregate demand shock, a money demand shock and an interest rate shock with the first two having permanent effects.[32]

Following Vlaar and Schuberth (1999), the two monetary policy shocks are assumed to have either permanent or merely transitory effects. The first, which is labelled as a change in the monetary policy objective, is thought to capture the nominal stochastic trend within our sample. The second is assumed to capture the transitory element of monetary policy. Our label suggests that it should be likely to show up as an innovation to the short-term nominal interest rate which is supposed to be the monetary policy instrument.[33] Aggregate demand and aggregate supply shocks are considered to be the most important factors driving real GDP which, in turn, acts as the scale variable determining money demand. Beyond being orthogonal to the other structural innovations, the money demand shock is assumed to be orthogonal to any contemporaneous variable affecting money demand. It should thus closely resemble the error term in our conditional money demand model.

Starting from an estimate of the system's covariance matrix Σ and its so-called total impact matrix $C(1)$, we have to impose a set of at least 10 independent restrictions on either its contemporaneous impact matrix $B_0 = B$, where $\Sigma = BB'$, or its long-run impact matrix

[32] As a matter of fact, the economic variables comprising the system will be subject to a larger number of shocks (shocks to the term structure, foreign interest rate shocks, exchange rate shocks, etc.). In this connection, it is clear that the empirical model under investigation is unlikely to encompass all the relevant aspects of the monetary transmission process. From a methodological point of view, that raises the question of whether the identified shocks can be interpreted as structural as incorporated in the underlying assumption of their orthogonality. The choice of the above labels reflects the underlying untested assumption that the respectively labelled shocks are the most relevant ones for purposes at hand. Ericsson, Hendry and Mizon (1998) provide an insightful discussion of the problems involved in interpreting impulse response analysis.

[33] The identification of the monetary policy shocks disregards any problem which may be due to the operation of the EMS according to which a number of participating countries had to use the short-term interest rate as an instrument for fixing the exchange rate instead of gearing it directly to the ultimate objective of monetary policy.

$B_\infty = C(1)B$ in order to identify the vector of structural innovations.[34] Since any particular set of identifying restrictions will be subject to controversial discussion, we have chosen to follow a conservative two-step identification strategy. In the first step, we impose a set of exactly identifying restrictions which borrow as far as possible from the outcome of our empirical analysis but leave the effects of the two monetary policy shocks unconstrained. In the second step, we then impose a set of additional, over-identifying restrictions which are designed to characterise the latter. As these additional restrictions must be considered crucial a priori, they will be tested for their compatibility with the data.

To separate the 2 shocks having permanent effects from the 3 shocks with merely transitory effects we start with restricting the long-run impact of the latter:

(R.1) The aggregate demand shock, the money demand shock and the interest rate shock have a zero long-run impact on the levels of all the variables within the cointegrated system, i.e. $(C(1)B)_{ij} = 0$ for $i = 1,...,5$ and $j = 3,4,5$.

In order to mutually identify the 2 shocks with permanent effects, i.e. to separate the effects of the aggregate supply shock from those of the change in the monetary policy objective, we impose a single restriction on the long-run impact of the former:

(R.2) The aggregate supply shock does not affect GDP inflation in the long run, i.e. $(C(1)B)_{51} = 0$.

Restriction (R.2) implies that monetary policy anchors inflation which is consistent with the widely held view that inflation is a monetary phenomenon in the long run.

To mutually identify the 3 shocks with merely transitory effects, a set of at least 3 restrictions have to be imposed on their contemporaneous impacts. Here, we assume:

(R.3) GDP inflation is not contemporaneously affected by the aggregate demand shock, i.e. $B_{53} = 0$.

(R.4) The money demand shock does not contemporaneously affect real GDP, the nominal short and long-term interest rates and GDP inflation, i.e. $B_{24} = B_{34} = B_{44} = B_{54} = 0$.

Restriction (R.3) corresponds to the assumption that inflation adjustment is sluggish,[35] while the restrictions given by (R.4) reflect our assumption that the money demand shock is

[34] Refer to Annex 2 for details.

[35] This view is widely supported by empirical evidence and has been adopted, for instance, by Fuhrer and Moore (1995) in modelling the inflation process.

orthogonal to all contemporaneous variables entering the conditional money demand model. Thereby it is implicitly assumed that money supply fully accommodates contemporaneous movements in money demand which, in turn, is in line with our presumption that the short-term nominal interest rate serves as the monetary policy instrument.

Up to now a set of 21 restrictions are imposed. Due to the reduced rank of the total impact matrix $C(1)$, however, only 6 out of the 15 zero restrictions given by (R.1) are linearly independent. Similarly, only 2 out of the 4 zero restrictions given by (R.4) turn out to be independent.[36] Disregarding the 11 redundant restrictions, there are in total 10 independent restrictions which indeed exactly identify B and, thus, the structural innovations.

The effects of the monetary policy shocks have been left unrestricted for the time being. The identification of these shocks may thus suffer from capturing influences which affect the variables of our money demand system but which are not due to monetary policy. Therefore, in order to end up with a more thorough characterisation of the effects of monetary policy, we impose a set of over-identifying restrictions. Firstly, we assume:

(R.5) Real GDP is not contemporaneously affected by both the change in the monetary policy objective and the interest rate shock, i.e. $B_{22} = B_{25} = 0$.

These restrictions are consistent with the assumption that monetary policy shocks affect real activity only with some delay which seems plausible for a system based on quarterly data. In structural VAR analyses of quarterly US data, for instance, the restriction on the impact response of real GDP proves to be an identifying assumption which makes the impulse responses of the VAR models match the a priori views on the effects of a monetary policy shock.[37]

Secondly, we impose:

(R.6) A change in the monetary policy objective does not affect real GDP in the long run, i.e. $(C(1)B)_{22} = 0$.

[36] This outcome is due to the particular set of zero restrictions we have imposed on the vector of the loading matrix corresponding to the long-run money demand relationship. These zero restrictions imply that the orthogonal complement of the loading matrix has a zero row which, in turn, implies that the total impact matrix of the system's reduced-form representation has a zero column. Given two zero restrictions on the contemporaneous impact vector of the money demand shock, we then end up with a homogenous linear equation system which determines the two remaining elements of the contemporaneous impact vector that do not correspond to real money holdings to be zero too. But this means that any pair of zero restrictions imposed on its contemporaneous impact will suffice to identify the money demand shock.

[37] See Bernanke and Blinder (1992) and Bernanke and Mihov (1995).

This restriction implies that real GDP will be exclusively driven by supply shocks in the long run, where the supply shocks may be interpreted as realisations of an underlying stochastic productivity trend.

Since the level of real GDP will not be affected by the level of the inflation rate in the long run, neutrality, indeed super-neutrality of monetary policy is implied by restriction (R.6), whereas restriction (R.2) guarantees that inflation is a monetary phenomenon in the long run. Taken together, these two restrictions imply that the long-run outcome of the impulse response analysis is dichotomous. Although this may be overly rigid in view of empirical findings, it will be useful in discussing the impulse response functions of the multivariate money demand system by providing a benchmark.

Individually as well as jointly conducted Likelihood Ratio tests accept the set of over-identifying restrictions given by (R.5) and (R.6). The statistics of the individually conducted tests which have asymptotic χ^2 - distributions with two and one degrees of freedom, respectively, are 0.23 and 0.51. The statistic of the jointly conducted test which is asymptotically χ^2 - distributed with three degrees of freedom yields 0.54. As the outcome of these tests gives evidence that the over-identifying restrictions are compatible with the data, we have chosen to rely on the full set of restrictions given by (R.1) - (R.6) for the remainder of our analysis. The estimate of the over-identified contemporaneous impact matrix is:

$$\hat{B} = \begin{bmatrix} \underset{(0.09)}{0.14} & \underset{(0.07)}{0.22} & \underset{(0.05)}{0.03} & \underset{(0.02)}{0.21} & \underset{(0.04)}{0.15} \\ \underset{(0.13)}{0.32} & \underset{(---)}{0.00} & \underset{(0.08)}{0.32} & \underset{(---)}{0.00} & \underset{(---)}{0.00} \\ \underset{(0.18)}{-0.47} & \underset{(0.25)}{-0.31} & \underset{(0.11)}{0.24} & \underset{(---)}{0.00} & \underset{(0.08)}{0.15} \\ \underset{(0.20)}{-0.12} & \underset{(0.15)}{-0.23} & \underset{(0.12)}{0.01} & \underset{(---)}{0.00} & \underset{(0.06)}{0.24} \\ \underset{(0.32)}{-0.52} & \underset{(0.19)}{-0.92} & \underset{(---)}{0.00} & \underset{(---)}{0.00} & \underset{(0.13)}{-0.45} \end{bmatrix}$$

with its estimated asymptotic standard errors given in parentheses.

As a first rough check of its plausibility, the above identification scheme should guarantee that at least the impact response of the variable which is considered to be most closely related to a particular structural shock is significant. This seems to hold true for all structural shocks but the interest rate shock. The impact responses of real GDP to both an aggregate supply shock and an aggregate demand shock turn out to be significant. Note that a positive aggregate supply shock is accompanied by a decrease in the nominal short-term interest rate, whereas a positive aggregate demand shock, which may build up inflationary pressures, is accompanied by an increase in the nominal short-term interest rate. Both interest rate responses which are consistent with an easing or a tightening of monetary policy, respectively, prove to be significant.

A downward change in the monetary policy objective results in a significant decline in the inflation rate though merely in an insignificant move of the nominal interest rates. As imposed, a positive money demand shock induces solely an increase in real money holdings. Note that the impact response of real money holdings is of the same order of magnitude as the estimated standard error of the residuals of the conditional money demand model. The impact effect turns out to be 0.21 while the estimated standard error of the residuals is 0.23. With respect to the identification of the structural innovation which was labelled as an interest rate shock some doubts arise. The estimated impact response of the nominal short-term interest rate proves to be merely borderline significant. On the contrary, the impact responses of real money holdings, the long-term nominal interest rate and, in particular, the inflation rate turn out to be significant. But let us postpone addressing this issue until discussing the results of the impulse response analysis. These are presented in the next subsection.

5.2 The structural impulse response functions

Based on the full set of restrictions given by (R.1) - (R.6), the Figures 8 - 12 show the structural impulse response functions of the variables of the cointegrated system, i.e. real money holdings $m-p$, real GDP y, the short- and long-term nominal interest rates s and l, and (annualised) GDP inflation π. In addition, we report the impulse response functions for a set of derived variables: the spread $l-s$, the short- and long-term real interest rates $s-\pi$ and $l-\pi$, the GDP deflator p, which is obtained by integrating the (quarterly) rate of GDP inflation $\pi/4$, and nominal money holdings m. In order to give an indication of the underlying statistical uncertainty, the figures display the impulse response functions plus/minus twice their asymptotic standard errors. As shown by these standard errors quite a few of the impulse response functions are statistically insignificant. Clearly this reflects the sampling variability in the estimated parameters of the cointegrated system and the large amount of uncertainty due to the system's stochastic trends. Any interpretation of the impulse response functions has therefore to bear in mind that they may provide only imprecise information about the true underlying economic relations.

