Core Inflation in the Euro Area

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Abstract

Using a common trends model, we estimate a forward-looking "core" inflation measure for the Euro area based on long-run relations among major macroeconomic variables, bearing the interpretation of long-run inflation forecast. The proposed measure may be particularly suitable for the "two-pillar" monetary policy strategy of the ECB which focuses on medium-term inflation prospects.

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1. Introduction

Long-run price stability is widely recognized as the primary goal of monetary policy even in countries where the central bank does not adopt an explicit inflation-targeting strategy. This is the case of the recently born European Central Bank (ECB), whose monetary policy strategy is aimed at maintaining an annual inflation rate below 2% over a medium-term horizon (ECB 1999). Short-run fluctuations of the observed inflation rate, however, may be due to only temporary disturbances to which monetary policy should not respond. How to construct a reliable empirical measure of the underlying, long-run trend of inflation -"*core*" inflation- has therefore become a crucial issue in monetary policy design.

Several measures of the core inflation process have been proposed and implemented (see Wynne 1999 for a thorough overview). Some measures are derived from univariate statistical techniques, such as simple moving averages computed over a variable time span (from 3-6 up to 36 months) or more sophisticated methodologies (i.e. unobserved component models). Other measures are obtained from the cross-sectional distribution of individual price items, either by excluding from the price index some categories of goods (such as energy and food items) which are believed to be high-variance components, or by computing more efficient, "limited influence" estimators of the central tendency of the distribution, such as the (weighted) median popularized by Bryan and Cecchetti (1994) and Cecchetti (1997) for the US. Finally, Quah and Vahey (1995) applied to the UK a bivariate structural VAR approach to core inflation estimation based on long-run output neutrality of permanent shocks to the inflation rate.

In this paper we define and estimate a different, explicitly *forward-looking*, measure of core inflation, based on (appropriately estimated and tested) long-run relations among major macroeconomic variables. This measure may provide useful information in the light of the "two-pillar" monetary policy strategy of the ECB. In fact, in conducting monetary policy, the ECB considers: (*i*) the deviations of M3 growth from a reference value (a money growth indicator), and (*ii*) a broadly-based assessment of the outlook for future price developments in the Euro area as a whole (ECB 1999, 2000). This framework is motivated by the (alleged) close long-run relationship between money growth and inflation. Recent results have provided some evidence of stable long-run relationships among money, output, interest rates and inflation over the last two decades for the EMU countries (Brand and Cassola 2000, Gerlach and Svensson 2001, Golinelli and Pastorello 2000). We use such information to construct a forward-looking measure of core inflation consistent with the long-run features of the Euro area macroeconomy.

To this aim, we consider a multivariate framework, capturing the dynamic interactions among inflation, money, interest rates and output. The existence of long-run cointegrating relations among these variables allows us to use the common trends approach, outlined in Section 2, to decompose the inflation rate into a non-stationary (stochastic) trend component, capturing the effect of permanent shocks only, and a stationary transitory element. The former, "core", component bears the interpretation of the long-run inflation forecast conditional on an information set including several important macroeconomic variables. The main advantage of this common trends measure of core inflation lies in its *forward-looking* nature, capturing the long-term element of the inflation process (of particular interest from the monetary policy perspective) consistent with the long-run properties of the macroeconomic system.

This methodology is applied to the Euro area for the 1979-2000 period and results are discussed in Section 3. The message of this paper is that the ECB should take into proper account a forward-looking measure of the core inflation rate consistent with its whole monetary policy framework, based on a strong and stable long-run relationships between inflation and other major macroeconomic variables.

