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Measuring US core inflation: A common trends approach

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Abstract

In this paper the long-run trend in CPI inflation (*core inflation*) for the US over the 1960–2000 period is estimated using a common trends model. In this framework, core inflation is interpreted and constructed as the long-run forecast of inflation conditional on the information contained in nominal money growth, output fluctuations and movements in the oil price. Unlike other commonly used measures of core inflation, the common-trends core inflation rate exploits the long-run link between inflation and monetary growth, a strong feature of the data.

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1. Introduction

In the recent debate about monetary policy targets, one prominent view favours the direct formulation of the central bank's objective in terms of the ultimate policy goal, price stability. Inflation targeting policies, setting precise quantitative targets for monetary authorities, have been advocated and implemented in several countries (see Bernanke et al. (1999) for a cross-country assessment and a review of the main implementation issues). Though the essence of inflation-targeting policies can be simply stated, qualifications are needed once it is recognized that observed inflation may

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fluctuate in the short run due to only temporary disturbances of both real and nominal nature. As argued by Cecchetti (1997), transitory phenomena should not affect policymakers' actions and short-run changes in the observed inflation rate should be carefully analyzed in order to extract the long run, trend component of inflation, commonly referred to as the "underlying" or "*core*" inflation rate. Accordingly, the empirical study of inflation has become a crucial issue in monetary policy analysis, with the aim of distinguishing persistent sources of inflationary pressures from only transient price fluctuations.

Several measures of the core inflation rate have been put forward and used in practical monetary policy conduct (see the collection of papers by central banks' model builders published by the Bank for International Settlements, 1999). One approach relies on the use of limited influence estimators, such as trimmed means or the (weighted) median, instead of the conventional weighted mean calculated over the complete cross-sectional distribution of the individual price components (Bryan and Cecchetti, 1994). Other approaches apply various techniques to the aggregate price change series to measure the core inflation component. For example, univariate techniques, such as simple moving averages calculated over a variable time span (from 3 to 6 up to 36 months) or more sophisticated methodologies (e.g. unobserved component models), are used to eliminate the noise component of inflation fluctuations. Other measures are based on econometric methods aimed at decomposing economic time-series into permanent and transitory components. In particular, Quah and Vahey (1995) applied to the UK a bivariate structural vector autoregression (VAR) approach to core inflation estimation based on the assumption of long-run output neutrality of permanent shocks to the inflation rate.

The present paper extends the bivariate, output-inflation setting of Quah and Vahey (1995) to a multivariate framework applied to the US inflation from 1960 to 2000. In this context, we interpret US *core* inflation as the long-term inflation forecast obtained from a small-scale common trends model (Stock and Watson, 1988; King et al., 1991), built around a long-run equilibrium relation (appropriately tested) between the inflation rate and what is believed to be its main long-run determinant, the rate of nominal money growth. [1] In so doing, we follow the lead of Bryan and Cecchetti (1994), who define core inflation as the long-run, persistent, component of the measured inflation rate, "which is tied in some way to monetary growth" (p. 197). Also Quah and Vahey (1995) argue that it would be informative to allow for more—in particular monetary—variables in the VAR system used to estimate core inflation. Hallman et al. (1991) have already provided evidence of a strong long-run link between growth in $M2$ and inflation in the US since the early '50s: we interpret and test this relationship in terms of cointegration within a system including also an output measure and the price of oil as an additional major source of inflationary pressures. In this framework, identification of permanent shocks is

[1] Also Blix (1995) uses a common trends framework to implement the Quah–Vahey long-run identification scheme for several countries, including the US.

achieved and a measure of inflation is constructed which reflects only the effect of permanent disturbances.

The rest of the paper is organised as follows. The common trends approach to core inflation estimation is outlined and implemented in Section 2. The resulting common trends measure of core inflation is discussed in Section 3 and compared with two commonly used core inflation series (the Bryan–Cecchetti median inflation and the CPI excluding food and energy). Section 4 briefly concludes.