(a) A positive aggregate supply shock

A shock to aggregate supply (Figure 8) moves real GDP and GDP inflation in opposite directions on impact. These movements are consistent with the assumption that firms are operating under monopolistic competition by setting prices as a mark-up on average production costs: If a positive aggregate supply shock resembles an upward shift in productivity, production will increase while prices will decline due to the reduction in average production costs. Whereas the effect on GDP is permanent, inflation returns to its steady-state level in the long run, as has been imposed in (R.2) above. Due to the temporary fall in inflation there is some leeway for monetary policy to ease. While there is a significant instantaneous fall in the nominal short-term interest rate, however, the short-term real interest rate declines only gradually. Consistent with the expectation theory of the term structure, the long-term nominal interest rate decreases but less than the short-term rate, thereby inducing a temporary increase in their spread. Both the short and long-term nominal interest rates return to their original steady-state levels. Given that the effect of the supply shock on GDP inflation dies out, this outcome reflects our empirical findings that the long-term real interest rate as well as the spread constitute cointegrating relationships which are automatically accounted for when computing the system's impulse response functions. Similarly, real money holdings increase in line with the long-run money demand specification as obtained from our conditional analysis. Whereas the effect on nominal money holdings can be neglected, there is a permanent decrease in the price level which is due to the temporary fall in inflation.

(b) A downward change in the monetary policy objective

Changes in the monetary policy objective are thought to capture the nominal stochastic trend in the sample. Given the historical evidence for the member countries of the monetary union, a downward change in the monetary policy objective has a natural interpretation in terms of a possibly pre-announced commitment to lower inflation. Of course, as there was no single monetary policy authority which could have announced a decline in its inflation objective, the chosen label remains fictitious but nevertheless serves as an intuitive means of interpreting the average disinflation policies within the member countries. At a theoretical level, we can separate the liquidity effect of a disinflation policy from its expectation effect.[38] As long as the commitment to decrease inflation is perfectly credible and prices are flexible it would be feasible to reduce inflation on impact. Moreover, the lowering of inflation would even be achievable by adjusting the nominal short-term interest

[38] See Christiano and Eichenbaum (1992) for a thorough discussion of the operation of the liquidity and expectation effects in a dynamic general equilibrium setting; and Christiano, Eichenbaum and Evans (1996) for related empirical evidence.

rate downwards instantaneously if monetary policy operated exclusively via the expectation channel.

Although disregarding the liquidity effect seems somewhat unrealistic in view of empirical findings, the computed impulse response functions (Figure 9) are indeed in line with such a perception: Inflation falls on impact, even overshooting its new steady state level, and the nominal short-term interest rate declines. Real GDP is not contemporaneously affected since it has been imposed in (R.5) but the short-term real interest rate rises on impact. Due to the temporary rise of the short-term real interest rate, real GDP is transitorily suppressed in the subsequent periods but returns to its steady state level in the long run, as imposed in (R.6). This hump-shaped adjustment path of real GDP suggests that disinflation is costly even if there is a strong initial expectation effect. Given the stationarity of the real rates, both nominal interest rates decline by the same amount as the inflation rate in the long run. Furthermore, consistent with our long-run money demand relationship, real money holdings stabilise at a higher level reflecting the lower opportunity costs of holding money due to the decline in the steady-state inflation rate. The price level and nominal money holdings drift away as implied by the permanently lower inflation rate.

(c) A positive aggregate demand shock

The impulse responses to the aggregate demand shock (Figure 10) are in line with what we would have expected a priori. The significant increase in real GDP creates inflationary pressures which are counteracted by a tightening of monetary policy in a forward-looking manner via an increase in the nominal short-term interest rate. Due to the sluggishness of inflation, as imposed in (R.3), this response feeds through to the real short-term interest rate one to one. In line with the expectation theory, the long-term nominal interest rate shows a more moderate rise than the short-term rate, thereby implying a flattening of the yield curve. Whereas real GDP declines monotonously, GDP inflation follows a typical humped-shaped adjustment path. Real money holdings build up with some delay but return to their steady state in the long run. On the other hand, nominal money holdings increase permanently. As real, but not nominal money holdings are restricted to be unaffected in the long run, the increase in the latter just compensates for the permanent increase in the GDP deflator which is caused by the temporary rise in inflation.

(d) A positive money demand shock

Given the restrictions imposed in (R.4), the shock in money demand (Figure 11) affects only real money holdings contemporaneously. In the subsequent periods, however, real GDP rises significantly. This response may be interpreted in terms of a wealth effect operating through temporarily accumulated real balances. The increase in real GDP, in turn, causes a rise in GDP inflation which is counteracted by a tightening of monetary policy. In view of these findings, it is evident that the system's impulse responses to the money demand shock closely resemble the pattern of the impulse responses to an aggregate demand shock as far as the lag due to the operation of the real balance effect is taken into account. Accordingly, both the GDP deflator and nominal money holdings increase permanently, whereas the effects on real money holdings die out in the long run. Note that the rise in nominal money holdings proves to lead the increase in the price level due to the sluggish adjustment of inflation. Of course, as this outcome is specific to the money demand shock under investigation it does not follow that nominal money is in general a leading indicator for future price level developments.

(e) A positive interest rate shock

The interest rate shock (Figure 12) is assumed to capture the transitory element of monetary policy. Consistent with the commonly-held view on the outcome of a monetary contraction, the impulse responses of real GDP show a hump-shaped pattern: Real GDP starts to decline with some delay, as implied by the contemporaneous restriction in (R.5), reaches a trough after a few periods and starts to recover subsequently. As in the case of the change in the monetary policy objective, the decline in real GDP follows an increase in the short-term real interest rate driven by both the tightening of monetary policy and the initial drop in the inflation rate.

Real money holdings increase on impact in response to the interest rate shock. Since real GDP is not affected contemporaneously, there remain three variables which simultaneously account for the response of real money holdings. Firstly, the demand for real money holdings is stimulated by the increase in the nominal short-term interest rate (measuring the rate of return of the interest bearing components within the broad monetary aggregate under investigation). Secondly, the increase in the nominal long-term interest rate (measuring the opportunity cost of real money holdings in terms of alternative financial assets) dampens the demand for real money holdings. Thirdly, the demand for real money holdings is increased by the initial drop in inflation (measuring the opportunity cost of real money holdings in terms of real assets). Subsequent to their initial increase, real money holdings start to fall. This fall is presumably driven by the temporary decline in real GDP since the joint effect of the nominal interest rates, as measured by their spread, and inflation appears to be of second order magnitude.

Given the impulse responses of real money holdings and the GDP deflator, nominal money holdings, though imprecisely estimated, decline in the medium to long run. This outcome gives some tentative evidence that, over those horizons, the money stock is controllable via the various channels of transmission through which monetary policy operates. To the contrary, the mute short-run response of nominal money holdings appears to indicate that short-term controllability of M3 would likely be problematic. These results, however, may be adversely affected by the strong impact response of the inflation rate which turns out to be implausible. On the one hand, the impact response of inflation might capture effects which are due to monetary policy but which are operating through channels not accounted for in our system as, for instance, the exchange rate channel. On the other hand, the impact response of inflation might be the outcome of shocks which are of relevance for the system but which are not separately identified. Therefore, further investigation is required to obtain a more reliable characterisation of the monetary policy shock and its ultimate effects on money holdings.

6 Conclusions

Given the prominent role assigned to monetary aggregates in the monetary policy strategy of the Eurosystem, the present paper has analysed some empirical evidence on the demand for broad money M3 in the euro area. An attempt to account for the different dynamic interactions between broad money, prices, real income and short and long-term interest rates was also made using SVAR techniques by means of an impulse response analysis of the cointegrated system. A number of conclusions can be drawn from the analysis.

Starting with the system results, three long-run relationships with economic content were found: first, the spread between the long- and the short-term nominal interest rates, consistent with the term structure of interest rates; second, the long- and short-term real interest rates, consistent with the Fisher parity; and third, a long-run demand for broad money M3 in the euro area. Furthermore, the impulse response analysis of the cointegrated system provided an overall picture of the responses of the different variables to a number of structural shocks which we tentatively labelled as: an aggregate demand shock, an aggregate supply shock, a money demand shock, a transitory monetary policy shock and a permanent monetary policy shock. The general patterns of the estimated dynamic responses were broadly consistent with economic theory, though some of the results deserve further investigation.

Turning to our central relation of interest, money demand, the empirical findings are plausible from a number of perspectives. From an economic point of view, the long-run demand for M3 in the euro area shows a simple specification whereby holdings of M3 real balances are determined by real income, the opportunity cost of holding money relative to financial assets not included in the definition of the monetary aggregate and the inflation rate. The signs and magnitudes of the estimated long-run coefficients also appear quite plausible on theoretical grounds. The long-run income elasticity is estimated significantly above one (1.14), a finding which is often interpreted as proxying omitted wealth effects in the demand for money function. The estimated long-run semi-elasticities with respect to the spread and the inflation rate are, respectively, -0.82 and -1.46.

In the short run, changes in M3 holdings were modelled via an equilibrium correction model whereby growth in real M3 is explained by the changes in income, interest rates and inflation, on the one hand, and the extent to which M3 differs from its equilibrium level, on the other. As surveyed in Browne et al. (1997), equilibrium correction models have proven successful in modelling the demand for money in a number of EU countries and in the area as a whole during the last decade. From a statistical point of view, the estimated models for money demand show no sign of mis-specification and appear to be a congruent statistical representation of the ‘average’ process which has driven broad money M3 in the euro area during the sample period under investigation. When compared with the existing empirical evidence on money demand in individual euro area-countries and in the euro-area as a whole, as surveyed for instance in Browne et al. (1997), the estimated models fare quite well in terms of stability, fit, residual standard errors, etc. In particular, the estimates presented support the view reached by many previous studies that area-wide money demand equations outperform most national equations.

The demand for broad money M3 equation turns out to be reasonably constant when estimated recursively. This holds in the face of the dramatic economic changes which euro area countries undertook during the estimation sample and poses the question of whether the estimated models are of any use for monetary policy purposes in the context of Stage Three. This issue is of utmost relevance, as it has been argued that the adoption of a single monetary policy marks a significant regime shift, which could affect empirical economic relationships estimated on the basis of pre-1999 data. The potential effects of the adoption of a single currency are also argued to be of particular importance at the outset of Stage Three in the financial sector of the economy, thereby affecting the behaviour of monetary aggregates in the euro area. Unfortunately, it is not possible to test explicitly for parameter invariance against such historical event, since it has no parallel within our estimation sample, but the estimated models constitute useful benchmarks for monitoring structural change in the early stages of Stage Three and for the evaluation of rival models as they become available. Nonetheless, the empirical finding that the area-wide money demand equation remains constant against a changing background in monetary policy co-ordination between euro area countries during the sample period under investigation suggests that instabilities in the models should not arise because of the introduction of the single currency itself. Parameter constancy when the estimation sample is extended to cover 1997:Q3-1998:Q4, a period in which the alleged instabilities should likely have shown up in anticipation of EMU, is a further argument along this line of reasoning.