2. Methodology

Consider a vector $\mathbf{x}_t$ of n $I(1)$ variables of interest. If there exist $0 < r < n$ cointegrating relations among the variables, the following cointegrated VAR representation for $\mathbf{x}_t$ holds (deterministic terms are omitted for ease of exposition):

$$\Delta \mathbf{x}_t = \mathbf{\Pi}(L)\Delta \mathbf{x}_{t-1} + \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{2.1}$$

where the $n \times r$ matrix $\boldsymbol{\beta}$ contains the cointegrating vectors, such that $\boldsymbol{\beta}'\mathbf{x}_t$ are stationary linear combinations of the variables, and $\boldsymbol{\alpha}$ is the $n \times r$ matrix of factor loadings. $\boldsymbol{\varepsilon}_t$ is a vector of i.i.d. reduced form disturbances. As shown in Mellander, Vredin and Warne (1992), the cointegrated VAR in (2.1) can be inverted to yield the following stationary Wold representation for $\Delta \mathbf{x}_t$:

$$\Delta \mathbf{x}_t = \mathbf{C}(L)\boldsymbol{\varepsilon}_t \tag{2.2}$$

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + ...$ From the representation in (2.2) the following expression for the levels of the variables can be derived by recursive substitution:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{C}(1)\sum_{j=0}^{t-1}\boldsymbol{\varepsilon}_{t-j} + \mathbf{C}^*(L)\boldsymbol{\varepsilon}_t \tag{2.3}$$

where $\mathbf{C}^*(L) = \sum_{j=0}^{\infty}\mathbf{C}_j^* L^j$ with $\mathbf{C}_j^* = -\sum_{i=j+1}^{\infty}\mathbf{C}_i$. $\mathbf{C}(1)$ captures the long-run effect of the reduced form disturbances in $\boldsymbol{\varepsilon}$ on the variables in $\mathbf{x}$.

In order to obtain an economically meaningful interpretation of the dynamics of the variables of interest from the reduced form representations in (2.2) and (2.3), the vector of reduced form disturbances $\boldsymbol{\varepsilon}$ must be transformed into a vector of underlying, "structural" shocks, some of which with *permanent* effects on the level of $\mathbf{x}$ and some with only *transitory* effects. Let us denote this vector of i.i.d. structural disturbances as $\boldsymbol{\varphi}_t \equiv \begin{pmatrix} \boldsymbol{\psi}_t \\ \boldsymbol{\nu}_t \end{pmatrix}$, where $\boldsymbol{\psi}$ and $\boldsymbol{\nu}$ are subvectors of k and r elements respectively, with $k = n - r$. The structural form for the first difference of $\mathbf{x}_t$ is:

$$\Delta \mathbf{x}_t = \mathbf{\Gamma}(L)\boldsymbol{\varphi}_t \tag{2.4}$$

where $\mathbf{\Gamma}(L) = \mathbf{\Gamma}_0 + \mathbf{\Gamma}_1 L + ...$ The relationship between the reduced form and the structural shocks is given by:

$$\boldsymbol{\varepsilon}_t = \mathbf{\Gamma}_0 \boldsymbol{\varphi}_t$$

where $\mathbf{\Gamma}_0$ is an invertible matrix. Hence, comparison of (2.4) and (2.2) shows that

$$\mathbf{C}(L)\mathbf{\Gamma}_0 = \mathbf{\Gamma}(L)$$

implying that $\mathbf{C}_i \mathbf{\Gamma}_0 = \mathbf{\Gamma}_i$ $(\forall i > 0)$ and $\mathbf{C}(1)\mathbf{\Gamma}_0 = \mathbf{\Gamma}(1)$. In order to identify the elements of $\boldsymbol{\psi}_t$ as the permanent shocks and the elements of $\boldsymbol{\nu}_t$ as the transitory disturbances the following restriction on the long-run matrix $\mathbf{\Gamma}(1)$ must be imposed:

$$\mathbf{\Gamma}(1) = (\mathbf{\Gamma}_g \quad \mathbf{0}) \tag{2.5}$$

with $\mathbf{\Gamma}_g$ an $n \times k$ submatrix. The disturbances in $\boldsymbol{\psi}_t$ are then allowed to have long-run effects on (at least some of) the variables in $\mathbf{x}_t$, whereas the shocks in $\boldsymbol{\nu}_t$ are restricted to have only transitory effects.