2. The common trends approach to core inflation estimation

The econometric literature on the decomposition of economic time series into permanent and transitory components provides the necessary empirical framework for the estimation of a long-run inflation forecast (see Quah (1992) for a general treatment of this issue). Starting with the seminal work of Beveridge and Nelson (1981), different approaches to the permanent–transitory decomposition have been proposed. Blanchard and Quah (1989) have shown how a trend-cycle decomposition may be attained for non-cointegrated $I(1)$ variables by constraining their long-run responses to different shocks obtained from the VAR representation. Quah and Vahey (1995) applied this methodology to obtain an estimate of the core inflation component from a VAR model including only industrial production and inflation. In their framework, core inflation is defined as that component of the observed inflation rate that has no long-run effect on output, consistent with a vertical long-run Phillips curve relationship between output and inflation.

In the present paper we construct an estimate of core inflation on the basis of a larger information set, including other macroeconomic variables that can play an important role in determining the long-run inflation rate. The long-run (cointegration) properties of the data may then be used to disentangle the short- and long-run components of the variables analyzed, as shown by Stock and Watson (1988) and Gonzalo and Granger (1995). To this aim, we apply the common trends methodology of King et al. (1991) and Mellander et al. (1992) to a small-scale macroeconomic system including the inflation rate, output, the growth rate of the nominal money stock and the oil price (as a source of supply-side inflation disturbances). In this context, core inflation is interpreted as the long-run forecast of the inflation rate conditional on the information contained in the variables of the system and consistent with the long-run cointegration properties of the data. [2] A similar definition of core inflation is adopted by Cogley and Sargent (2000) in their analysis of the dynamic behavior of post-war US inflation. Moreover, in a multivariate system, structural shocks are likely to be identified more precisely than in the bivariate approach of Quah and Vahey (1995), and the forecast error variance decomposition can yield meaningful

[2] Evans and Reichlin (1994) show that the Quah–Vahey approach yields a measure of core inflation that cannot bear the interpretation of a long-run forecast for the inflation series.

information about the dynamic effects of different disturbances on the inflation process. [3] The rest of this section outlines and applies the econometric methodology.

2.1. 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):

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

where $\mathbf{\Pi}(L) = \mathbf{\Pi}_1 + \mathbf{\Pi}_2 L + \cdots + \mathbf{\Pi}_p L^{p-1}$ is a polynomial in the lag operator L, 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, $\boldsymbol{\alpha}$ is the $n \times r$ matrix of factor loadings, and $\boldsymbol{\varepsilon}_t$ is a vector of i.i.d., serially uncorrelated, reduced form disturbances. As shown in Mellander et al. (1992), the cointegrated VAR in (1) can be inverted to yield the following stationary Wold representation for $\Delta\mathbf{x}_t$:

$$\Delta\mathbf{x}_t = \mathbf{C}(L)\boldsymbol{\varepsilon}_t, \tag{2}$$

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + \cdots$ with $\sum_{j=0}^{\infty} j|\mathbf{C}_j| < \infty$. From the representation in (2) the following expression for the levels of the variables can be derived by recursive substitution:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{C}(1)\sum_{j=0}^{t-1}\boldsymbol{\varepsilon}_{t-j} + \mathbf{C}^*(L)\boldsymbol{\varepsilon}_t, \tag{3}$$

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 $\mathbf{x}$ and $\mathbf{x}_0$ is the initial observation in the sample.

In order to obtain an economically meaningful interpretation of the dynamics of the variables of interest from the reduced form representations in (2) and (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 $\boldsymbol{\varphi}_t \equiv (\boldsymbol{\psi}_t \mathbf{v}_t)'$, where $\boldsymbol{\psi}$ and $\mathbf{v}$ are subvectors of k and r elements respectively, with $k = n - r$. The structural form for the first difference of $\mathbf{x}_t$ is

$$\Delta\mathbf{x}_t = \mathbf{\Gamma}(L)\boldsymbol{\varphi}_t, \tag{4}$$

where $\mathbf{\Gamma}(L) = \mathbf{\Gamma}_0 + \mathbf{\Gamma}_1 L + \cdots$ Since the first element of $\mathbf{C}(L)$ in (2) is $\mathbf{I}$, equating the first term of the right-hand sides of (2) and (4) yields the following relationship between the reduced form and the structural shocks:

$$\boldsymbol{\varepsilon}_t = \mathbf{\Gamma}_0\boldsymbol{\varphi}_t, \tag{5}$$

[3] Quah (1995) clarifies that, though both the Quah–Vahey model and the common trends model use long-run restrictions to achieve identification, in general they do not provide equivalent representations of the data. Crowder (1995) discusses the special case in which the two representations are equivalent.

where $\boldsymbol{\Gamma}_0$ is an invertible matrix. Hence, comparison of (4) and (2) shows that

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

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

$$\boldsymbol{\Gamma}(1) = (\boldsymbol{\Gamma}_g\mathbf{0}) \tag{6}$$

with $\boldsymbol{\Gamma}_g$ 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 $\mathbf{v}_t$ are restricted to have only transitory effects.