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Annex 1: The data

	m*	p	y	s	l	m* *
80Q1	13.90670	3.97120	13.63810	0.12720	0.12060	13.9660
80Q2	13.93170	3.99680	13.63190	0.13220	0.12100	13.9877
80Q3	13.95660	4.02100	13.63200	0.12160	0.12150	14.0099
80Q4	13.98360	4.04170	13.62990	0.12570	0.12770	14.0348
81Q1	14.00840	4.06530	13.62960	0.13360	0.13390	14.0575
81Q2	14.03660	4.08830	13.63620	0.15530	0.14530	14.0846
81Q3	14.05940	4.11470	13.64040	0.16070	0.15150	14.1057
81Q4	14.08420	4.14320	13.64080	0.15360	0.15050	14.1288
82Q1	14.11580	4.16920	13.64270	0.14270	0.14720	14.1584
82Q2	14.14490	4.18950	13.64570	0.14330	0.14470	14.1845
82Q3	14.17560	4.20870	13.64270	0.13280	0.14170	14.2126
82Q4	14.20420	4.22750	13.64210	0.12490	0.13630	14.2373
83Q1	14.23030	4.25180	13.64740	0.11790	0.13090	14.2612
83Q2	14.25170	4.26860	13.65560	0.11600	0.13000	14.2812
83Q3	14.27570	4.28830	13.65790	0.11890	0.13070	14.3028
83Q4	14.29800	4.30740	13.66550	0.11920	0.13050	14.3244
84Q1	14.31670	4.32560	13.67710	0.11300	0.12560	14.3410
84Q2	14.33810	4.33860	13.67050	0.10790	0.12450	14.3608
84Q3	14.36310	4.35100	13.67860	0.10360	0.12070	14.3838
84Q4	14.38490	4.36440	13.68420	0.10070	0.11270	14.4032
85Q1	14.41060	4.37740	13.68610	0.09810	0.10730	14.4271
85Q2	14.42990	4.38920	13.69590	0.09840	0.10640	14.4443
85Q3	14.44690	4.40450	13.70560	0.09240	0.10520	14.4600
85Q4	14.46180	4.41890	13.71140	0.08570	0.10150	14.4741
86Q1	14.47800	4.43670	13.70630	0.08690	0.09720	14.4897
86Q2	14.49320	4.44780	13.72400	0.07920	0.08560	14.5045
86Q3	14.51210	4.45830	13.73020	0.07630	0.08320	14.5233
86Q4	14.53310	4.46570	13.73660	0.07740	0.08330	14.5431
87Q1	14.55290	4.47510	13.72880	0.07950	0.08320	14.5618
87Q2	14.57540	4.48240	13.74550	0.08180	0.08590	14.5837
87Q3	14.59460	4.48790	13.75530	0.08250	0.09320	14.6029
87Q4	14.61470	4.49820	13.76870	0.08230	0.09400	14.6226
88Q1	14.63220	4.50740	13.77590	0.07300	0.08790	14.6394
88Q2	14.65430	4.51580	13.78280	0.07130	0.08740	14.6615
88Q3	14.67580	4.52460	13.79410	0.07870	0.08880	14.6828
88Q4	14.69730	4.53510	13.80620	0.08340	0.08820	14.7036
89Q1	14.71810	4.54710	13.81440	0.09360	0.09280	14.7236
89Q2	14.74180	4.55440	13.82470	0.09540	0.09510	14.7475
89Q3	14.76470	4.56410	13.82740	0.09940	0.09550	14.7714
89Q4	14.78610	4.57680	13.84190	0.10920	0.10080	14.7925
90Q1	14.80890	4.58930	13.85480	0.11060	0.10820	14.8150
90Q2	14.82410	4.60050	13.85890	0.10490	0.10840	14.8308
90Q3	14.84380	4.61110	13.86630	0.10390	0.10930	14.8509
90Q4	14.86640	4.61930	13.86990	0.10890	0.11030	14.8722
91Q1	14.88570	4.63380	13.87530	0.11010	0.10640	14.8920
91Q2	14.90230	4.64840	13.87900	0.10390	0.10140	14.9085
91Q3	14.92250	4.66060	13.88160	0.10440	0.10230	14.9278
91Q4	14.94520	4.67320	13.88710	0.10670	0.09940	14.9477
92Q1	14.96140	4.68460	13.89730	0.10790	0.09600	14.9634

92Q2	14.98560	4.69340	13.89530	0.10920	0.09720	14.9873
92Q3	15.00530	4.70360	13.89370	0.11870	0.10100	15.0066
92Q4	15.02390	4.71160	13.88930	0.11340	0.09670	15.0240
93Q1	15.03940	4.72230	13.88010	0.10610	0.09020	15.0395
93Q2	15.05760	4.73210	13.88120	0.09020	0.08640	15.0568
93Q3	15.06870	4.73800	13.88490	0.08070	0.07600	15.0672
93Q4	15.08550	4.74500	13.88920	0.07380	0.06950	15.0834
94Q1	15.09830	4.75330	13.89700	0.06810	0.07010	15.0960
94Q2	15.10630	4.75740	13.90690	0.06340	0.07950	15.1039
94Q3	15.11240	4.76320	13.91460	0.06360	0.08810	15.1098
94Q4	15.11660	4.76900	13.92300	0.06490	0.09210	15.1138
95Q1	15.12290	4.77790	13.92840	0.06950	0.09330	15.1203
95Q2	15.13390	4.78700	13.93180	0.07150	0.08950	15.1314
95Q3	15.14780	4.79440	13.93440	0.06700	0.08560	15.1455
95Q4	15.16440	4.80060	13.93490	0.06500	0.08130	15.1619
96Q1	15.18100	4.80620	13.94160	0.05630	0.07680	15.1788
96Q2	15.18970	4.81050	13.94520	0.05140	0.07540	15.1879
96Q3	15.19930	4.81370	13.95150	0.05000	0.07270	15.1974
96Q4	15.20810	4.81760	13.95460	0.04590	0.06460	15.2065
97Q1	15.21660	4.82090	13.95790	0.04440	0.06190	15.2158
97Q2	15.22750	4.82430	13.97090	0.04320	0.06190	15.2272
97Q3	15.23910	4.82840	13.97920	0.04320	0.05820	15.2391
97Q4	15.24770	4.83250	13.98520	0.04430	0.05660	15.2477
98Q1	15.25980	4.83610	13.99380	0.04190	0.05110	15.2598
98Q2	15.27370	4.84020	13.99910	0.04050	0.05010	15.2737
98Q3	15.28240	4.84330	14.00590	0.03930	0.04550	15.2824
98Q4	15.29360	4.84740	14.01090	0.03620	0.04150	15.2936

*** computed using fixed weights based on GDP at PPP rates**
**** computed in euro using fixed conversion rates**

Annex 2: A brief review of the methodology

The impulse response analysis of cointegrated systems, which was originally explored in Lütkepohl and Reimers (1992), relied on the Cholesky decomposition of the systems' covariance matrix in order to end up with a structural interpretation of their innovations. It is well known, however, that the identification scheme thus obtained is to some extent arbitrary. In particular, it implies a contemporaneous recursive structure among the systems' variables that depends on their ordering. Similar in vein, Warne (1993) proposed a recursive identification scheme based on the common trend representation of cointegrated systems, which separately identifies the innovations having permanent effects — the so-called common trends — and the innovations with only transitory effects.

In an extension of these approaches, Vlaar (1998) has recently suggested to identify the innovations of cointegrated systems in terms of an underlying theoretical structure by imposing contemporaneous as well as long-run restrictions on their impulse response functions, i.e. the coefficients of their vector moving average representation. Similar to the one proposed by Warne, this identification scheme proves to be more parsimonious than a scheme merely based on contemporaneous restrictions. But in addition, it proves to be more flexible than an identification scheme relying on a particular recursive ordering of the system's variables. Thereby, Vlaar closes the gap between the impulse response analysis of cointegrated systems, which has suffered from the lack of a sound structural interpretation so far, and the analysis of 'structural' vector autoregressive (SVAR) systems as introduced by Blanchard (1989) and Blanchard and Quah (1989), which was confined to stationary variables.

In order to discuss the method proposed by Vlaar in a bit more detail, let $\{z_t\}$ be a sequence of a n - dimensional vector of variables which is generated by an unrestricted vector autoregressive (VAR) process of order k,

$$z_t = \mu + A_1 z_{t-1} + \cdots + A_k z_{t-k} + \varepsilon_t$$

where the sequence of the n - dimensional vector of innovations $\{\varepsilon_t\}$ has a positive-definite covariance matrix $\mathrm{E}[\varepsilon_t \varepsilon_t'] = \Sigma$.

If $\{z_t\}$ is cointegrated of order $(1,1)$ with cointegrating rank r, then, according to the Engle/Granger representation theorem (see Engle and Granger (1987) or Johansen (1995)), the data generating process has the following two equivalent representations:

(a) the vector error correction (VEC) representation

$$\Delta z_t = \mu + \Gamma_1 \Delta z_{t-1} + \cdots + \Gamma_{k-1} \Delta z_{t-k+1} + \alpha\beta' z_{t-k} + \varepsilon_t$$

where α, β are $(n \times r)$ - dimensional matrices with $0 < \text{rank}[\alpha\beta'] = r < n$, and

(b) the vector moving average (VMA) representation

$$\Delta z_t = C(L)(\mu + \varepsilon_t)$$

where $C(L) \equiv I_n - \sum_{i=1}^{\infty} C_i L^i$ is a matrix polynomial of infinite order in the lag operator L, such that $L^i \varepsilon_t = \varepsilon_{t-i}$, and $C(1) = \beta_\perp (\alpha'_\perp (I_n - \sum_{i=1}^{k-1} \Gamma_i) \beta_\perp)^{-1} \alpha'_\perp$ with $\text{rank}[C(1)] = n - r$.

The infinite order VMA representation in first differences can be integrated to obtain a finite order VMA representation in levels,

$$z_t = z_0 + C(1)\mu t + C(1) \sum_{\tau=1}^{t} \varepsilon_\tau + \tilde{C}(L)\varepsilon_t$$

where $C(L) = C(1) + \tilde{C}(L)(1 - L)$ with $\tilde{C}_i = -\sum_{j=i+1}^{\infty} C_j$.

Due to the reduced rank of the total impact matrix $C(1)$ only $n - r$ linear combinations of the innovations — the common trends — have permanent effects on the levels, whereas the effects of the remaining r linear combinations die out in the long run.[1]

The above representations are reduced-form representations. In order to give them a structural interpretation, the sequence of the vector of reduced-form innovations $\{\varepsilon_t\}$ is assumed to be related to the sequence of a n - dimensional vector of orthogonal innovations $\{e_t\}$ with covariance matrix normalised to be $\text{E}[e_t e'_t] = I_n$. The contemporaneous relationship between the reduced-form and the structural innovations is modelled via the non-singular $(n \times n)$ - dimensional matrix B,

$$\varepsilon_t = B e_t$$

In view of this relationship, the covariance matrices of the reduced-form and the structural innovations are related by $\text{E}[\varepsilon_t \varepsilon'_t] = B\text{E}[e_t e'_t]B'$, i.e. $\Sigma = BB'$. Due to the symmetry of Σ and the normalisation of the variances of the structural innovations, already $n(n+1)/2$ non-linear restrictions are imposed on the n^2 elements of the matrix B. Therefore, we

[1] This system property is exploited extensively in the common trend literature. See, e.g., Stock and Watson (1988) and Warne (1993).

have to impose at least $n(n-1)/2$ additional independent restrictions on B in order to identify its elements.