From (2.4), the structural form representation for the endogenous variables in levels is derived as

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{\Gamma}(1) \sum_{j=0}^{t-1} \boldsymbol{\varphi}_{t-j} + \mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t = \mathbf{x}_0 + \mathbf{\Gamma}_g \sum_{j=0}^{t-1} \boldsymbol{\psi}_{t-j} + \mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t \tag{2.6}$$

where the partition of $\boldsymbol{\varphi}$ and the restriction in (2.5) have been used and $\mathbf{\Gamma}^*(L)$ is defined analogously to $\mathbf{C}^*(L)$ in (2.3). The permanent part in (2.6), $\sum_{j=0}^{t-1} \boldsymbol{\psi}_{t-j}$, may be expressed as a k-vector random walk with innovations $\boldsymbol{\psi}$:

$$\boldsymbol{\tau}_t = \boldsymbol{\tau}_{t-1} + \boldsymbol{\psi}_t = \boldsymbol{\tau}_0 + \sum_{j=0}^{t-1} \boldsymbol{\psi}_{t-j} \tag{2.7}$$

Using (2.7) in (2.6) we finally obtain the common trends representation for $\mathbf{x}_t$:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{\Gamma}_g \boldsymbol{\tau}_t + \mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t \tag{2.8}$$

As shown in detail by Stock and Watson (1988), King, Plosser, Stock and Watson (1991) and Warne (1993), the identification of separate permanent shocks requires a sufficient number of restrictions on the long-run impact matrix $\mathbf{\Gamma}_g$ in (2.8). Part of these restrictions are provided by the cointegrating relations and the consistent estimate of matrix $\mathbf{C}(1)$; additional ones are suggested by economic theory. Finally, having estimated $\mathbf{\Gamma}_g$, the behavior of the variables in $\mathbf{x}_t$ due to the permanent disturbances only, interpreted as the long-run forecast of $\mathbf{x}_t$, may be computed as $\mathbf{x}_0 + \mathbf{\Gamma}_g \boldsymbol{\tau}_t$.

3. Results

In the empirical analysis we consider a small-scale quarterly macroeconomic model for the 11 countries of the European Monetary Union including the following variables: the log of real money balances obtained by deflating the stock of M3 with the consumer price level $(m-p)$, the log of real GDP (y), the nominal long-term interest rate on government bonds (l), the quarterly rate of consumer price inflation measured by the Harmonized Index of Consumer Prices (HICP) used by the ECB (π), and the rate of capacity utilization in the manufacturing sector measured by the OECD (qr) as an indicator of the business cycle. The sample spans the post-EMS period, from 1979(2) to 2000(2). For the pre-Euro period (up to 1998(4)) aggregate variables for the Euro area were constructed by aggregating the historical data of the 11 member countries.[1] This approach, recently followed by Gerlach and Svensson (2001) and Galì, Gertler and Lopez-Salido (2000), rests on the hypothesis that the artificial Euro-area data are still suitable for analyzing and forecasting EMU-wide aggregates (Bodo, Golinelli and Parigi (2000) provide evidence in favor of this view for industrial production).

Standard ADF and Phillips-Perron unit-root tests show that all variables can be treated as $I(1)$ processes, with the exception of the stationary output gap measure qr. The vector of endogenous variables is then specified as $\mathbf{x}_t = (y_t \ l_t \ m_t - p_t \ \pi_t \ qr_t)'$. A $VAR(3)$ system for $\mathbf{x}_t$ was estimated with unrestricted constant term and seasonal dummies. The choice of order three is supported by the Akaike information criterion (AIC) and the inclusion of a linear trend in the system was not significant. Diagnostic tests on each individual equation and on the system as a whole do not detect any sign of residual autocorrelation, non-normality and heteroscedasticity, supporting our dynamic specification of the system. Moreover, recursive estimation shows coefficient stability over time and the absence of structural breaks.