From (4), the structural form representation for the endogenous variables in levels is derived as

$$\mathbf{x}_t = \mathbf{x}_0 + \boldsymbol{\Gamma}(1)\sum_{j=0}^{t-1}\boldsymbol{\varphi}_{t-j} + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t = \mathbf{x}_0 + \boldsymbol{\Gamma}_g\sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j} + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t, \tag{7}$$

where the partition of $\boldsymbol{\varphi}$ and the restriction in (6) have been used and $\boldsymbol{\Gamma}^*(L)$ is defined analogously to $\mathbf{C}^*(L)$ in (3). The permanent part in (7), $\sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j}$, may be expressed as a k-vector random walk $\boldsymbol{\tau}$ with innovations $\boldsymbol{\psi}$:

$$\boldsymbol{\tau}_t = \boldsymbol{\tau}_{t-1} + \boldsymbol{\psi}_t = \boldsymbol{\tau}_0 + \sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j}. \tag{8}$$

Using (8) in (7) we finally obtain the common trend representation for $\mathbf{x}_t$:

$$\mathbf{x}_t = \mathbf{x}_0 + \boldsymbol{\Gamma}_g\boldsymbol{\tau}_t + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t. \tag{9}$$

As shown in detail by Stock and Watson (1988), King et al. (1991) and Warne (1993), the identification of separate permanent shocks requires a sufficient number of restrictions on the long-run impact matrix $\boldsymbol{\Gamma}_g$ in (9). Part of these restrictions are provided by the cointegrating relations and the consistent estimation of $\mathbf{C}(1)$; additional ones are suggested by economic theory (e.g. long-run neutrality assumptions). Finally, having estimated $\boldsymbol{\Gamma}_g$, 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 $\mathbf{x}_0 + \boldsymbol{\Gamma}_g\boldsymbol{\tau}_t$. Formally, such long-run forecast can be expressed as

$$\lim_{h\to\infty} E_t\mathbf{x}_{t+h} = \mathbf{x}_0 + \boldsymbol{\Gamma}_g\boldsymbol{\tau}_t, \tag{10}$$

capturing the values to which the series are expected to converge once the effect of the transitory shocks have died out (Cogley and Sargent, 2000). Moreover, from the moving average representation in (4), impulse responses and forecast error variance decompositions may be calculated to gauge the relative importance of permanent and transitory innovations in determining fluctuations of the endogenous variables.

2.2. Core inflation estimation

In the empirical analysis we consider a four-variable system including the log of the oil price in US dollars (oil), the log of the industrial production index (y), the monthly rate of change of nominal $M2$ (m), and price inflation measured by the monthly rate of change of the CPI all-items price index (π). With the exception of oil, all series are seasonally adjusted. The evidence provided by Hallman et al. (1991) of a strong long-run link between US inflation and the growth rate of $M2$ motivates our choice of this monetary aggregate.

Standard unit-root tests show that all variables can be treated as $I(1)$ processes. In particular, an *ADF* test on the monthly CPI inflation rate over the sample 1960(2)–2000(4) yields a statistic of −2.39 (with a 5% critical value of −2.87), supporting the non-stationarity of π and confirming the results obtained by Hallman et al. (1991) on annual data from 1955 to 1988 and by Freeman (1998) on monthly data from 1967 to 1996. The vector of endogenous variables is then specified as $\mathbf{x}_t = (\text{oil}_t\, y_t\, m_t\, \pi_t)'$. The main rationale for the inclusion of oil as an endogenous variable in the system is to evaluate the response of the other variables to a major source of supply side shocks by means of impulse responses and forecast error variance decomposition techniques. This should provide valuable additional information about the long-run determinants of inflation.