The sequence of the system's impulse responses to the structural innovations, $\{B_i\}_{i=0}^{\infty}$, has an illuminating representation in terms of the coefficients of its VMA representation and the matrix B,[2]

$$B_i \equiv (C(1) + \tilde{C}_i)B, \qquad i = 0,1,\ldots$$

Of course, since $\tilde{C}_0 = I_n - C(1)$ and $\lim_{i\to\infty} \tilde{C}_i = 0$, the contemporaneous impact of the structural innovations is just given by $B_0 = B$, while their long-run impact is measured by $\lim_{i\to\infty} B_i = C(1)B$. In view of these relationships, the identification of the structural innovations can thus be achieved by imposing appropriate restrictions on their long-run impact, i.e. on the elements of the matrix $C(1)B$, and by imposing appropriate restrictions on their contemporaneous impact, i.e. on the elements of the matrix B.

If it is assumed that the $n-r$ common trends of the cointegrated system relate to a particular subset of $n-r$ structural innovations, the effects of the r remaining structural innovations will die out in the long run. This can easily be accomplished by imposing zero restrictions on the corresponding columns of the long-gun impact matrix $C(1)B$. Due to the reduced rank of $C(1)$, however, only $(n-r)r$ out of these nr zero restrictions will be linearly independent. Given the separation of the innovations with permanent effects from those with transitory effects, it will then suffice to impose any set of at least $(n-r)(n-r-1)/2$ independent restrictions on $C(1)B$ or B to mutually identify the innovations with permanent effects. To mutually identify the innovations with transitory effects, we can impose any set of at least $r(r-1)/2$ independent restrictions on B.

If B is identified, its estimate can subsequently be obtained by means of a constrained two-step Maximum Likelihood procedure starting from the reduced-form estimates of the covariance matrix Σ and the long-run impact matrix $C(1)$, where the latter is incorporated in the formulation of the constraints for estimating B. This procedure also permits to test over-identifying restrictions (given a particular set of exactly identifying restrictions). Thereby, the identification of the system's structure is completed and its impulse response functions can be computed.

[2] Whereas the representation of the impulse response functions in terms of the VMA coefficients is illustrative, it is more convenient, from a computational point of view, to derive the impulse response functions from the companion form of the VAR representation of the cointegrated system. For details see, for instance, Lütkepohl (1991), Chap. 2.

To assess the precision of the computed impulse response functions, it is convenient to report their confidence intervals. These may alternatively be obtained by analytical methods based on asymptotic theory or by Bootstrap and simulation-based methods. As the numerical burden of the latter proves to be prohibitive, the approach proposed by Vlaar confines itself to analytical methods. The analytical derivation of confidence intervals starts from the result that the distribution of the estimated coefficients of the underlying VAR model is asymptotically normal. This result, which was originally proved for stationary VAR models, was extended to cointegrated VAR models by Lütkepohl and Reimers. They derived the distribution of the VAR coefficients from the asymptotic distribution of the estimated VEC coefficients assuming the coefficients of the cointegrating relationships known.

In extension of the results by Lütkepohl and Reimers, Vlaar showed that the covariance matrix of the estimator of the VEC coefficients and the estimator of the contemporaneous impact matrix does not have a block-diagonal structure. This is due to the stochastic nature of the imposed long-run restrictions which are non-linear functions of the estimated VEC coefficients. To take account of this stochastic element, Vlaar proposes a first order correction based on the delta method in deriving the analytical confidence intervals of the impulse response functions. He demonstrates that this analytical correction has considerable impact on the shape of the confidence intervals. This holds in particular for the impulse response functions of the transitory shocks that are enforced to die out in the long run due to the imposed zero restrictions.

Table 1. Summary of results from the application of the Johansen procedure: M3 (fixed GDP PPP rates)

Informat. criteria	**k=1**	**k=2**	**k=3**	**k=4**		
SC	-51.87	**-51.93**	-50.66	-49.98		
HQ	-52.58	**-53.14**	-52.38	-52.51		
Test diagnostics (k=2)	**LM(1)**	**LM(4)**	**LM(1-8)**	**N**	**HET**	
	0.50	0.84	0.99	22.19**	0.72	
Cointegration rank	**r=0**	**r<=1**	**r<=2**	**r<=3**	**r<=4**	
trace test	122.99	76.13	39.46	15.44	0.13	
	[107.72**]	[62.25**]	[31.03*]	[11.84]	[0.01]	
Chi-square tests for	**m-p**	**y**	**s**	**l**	**π**	
exclusion	26.43**	23.28**	12.79**	16.73**	26.02**	
stationarity	23.70**	23.82**	15.61**	19.36**	19.61**	
weak-exogeneity	23.14**	1.92	11.41**	16.77**	12.72**	
Tests for structural hypotheses	H_0^1	H_0^2	$H_0^1 \wedge H_0^2$			
	(0,0,-1,1,0)	(0,0,0,1,-1)				
	3.93	0.52	5.21			
	(0.14)	(0.77)	(0.27)			
The cointegrating vectors	β_1	**s.e.**	β_2	**s.e.**	β_3	**s.e.**
m-p	1.000	-	-	-	-	-
y	-1.170	0.034	-	-	-	-
s	1.260	0.170	-1.000	-	-	-
l	-	-	1.000	-	1.000	-
π	-	-	-	-	-1.000	-
The loadings	α_1	**s.e.**	α_2	**s.e.**	α_3	**s.e.**
m-p	-0.138	0.026	-0.250	0.063	0.045	0.062
y	0.029	0.037	0.258	0.090	-0.162	0.084
s	-0.078	0.038	0.028	0.091	-0.186	0.086
l	-0.026	0.024	-0.023	0.059	-0.236	0.056
π	0.015	0.069	-0.009	0.167	0.480	0.159
	H_0^3		$H_0^1 \wedge H_0^2 \wedge H_0^3$			
	$(\alpha_{21}=\alpha_{31}=\alpha_{41}=\alpha_{51}=0)$					
	2.79		9.65			
	(0.25)		(0.29)			

Note: The superscripts * and ** indicate rejections at 10% and 5% significance levels, repectively.

Table 2. Summary of results from the application of the Johansen procedure: M3 (fixed conversion rates)

Informat. criteria	**k=1**	**k=2**	**k=3**	**k=4**		
SC	-51.83	**-51.89**	-50.67	-50.02		
HQ	-52.53	**-53.10**	-52.39	-52.24		
Test diagnostics (k=2)	**LM(1)**	**LM(4)**	**LM(1-8)**	**N**	**HET**	
	0.57	0.90	0.94	22.87**	0.67	
Cointegration rank	**r=0**	**r<=1**	**r<=2**	**r<=3**	**r<=4**	
trace test	119.31	76.54	39.45	15.12	0.19	
	[106.60**]	[64.09**]	[30.94*]	[12.06]	[.01]	
Chi-square tests for	**m-p**	**y**	**s**	**l**	**π**	
exclusion	24.10**	21.90**	10.56**	16.13**	22.63**	
stationarity	24.04**	24.10**	17.16**	19.62**	20.24**	
weak-exogeneity	22.85**	1.71	11.45**	17.05**	13.63**	
Tests for structural hypotheses	**H_0^1**	**H_0^2**	**$H_0^1 \wedge H_0^2$**			
	(0,0,-1,1,0)	(0,0,0,1,-1)				
	4.50	0.18	5.54			
	(0.11)	(0.92)	(0.24)			
The cointegrating vectors	**β_1**	**s.e.**	**β_2**	**s.e.**	**β_3**	**s.e.**
m-p	1.000	-	-	-	-	-
y	-1.160	0.036	-	-	-	-
s	-	-	-1.000	-	-	-
l	-	-	1.000	-	1.000	-
π	0.840	0.188	-	-	-1.000	-
The loadings	**α_1**	**s.e.**	**α_2**	**s.e.**	**α_3**	**s.e.**
m-p	-0.142	0.024	-0.042	0.041	-0.167	0.052
y	0.025	0.036	0.214	0.061	-0.133	0.077
s	-0.059	0.038	0.145	0.063	-0.297	0.080
l	-0.014	0.024	0.021	0.041	-0.276	0.052
π	0.069	0.068	-0.016	0.113	0.532	0.146
	H_0^3 ($\alpha_{21}=\alpha_{31}=\alpha_{41}=\alpha_{51}=0$)		**$H_0^1 \wedge H_0^2 \wedge H_0^3$**			
	1.27		8.97			
	(0.53)		(0.34)			

Note: The superscripts * and ** indicate rejections at 10% and 5% significance levels, repectively.