The existence of long-run relationships among the variables is tested by means

[1] The data used are downloadable at: http://www.spbo.unibo.it/pais/golinelli/macro.html

of the Johansen's (1988) cointegration trace test statistic reported in Table 1. As shown in the upper panel of the table, there is strong evidence of three cointegrating vectors; we then proceed in the analysis with $r = 3$. To identify the long-run relationships we impose and test a set of restrictions on the cointegrating vector parameters (in matrix $\boldsymbol{\beta}$) suggested both by economic theory and by the available evidence. We consider a relation between real money balances and income (as found by Gerlach and Svensson 2001), interpretable as a simple long-run money demand function, and a relation between the nominal interest rate and the inflation rate. Moreover, the stationarity of the output gap provides additional identifying restrictions for the third cointegrating vector. The estimated parameters of the restricted cointegrating vectors are reported in the lower panel of Table 1 together with the estimated loading factors (elements of $\boldsymbol{\alpha}$) and the test for the overidentifying restrictions, which supports the chosen identification assumptions (the p-value of the test is 0.14). The income coefficient in the money demand equation (1.57) is very precisely estimated (and stable in recursive analysis) and very similar to Gerlach and Svensson's (2001) estimate. The long-run effect of inflation on the long term interest rate is somewhat higher (1.32) than the unitary value suggested by a simple Fisher relation. This result, also found by Brand and Cassola (2000), may suggest that the level of inflation affects the real interest rate in the long-run through an "inflation premium" positively related to the level of π. In fact, the gradual disinflation process occurred in many European countries over the 1990s has been accompanied by a progressive decrease in long-term real interest rates. The estimated loading parameters show that positive deviations from the equilibrium relation between $m-p$ and y cause a strong upward pressure on inflation and an error-correcting reaction of real money balances. The interest rate shows a similar error-correcting behavior in response to positive deviations from the long-run relation with inflation. Finally, increases in the capacity utilization rate have a positive impact on inflation short-run dynamics.

In the common trends framework the existence of three cointegrating relationships among our five variables implies the presence of two distinct sources of shocks having permanent effects on at least some of the variables in $\mathbf{x}$. As previously mentioned, such (restricted) cointegrating vectors are used to identify the elements of $\boldsymbol{\Gamma}_g$ in (2.8) together with additional restrictions grounded on economic theory. In order to achieve identification of the common trends model, we then make the following assumptions on the nature of the two permanent shocks in the system: we consider a *real* shock (ψ_r) and a *nominal* disturbance (ψ_n). The permanent part (2.7) of the common trends representation is then the following bivariate random walk:

$$\begin{pmatrix} \tau_r \\ \tau_n \end{pmatrix}_t = \begin{pmatrix} \mu_r \\ \mu_n \end{pmatrix} + \begin{pmatrix} \tau_r \\ \tau_n \end{pmatrix}_{t-1} + \begin{pmatrix} \psi_r \\ \psi_n \end{pmatrix}_t \tag{3.1}$$

where $\boldsymbol{\mu}$ is a vector of constant drift terms. As additional restrictions we assume that output is not affected in the long-run by the nominal shock (a long-run neutrality assumption), and that the output gap is a stationary variable, therefore not affected by permanent shocks in the long-run. Letting γ_{ij} denote the generic element of $\boldsymbol{\Gamma}_g$, the two assumptions above imply $\gamma_{12} = 0$ and $\gamma_{51} = \gamma_{52} = 0$ respectively. The common trends representation of the variables in levels (2.8) becomes therefore the following:

$$\begin{pmatrix} y \\ r \\ m-p \\ \pi \\ qr \end{pmatrix}_t = \begin{pmatrix} y \\ r \\ m-p \\ \pi \\ qr \end{pmatrix}_0 + \begin{pmatrix} \gamma_{11} & 0 \\ \gamma_{21} & \gamma_{22} \\ \gamma_{31} & \gamma_{32} \\ \gamma_{41} & \gamma_{42} \\ 0 & 0 \end{pmatrix} \begin{pmatrix} \tau_r \\ \tau_n \end{pmatrix}_t + \boldsymbol{\Gamma}^*(L) \begin{pmatrix} \psi_r \\ \psi_n \\ \nu_1 \\ \nu_2 \\ \nu_3 \end{pmatrix}_t \tag{3.2}$$