Cointegration analysis has been carried out using the Johansen (1988) Maximum Likelihood approach over the period 1960(2)–2000(4). Twelve lags of each variable have been included in the short-run specification of the model on the basis of diagnostic tests of dynamic specification, showing that a 12-lag dynamic structure is capable of eliminating all residual serial correlation.

Table 1 reports the results of the cointegration analysis. As expected, the data suggest the existence of one cointegrating vector at the 5% level of significance. From the coefficients of the normalised eigenvector, a long-run relation between nominal money growth and price inflation clearly emerges. Both the oil price and industrial production show very small coefficients. As shown in Table 1, a formal test cannot reject the hypothesis that the cointegrating vector captures the constancy of the rate of growth of real money ($m - \pi$) in the long run. This restriction has therefore been imposed in the rest of the analysis.

In the common trends framework outlined in the previous section, the existence of one cointegrating relationship among four variables implies the presence of three distinct sources of shocks having permanent effects on at least some of the variables. We make the following assumptions on the nature of the three permanent shocks in the system: we consider a *foreign real* shock (ψ_{f}), motivated by the huge oil price movements in the sample period, a *domestic real* shock (ψ_{r}), and a *nominal* disturbance (ψ_{n}). The latter shock (which may be of a domestic or foreign nature) is assumed to have no long-run effect on output (a long-run neutrality assumption), and to affect the rate of money growth and the inflation rate with the same magnitude, given the $(1, -1)$ cointegrating restriction. The domestic real shock is therefore responsible for the long-run fluctuations of output not attributable to the foreign disturbance ψ_{f}. Finally, we assume that both ψ_{r} and ψ_{n} do not have long-run effects on the oil price.

Table 1
Cointegration analysis: 1960(2)–2000(4)

Cointegration tests					
Eigenvalue:	0.070	0.029	0.017	0.0017	
Hypothesis:	$r=0$	$r\leqslant 1$	$r\leqslant 2$	$r\leqslant 3$	
λ_{MAX}	34.8**	14.01	8.30	0.82	
95% critical value	27.1	21.0	14.1	3.8	
λ_{TRACE}	57.93*	23.12	9.11	0.82	
95% critical value	47.2	29.7	15.4	3.8	
Unrestricted cointegrating vector[a]					
	oil	y	m	π	
$\boldsymbol{\beta}_1'$	0.0002	0.0023	1	−1.0092	
	(0.0008)	(0.0021)		(0.2348)	
Restricted cointegrating vector					
	oil	y	m	π	χ^2 test (p-value)
$\boldsymbol{\beta}_1'$	0	0	1	−1	5.28 (0.15)

r denotes the number of valid cointegrating vectors; * and ** denote significance at the 5% and 1% level respectively.
[a] $\boldsymbol{\beta}'$ matrix; cointegrating vector normalised on m; standard errors in parentheses.

The permanent part of the common trends representation is then the following trivariate random walk:

$$\begin{pmatrix}\tau_{\mathrm{f}}\\ \tau_{\mathrm{r}}\\ \tau_{\mathrm{n}}\end{pmatrix}_t = \begin{pmatrix}\mu_{\mathrm{f}}\\ \mu_{\mathrm{r}}\\ \mu_{\mathrm{n}}\end{pmatrix} + \begin{pmatrix}\tau_{\mathrm{f}}\\ \tau_{\mathrm{r}}\\ \tau_{\mathrm{n}}\end{pmatrix}_{t-1} + \begin{pmatrix}\psi_{\mathrm{f}}\\ \psi_{\mathrm{r}}\\ \psi_{\mathrm{n}}\end{pmatrix}_t, \tag{11}$$

where $\boldsymbol{\mu}$ is a vector of constant drift terms, added to the model in estimation. Letting γ_{ij} denote the generic element of $\boldsymbol{\Gamma}_g$, the assumptions above imply $\gamma_{12}=\gamma_{13}=\gamma_{23}=0$. The common trends representation of the variables in levels is therefore the following:

$$\begin{pmatrix}\mathrm{oil}\\ y\\ m\\ \pi\end{pmatrix}_t = \begin{pmatrix}\mathrm{oil}\\ y\\ m\\ \pi\end{pmatrix}_0 + \begin{pmatrix}\gamma_{11} & 0 & 0\\ \gamma_{21} & \gamma_{22} & 0\\ \gamma_{31} & \gamma_{32} & \gamma_{33}\\ \gamma_{41} & \gamma_{42} & \gamma_{43}\end{pmatrix}\begin{pmatrix}\tau_{\mathrm{f}}\\ \tau_{\mathrm{r}}\\ \tau_{\mathrm{n}}\end{pmatrix}_t + \boldsymbol{\Gamma}^*(L)\begin{pmatrix}\psi_{\mathrm{f}}\\ \psi_{\mathrm{r}}\\ \psi_{\mathrm{n}}\\ v_1\end{pmatrix}_t, \tag{12}$$

where v_1 is a purely transitory disturbance (uncorrelated with the permanent shocks) to which, given the main focus of our analysis, we do not attribute any structural economic interpretation. The estimated core inflation series from the common trends model is then computed as $\pi^{\mathrm{c}}_{\mathrm{CT},t} = \pi_0 + \hat{\gamma}_{41}\hat{\tau}_{\mathrm{f},t} + \hat{\gamma}_{42}\hat{\tau}_{\mathrm{r},t} + \hat{\gamma}_{43}\hat{\tau}_{\mathrm{n},t}$. Such a measure captures the long-run effects on inflation of all the identified permanent disturbances and bears the interpretation of the (conditional) forecast of the inflation rate over a long-term (infinite) horizon, when all transitory fluctuations in the inflation rate have vanished.

The Quah and Vahey (1995) procedure, applied to a non-cointegrated bivariate system including only output y and inflation π, would allow identification of only

Table 2
The estimated common trends model

Variables	Shock		
	ψ_f	ψ_r	ψ_n
Long-run effects of permanent shocks (matrix $\boldsymbol{\Gamma}_g$)			
oil	0.0898** (0.0167)	0 (–)	0 (–)
y	−0.0070* (0.0037)	0.0116** (0.0021)	0 (–)
m	0.00011 (0.00006)	0.00009 (0.00006)	0.0003** (0.00003)
π	0.00011 (0.00006)	0.00009 (0.00006)	0.0003** (0.00003)
Long-run (∞) forecast error variance decomposition			
oil	1 (–)	0 (–)	0 (–)
y	0.2663 (0.1989)	0.7337 (0.1989)	0 (–)
m	0.1224 (0.1353)	0.0963 (0.1178)	0.7813 (0.1283)
π	0.1224 (0.1353)	0.0963 (0.1178)	0.7813 (0.1283)

Asymptotic standard errors in parentheses; * and ** denote significance at the 5% and 1% level respectively.

two permanent shocks and no purely transitory disturbance. [4] Core-inflationary shocks would be identified by imposing a zero restriction on their long-run output effect and the core inflation series would then be constructed using only this kind of disturbances. This identification scheme does not allow for long-run inflation movements attributable to real shocks (which affect output in the long run) and does not exploit the long-run link between monetary growth and inflation (a strong feature of the data).

The main results from the estimation of the common trends model are shown in Table 2, where the estimated elements of the long-run impact matrix $\boldsymbol{\Gamma}_g$ and the long-run forecast error variance decomposition of the variables are reported (with asymptotic standard errors in parentheses). [5] A number of features can be noticed from the estimated model. First of all, the foreign (oil price) shock ψ_f has a positive, though not strongly significant, permanent effect on π and m, and a negative and strongly significant long-run effect on industrial production. The real internal disturbance ψ_r affects output in the long run with no statistically relevant effect on inflation and money growth. [6] The latter two variables show a strong long-run reaction to the permanent nominal disturbance ψ_n. The forecast error variance decomposition results give some additional insight about the long-run behavior of the variables.

[4] In the model of Blix (1995), including the price level, output and money, the detection of one cointegrating relationship allows for the identification of two permanent and one transitory disturbances.

[5] Estimation has been carried out using the CT Rats routine of Warne and Hansen.

[6] However, these two shocks do contribute to the explanation of the changes in the core inflation rate at specific times (the '60s and first half of the '70s for the domestic real shock and the oil-shock episodes of the mid-'70s and early '80s for the foreign disturbance).