Figure 1. The data

Fig. 1.a

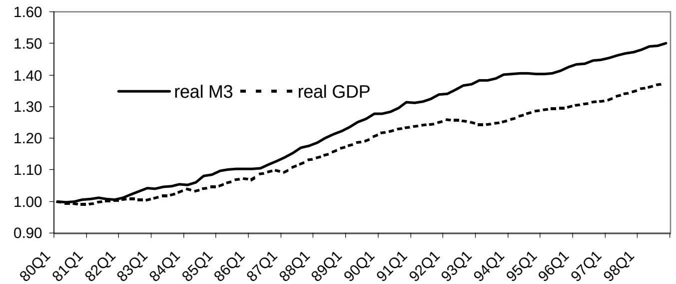


Fig. 1.b

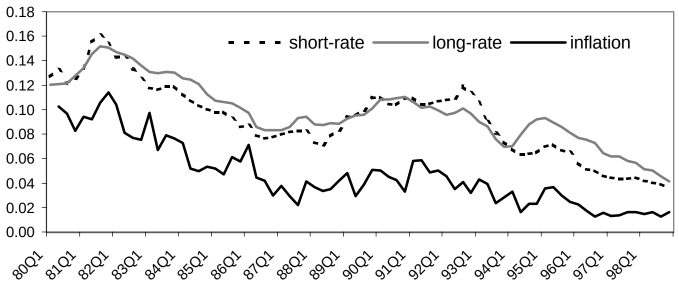


Fig. 1.c

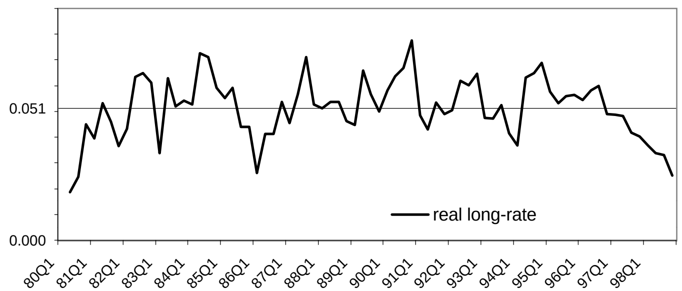


Fig. 1.d

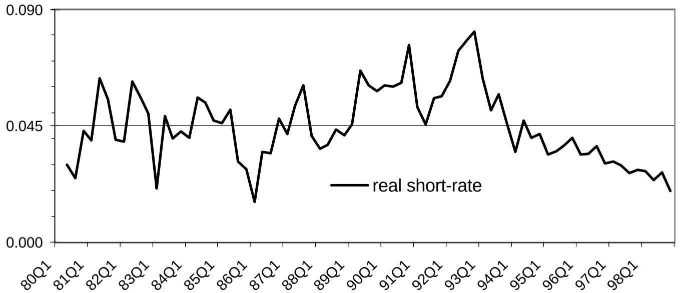


Fig. 1.e

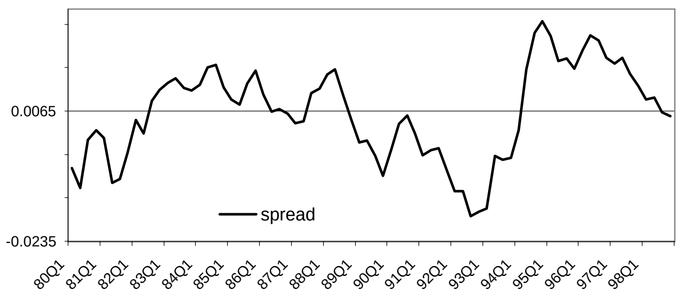


Fig. 1.f

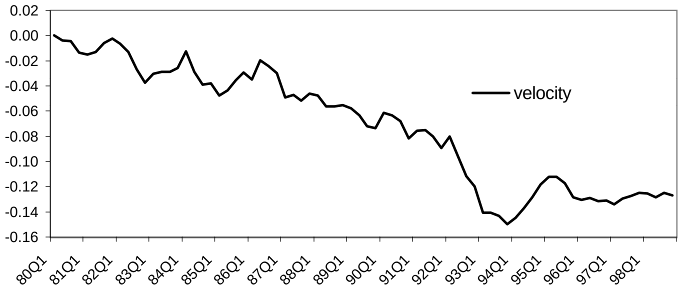


Figure 2. Recursive estimation of eq. (4) over 1994:Q1-1998:Q4: deviations from 1997:Q2 estimates

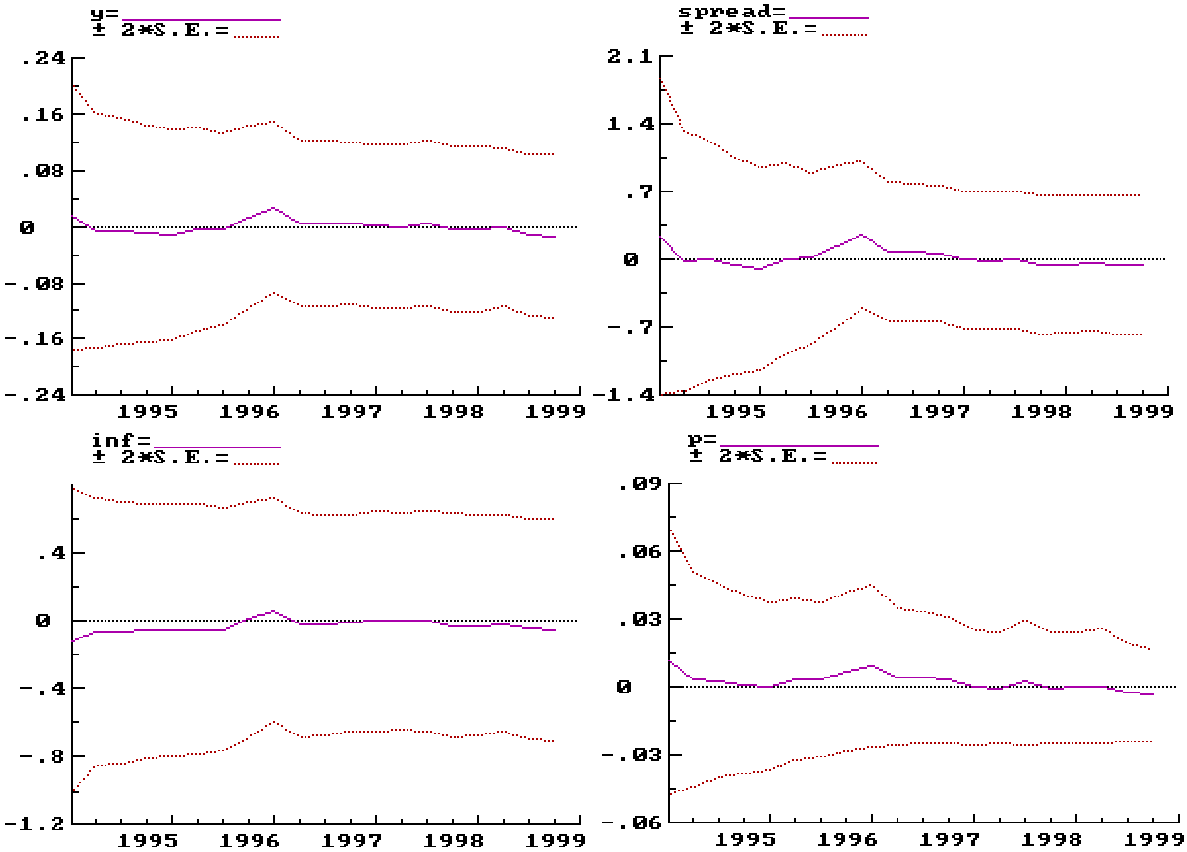


Figure 3. Graphical evaluation statistics for eq. (5)