where the ν_i's $(i = 1, 2, 3)$ are purely transitory disturbances (uncorrelated with the permanent shocks) to which, given the main focus of our analysis, we do not attribute any structural economic interpretation. The estimated elements of the long-run impact matrix $\boldsymbol{\Gamma}_g$ (not reported) show that the real shock (ψ_r), which is the only determinant of the long-run behavior of real money balances and output, plays only a marginal role in explaining the long-run features of interest rate and inflation, that are dominated by nominal disturbances (ψ_n). Our measure of "core" inflation, derived from the common trends model, is then computed as $\hat{\pi}_t^c = \pi_0 + \hat{\gamma}_{41}\hat{\tau}_{r,t} + \hat{\gamma}_{42}\hat{\tau}_{n,t}$. The forecast error variance decomposition reported in Table 2 shows that, while in the short-run the bulk of inflation variability is explained by transitory shocks (83% at the one-quarter and 56% at the one-year horizons), at the 2-3 year horizon permanent shocks play the dominant role, with the nominal shock accounting for about 70% of inflation variability at the three-year horizon. The common trend measure of core inflation, reflecting the effect of permanent disturbances only, adequately captures the medium/long term inflation dynamics that the policy maker should aim to stabilize.

Figure 1 compares the estimated core inflation series with the measured consumer price inflation, both expressed as annual rates (four-quarter lagged moving averages). In the 1980s the core inflation rate shows more limited fluctuations, ranging from 3.5% to 6.5%, with respect to actual inflation, which varies widely between 2% and 10%. In particular, core inflation displays a lower peak during the oil-shock episode of the early '80s (around 6.5% against a 10% actual inflation rate), whereas this pattern is reversed during the counter-shock in the mid-'80s. Starting in the early 1990s, the two inflation rates show a more similar behavior, though with some notable exceptions, namely in 1991-1992, when the core rate decreased in the face of an increasing actual inflation, and again in 1993, when the decline in the core rate occurred with actual inflation broadly constant.

Of particular interest is the relative behavior of the two inflation series since the introduction of the euro in January 1999: both rates increased from around 1.5% in mid-1999 up to a 2-2.5% range in 2000(2). Such an increase is largely attributed to the sharp rise in oil prices, as shown by the consumer price inflation rate "excluding food and energy items" (also portrayed in Figure 1), which was stable around 1.2%. However, our forward-looking measure of core inflation signals that the long-run inflation forecast as of 2000(2) was very close (in fact, a little higher) than observed inflation, even though other commonly used indicators of inflationary pressures, such as the "ex food and energy" index, showed a lower and stable inflation rate. This evidence can lend some support to the prudent monetary policy attitude of the ECB in 1999 and 2000 in the management of policy interest rates.

Of course, a core inflation rate estimated from a common trend model depends on the specification of the system in terms of variables included, sample period, dynamic specification, and other modelling choices. However, the core inflation series obtained from the small-scale macroeconomic model used in this paper, featuring long-run relationships between real money balances, output, inflation and interest rates, seems an useful benchmark to evaluate the properties of other measures of core inflation currently used in the monetary policy debate.

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Table 1

Cointegration tests: 1979(2)-2000(2)						
Eigenv.	Null	Altern.	Trace stat.	10% c.v.	5% c.v.	1% c.v.
0.4017	$r = 0$	$r \geq 1$	107.5		68.5	76.1
0.3266	r □ 1	$r \geq 2$	63.8	44.0	47.2	54.5
0.2130	r □ 2	$r \geq 3$	30.2	26.8	29.7	35.6
0.1029	r □ 3	$r \geq 4$	9.9	13.3	15.4	20.0
0.0074	r □ 4	$r = 5$	0.6	2.7	3.8	6.7