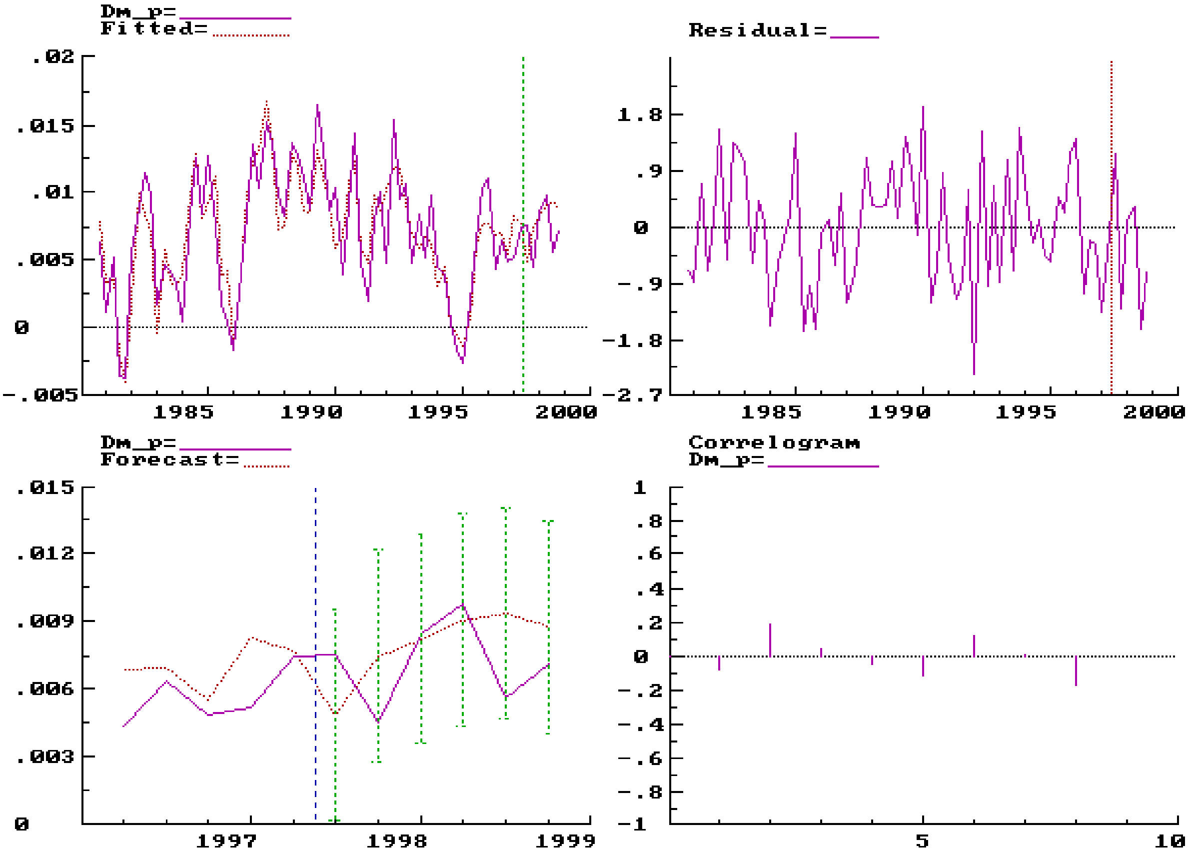


Figure 4. Cumulative normalised lag weights

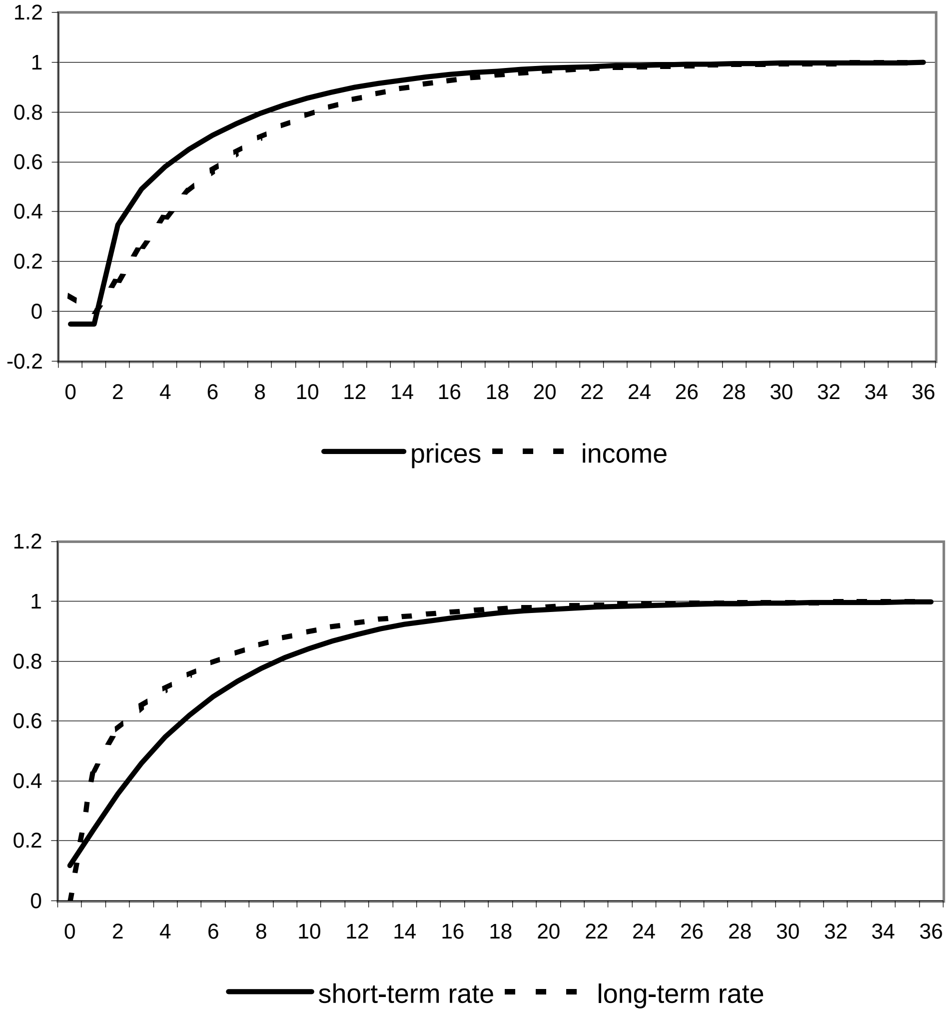


Figure 5. Recursive estimation of eq. 5 over 1990:Q1-1998:Q4

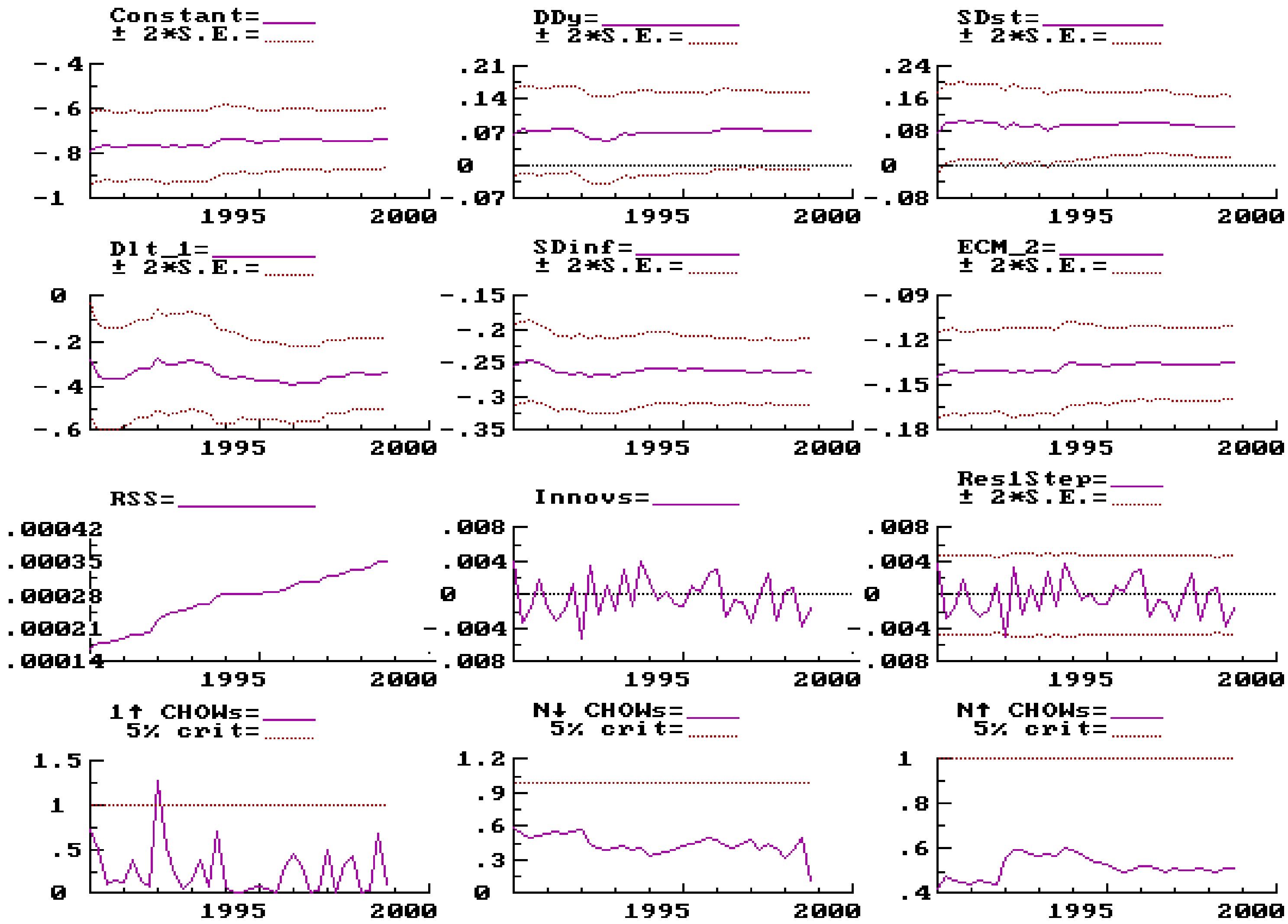


Figure 6

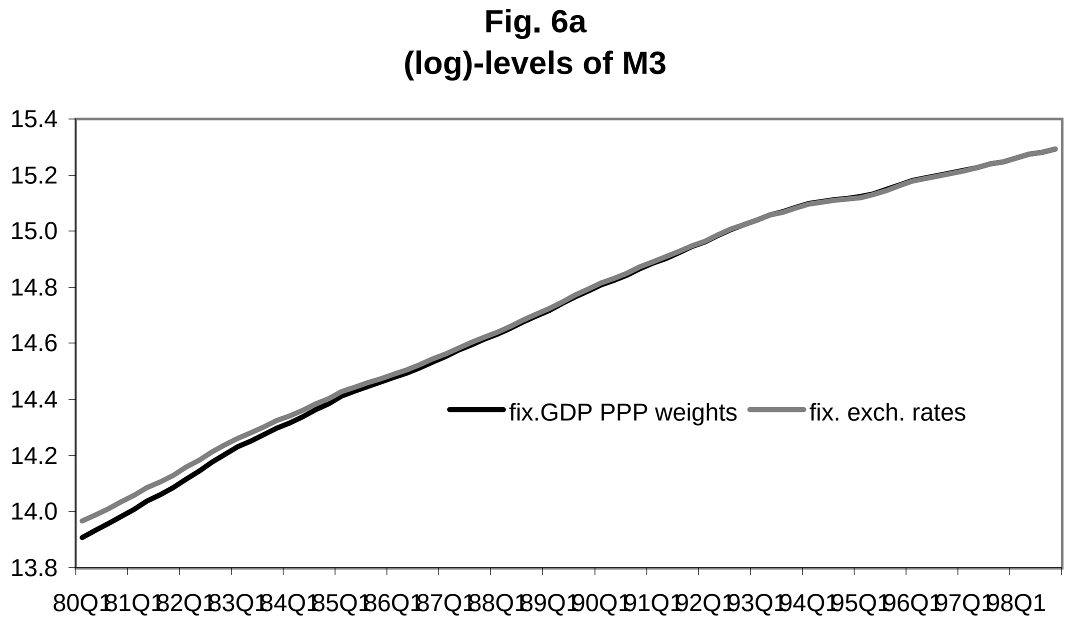


Fig. 6a
(log)-levels of M3
15.4
15.2
15.0
14.8
14.6
14.4
14.2
14.0
13.8
fix.GDP PPP weights
fix. exch. rates
80Q1
81Q1
82Q1
83Q1
84Q1
85Q1
86Q1
87Q1
88Q1
89Q1
90Q1
91Q1
92Q1
93Q1
94Q1
95Q1
96Q1
97Q1
98Q1

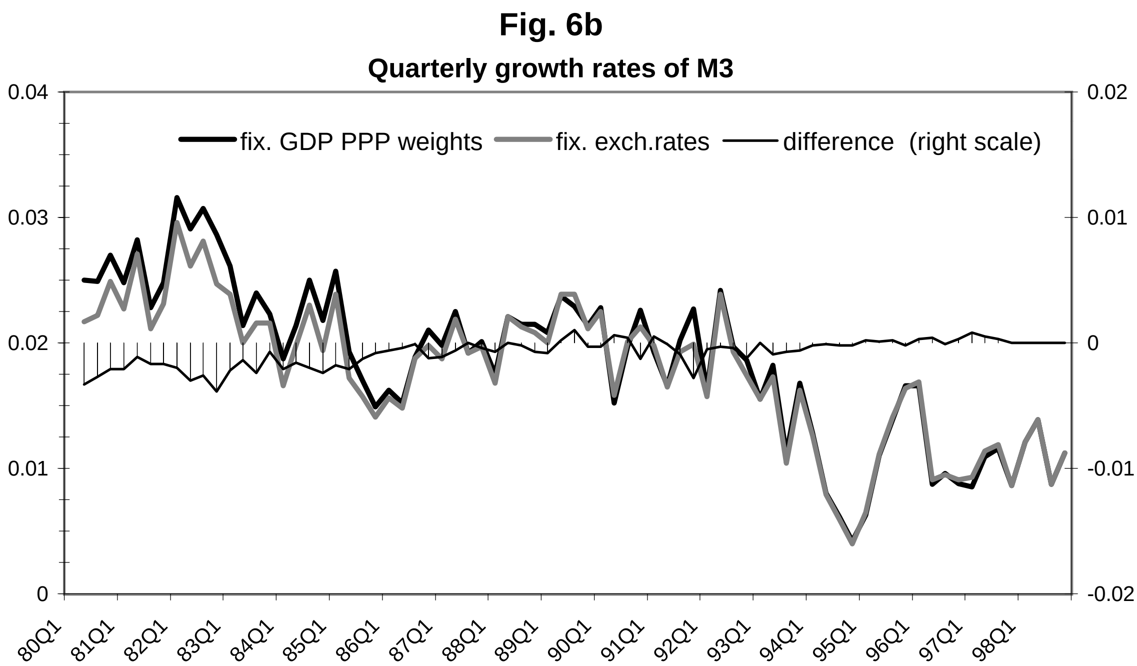


Fig. 6b
Quarterly growth rates of M3
0.04
0.03
0.02
0.01
0
0.02
0.01
0
-0.01
-0.02
fix. GDP PPP weights
fix. exch.rates
difference (right scale)
80Q1
81Q1
82Q1
83Q1
84Q1
85Q1
86Q1
87Q1
88Q1
89Q1
90Q1
91Q1
92Q1
93Q1
94Q1
95Q1
96Q1
97Q1
98Q1

Figure 7. Graphical evaluation and recursive estimates of equations (4b) and (5b)

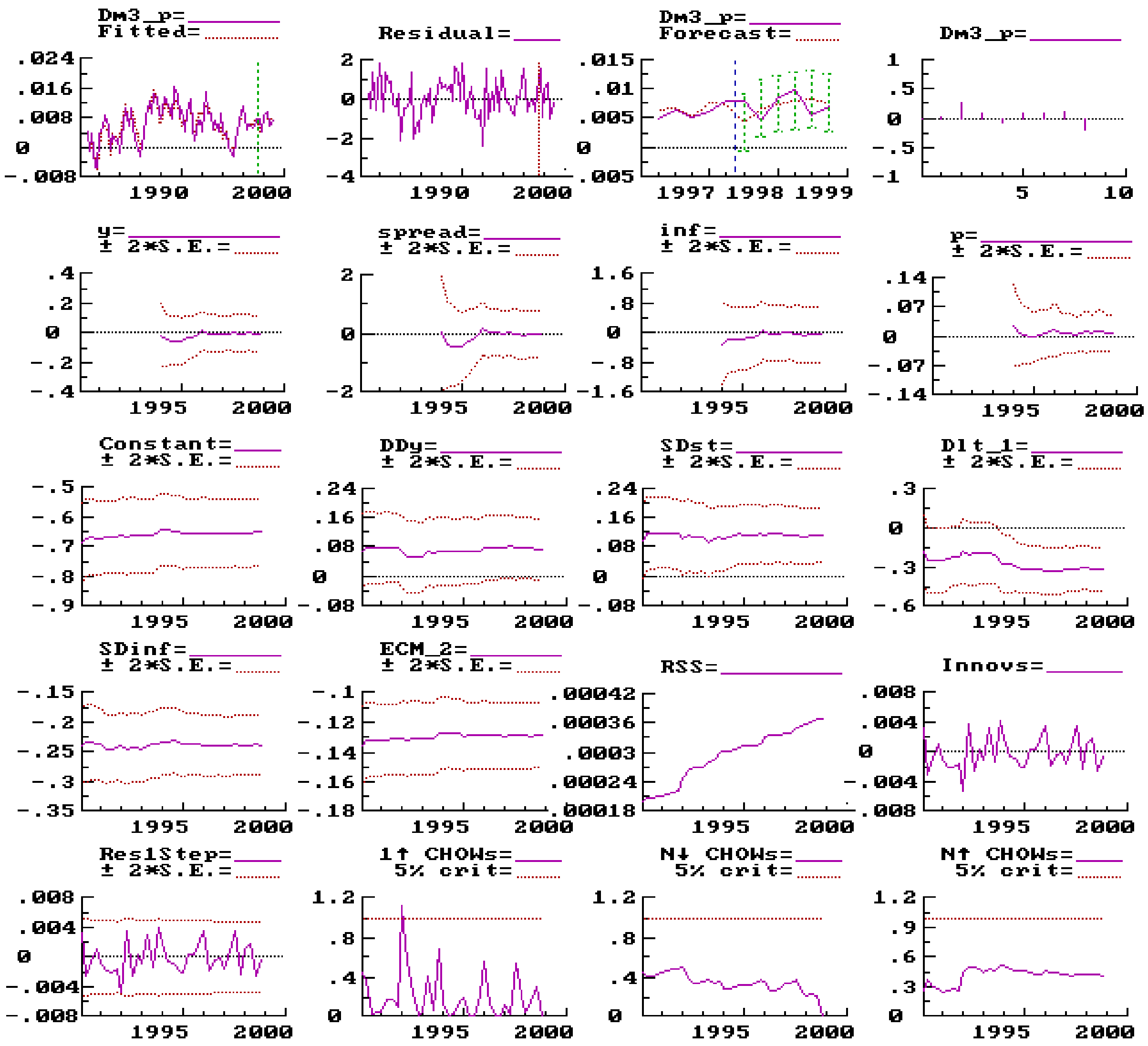


Figure 8. Impulse responses to a positive aggregate supply shock

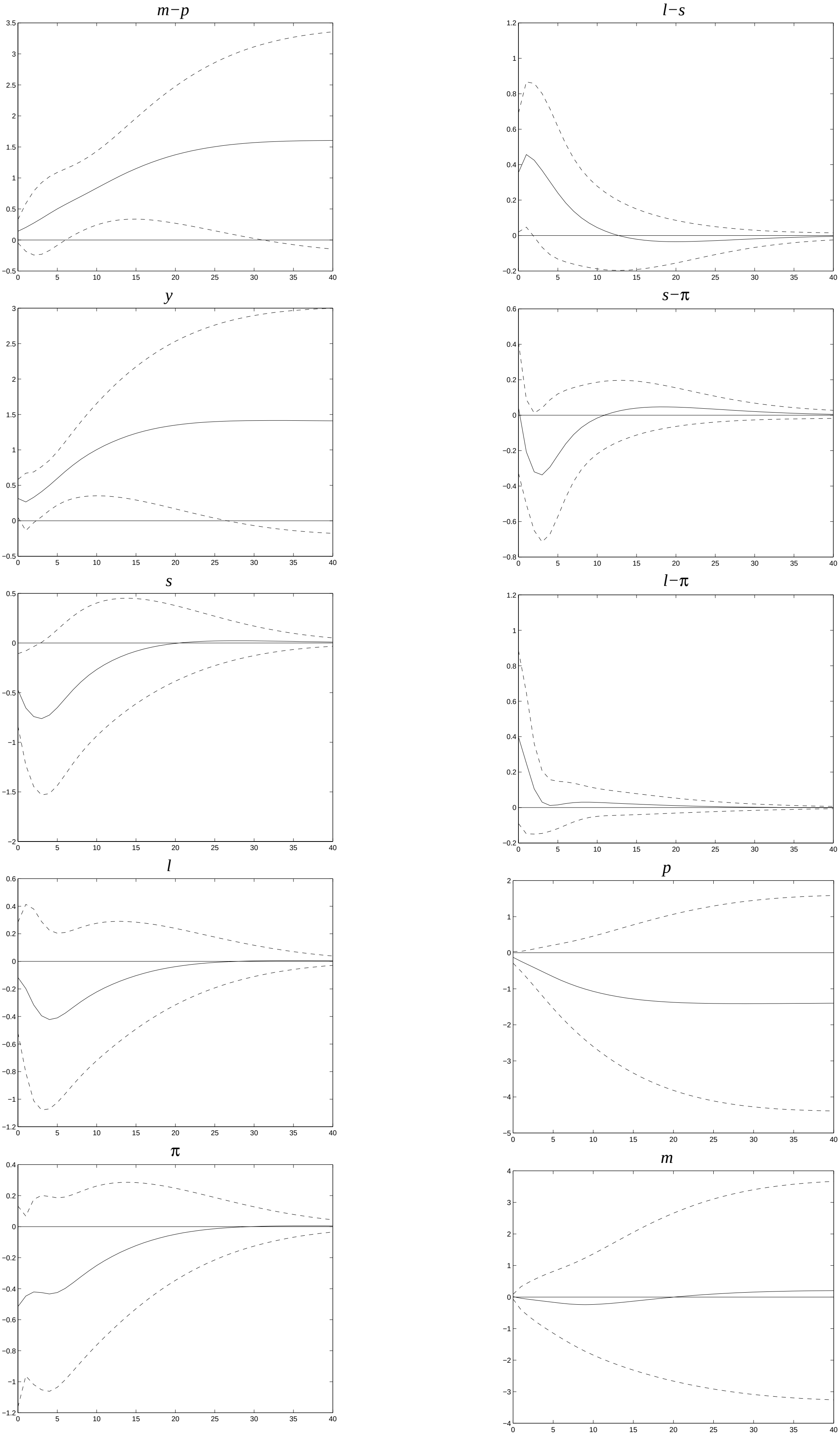


Figure 9. Impulse responses to a downward change in the monetary policy objective