Cointegration parameter estimates								
	Loading coeff. ($\boldsymbol{\alpha}$)			Restricted cointegrating vectors ($\boldsymbol{\beta}'$)				
				y	l	$m-p$	π	qr
y	0.0952 (0.0503)	0.0203 (0.0483)	0.0368 (0.0427)					
l	0.0737 (0.0319)	-0.0811 (0.0307)	0.0400 (0.0271)	-1.569 (0.016)	0	1	0	0
$m-p$	-0.0944 (0.0437)	0.0399 (0.0420)	-0.0017 (0.0371)	0	1	0	-1.324 (0.139)	0
π	0.3622 (0.0834)	0.1343 (0.0802)	0.2882 (0.0708)	0	0	0	0	1
qr	0.0172 (0.0455)	-0.0282 (0.0437)	-0.1529 (0.0386)					
Overidentifying restrictions test: $\chi^2(4) = 6.98$ [0.14]								

Table 2

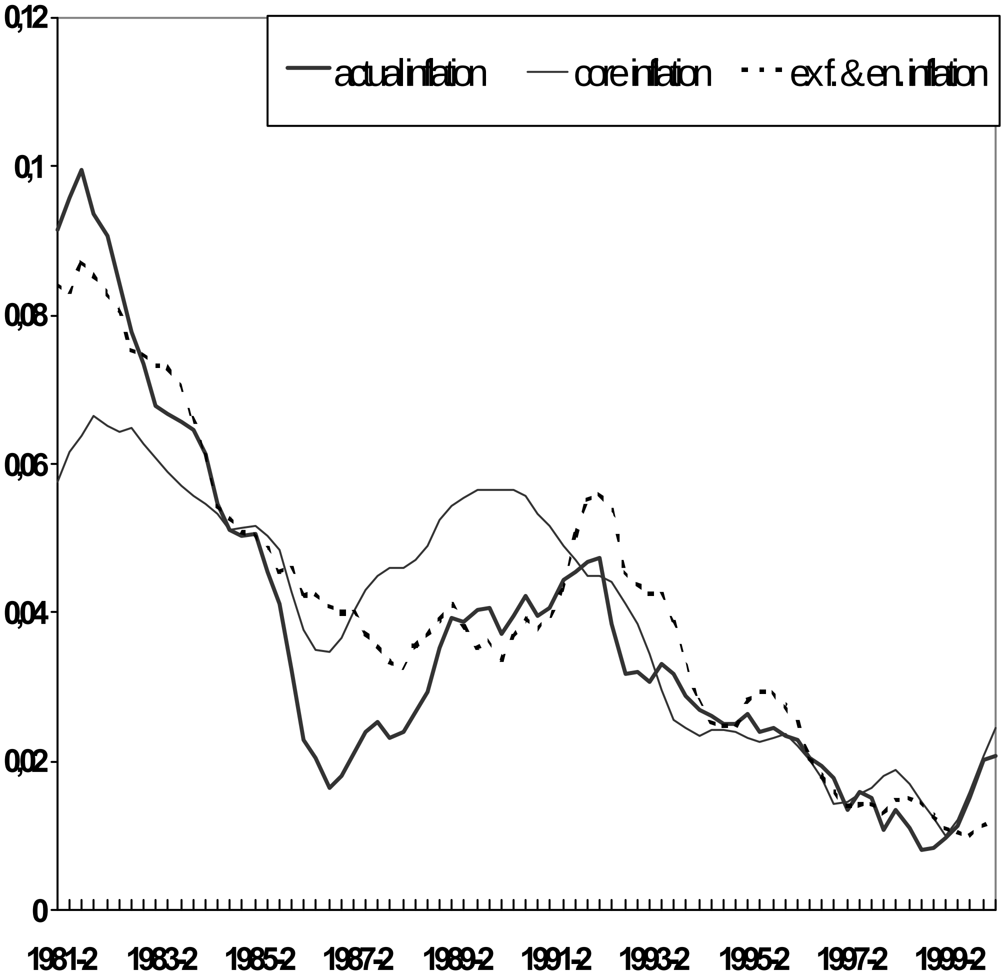
Forecast error variance decomposition of the inflation rate