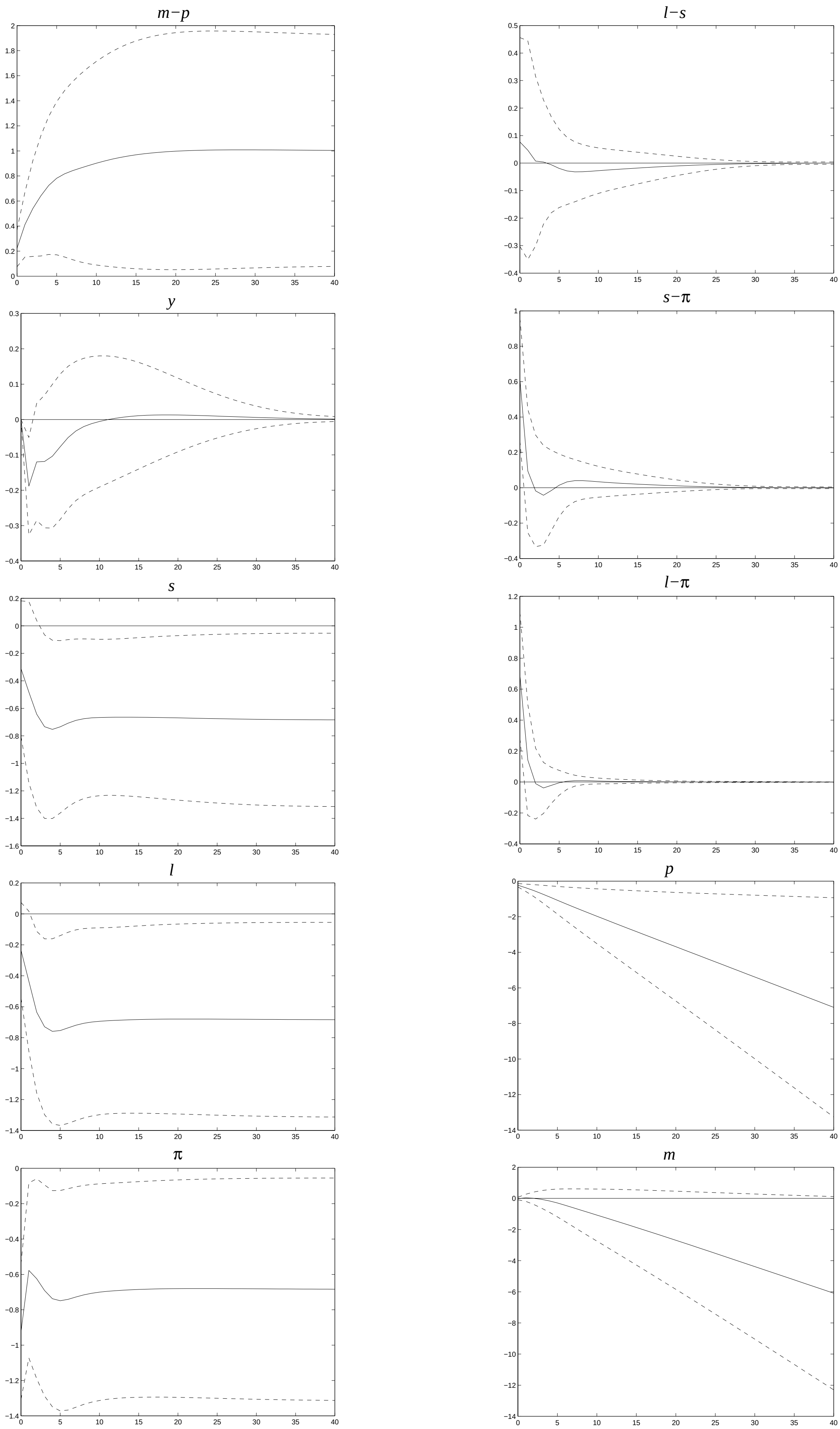


Figure 10. Impulse responses to a positive aggregate demand shock

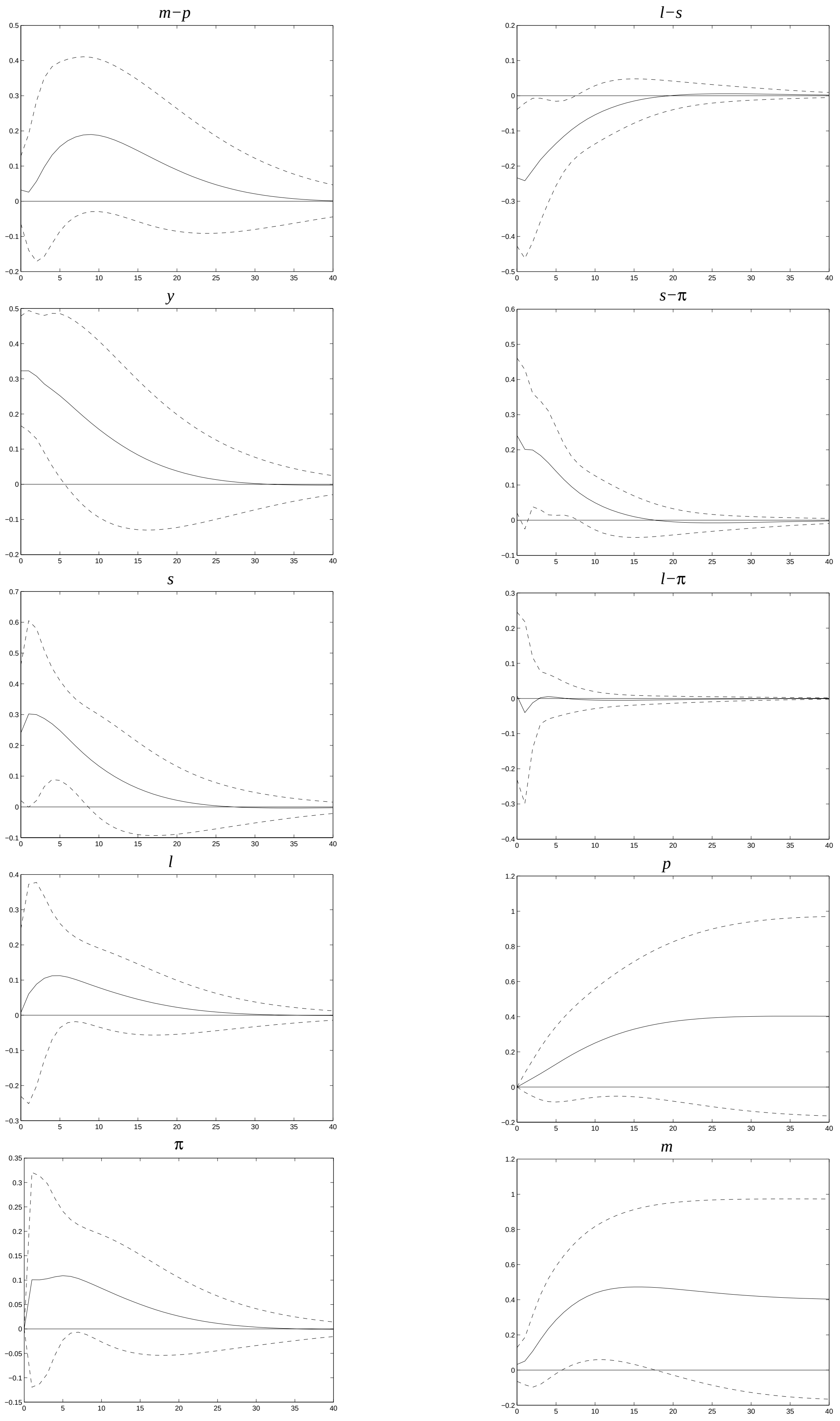


Figure 11. Impulse responses to a positive money demand shock

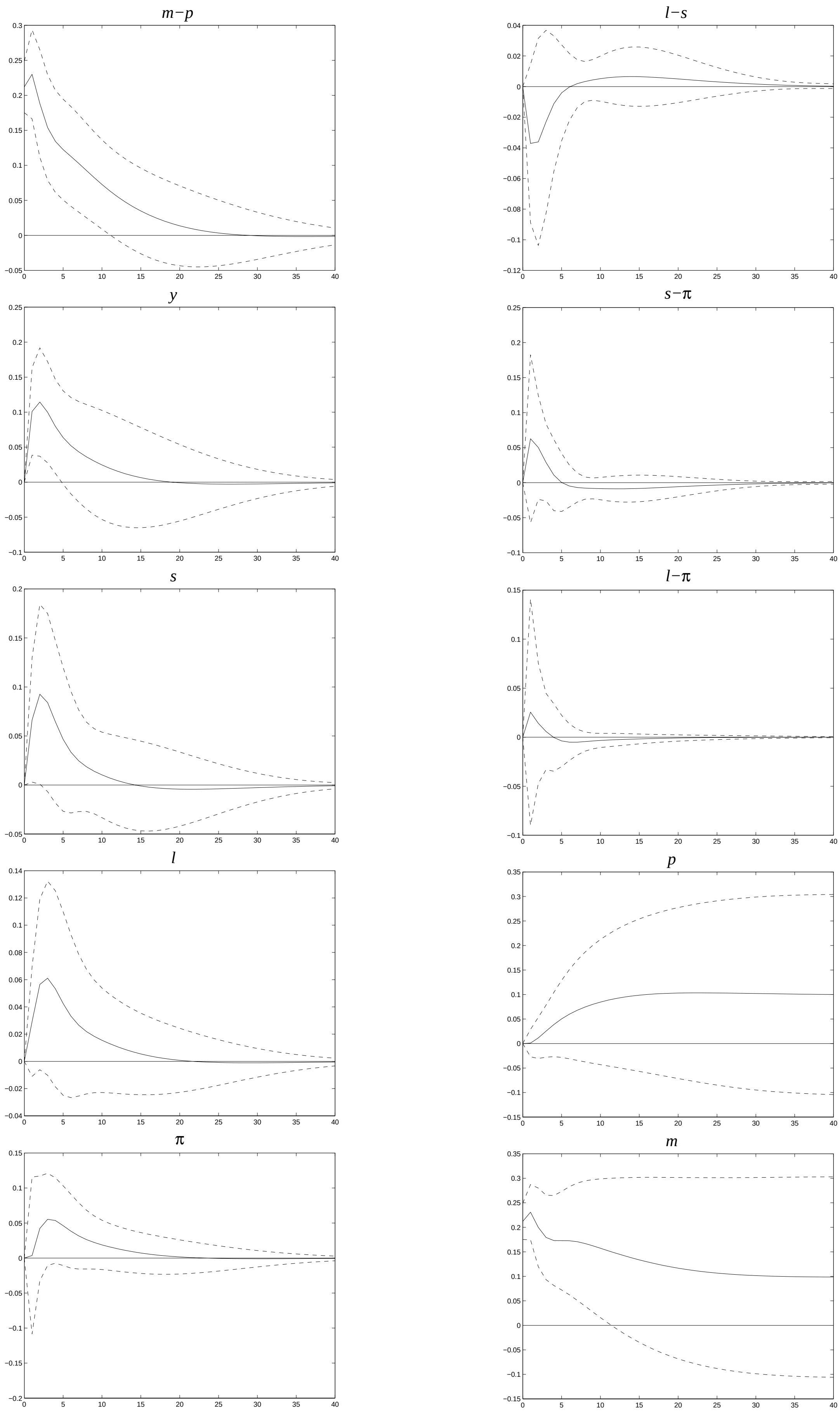
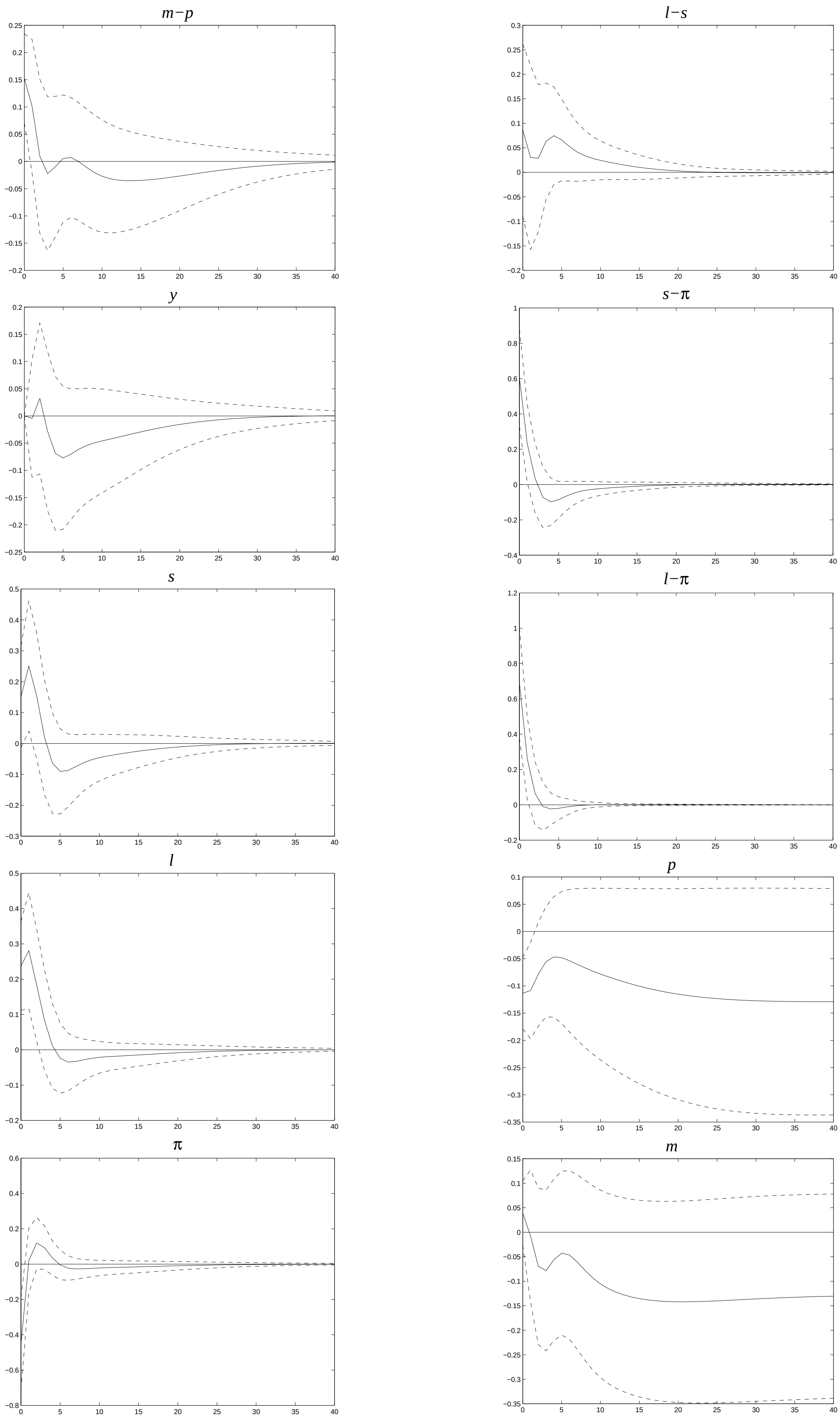


Figure 12. Impulse responses to a positive interest rate shock



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