	Fraction of variance accounted by:			
Horizon (quarters)	ψ_r	ψ_n	$\psi_{r,n}$	$\nu_{1,2,3}$
1	0.089	0.083	0.172	0.828
2	0.070	0.283	0.353	0.647
3	0.065	0.341	0.406	0.594
4	0.084	0.359	0.443	0.557
8	0.055	0.585	0.640	0.360
12	0.051	0.698	0.749	0.251
16	0.055	0.750	0.805	0.195
20	0.058	0.778	0.836	0.164
∞	0.085	0.915	1	0

ψ_r and $\psi_n$denote the real and nominal permanent shocks respectively; $\psi_{n,r}$ denotes the joint effect of permanent shocks; $\nu_{1,2,3}$ denotes the joint effect of transitory shocks.

Figure 1

Measured consumer price inflation rate, estimated core inflation rate and "ex food and energy" consumer price inflation rate (annual rates)



Appendix

This appendix reports additional results on the specification and stability analysis of the $VAR(3)$ system in (2.1) and on the estimated long-run effects of permanent disturbances in the common trends model (3.2).

Table A1: Diagnostic tests on the unrestricted VAR(3) system (p-values reported)

	Equation for: y	l	$m-p$	π	qr	Whole system	
AR(5) $F(5,61)$	0.64	0.16	0.08	0.15	0.10	AR(5) $F(125,187)$	0.17
Normality $\chi^2(2)$	0.06	0.99	0.43	0.97	0.06	Normality $\chi^2(10)$	0.28
Heterosc. $F(30,35)$	0.68	0.33	0.59	0.82	0.99	Heterosc. $F(450,340)$	1.00
ARCH(4) $F(4,58)$	0.74	0.13	0.91	0.77	0.30		

Figure A1: One-step residuals from recursive estimation of the VAR(3) system (1991(1)-2000(2)) with 95% confidence intervals and system break-point Chow stability tests

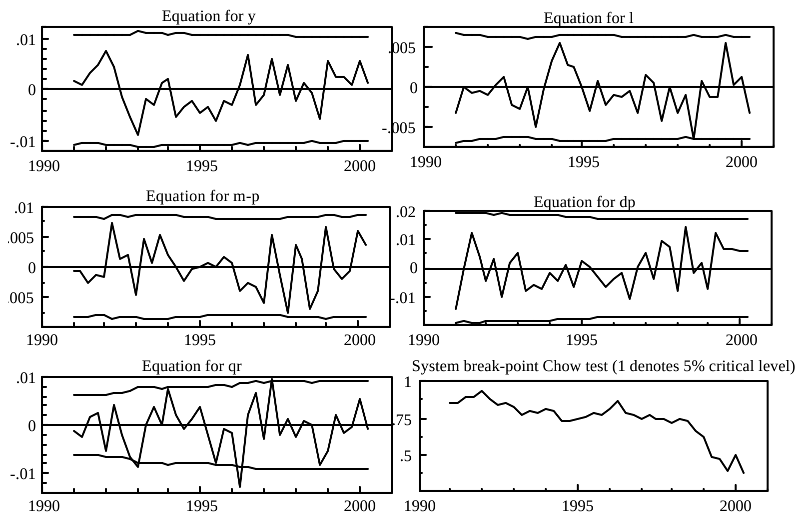


Table A2: Long-run effects of permanent shocks in the estimated common trends model

Elements of matrix $\mathbf{\Gamma}_g$ in (3.2)
(asymptotic standard errors in parentheses)

Variable	Shock	
	ψ_r	ψ_n
y	0.0062 (0.0027)	0
l	0.0016 (0.0030)	0.0055 (0.0016)
$m-p$	0.0097 (0.0042)	0
π	0.0012 (0.0023)	0.0041 (0.0012)
qr	0	0