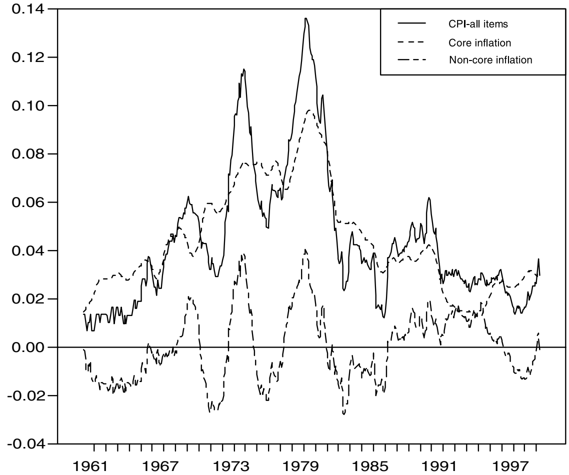

Fig. 1. Measured CPI (all-items) inflation and estimated core and non-core inflation rates from the common trends model (12-month lagged moving averages).

In particular, around 78% of the long-run variability in π and m is attributable to nominal disturbances, with the remaining 22% due to real shocks. The estimated core inflation series from the common trends model, π_{CT}^{c}, is shown in Fig. 1 together with the measured CPI (all-items) inflation and the estimated transitory ("non-core") inflation component, computed as $\pi - \pi_{CT}^{c}$. Twelve-month lagged moving averages of all series are plotted (an observation at month t is the sum of the monthly inflation from month $t-11$ to month t). The main features of this common trends measure of core inflation are discussed in the next section.

3. Core inflation: Discussion

To give reliable information for policy use, a core inflation measure must display some desirable properties, as pointed out by Bryan and Cecchetti (1994) and Wynne (1999). First, the estimated core inflation series should display lower variability and higher persistence than actual inflation. The common trends measure of core inflation portrayed in Fig. 1 is less volatile than measured CPI inflation, with lower peaks during the high-inflation episodes of the mid-'70s and early '80s, mainly due to surges in oil prices, whereas the non-core component closely follows the actual inflation rate, suggesting that the large movements of the latter were mainly of a transitory nature. Moreover, we note that, on some occasions, the core and observed inflation rates followed a very different pattern: for example in 1970–71, when core inflation showed a sharp rise and actual inflation a remarkable fall before starting to rise in the second half of 1972, and in 1975–76, when the large decline in measured inflation occurred with a core inflation broadly constant at the peak level reached in 1974. In the more recent part of the sample, the fall in inflation in 1986 is fully reflected in the transitory component, with only a slight decline in the core measure, whereas in 1991 the reduction in observed inflation is matched by a sharp decline of the core component. The constant observed inflation of 1995–96 is the result of a rising core inflation rate and a declining transitory component; the latter also determined the reduction in measured inflation in 1997, with core inflation fairly stable at the level attained in the previous year. Finally, the steady increase in measured inflation from the beginning of 1999 (from 1.6% in January 1999 up to 3.1% in February 2000) is only partly reflected in a rise in the core inflation measure (which started at 2.9% early in 1999, reaching 3.2% in August to come back to 2.9% in February 2000). The sharp rise in March 2000 (with a measured rate of 3.6%, followed by 3.0% in April) is entirely attributable to the transitory component, with a core inflation rate almost unchanged and very close to 3.0%.

The smoothing property of the estimated core inflation series is further illustrated in Table 3, which reports correlation coefficients among changes in the monthly and annual (12-month moving average) inflation rates, including observed inflation and the common trends core and non-core measures, denoted by $\Delta\pi^{c}_{CT}$ and $\Delta\pi^{nc}_{CT}$ respectively, with $\Delta\pi \equiv \Delta\pi^{c}_{CT} + \Delta\pi^{nc}_{CT}$. Standard deviations in percentage points are shown

Table 3
Correlations of measures of core and non-core inflation

	$\Delta\pi$	$\Delta\pi^{c}_{CT}$	$\Delta\pi^{nc}_{CT}$	$\Delta\pi^{c}_{NFE}$	$\Delta\pi^{nc}_{NFE}$	$\Delta\pi^{c}_{BC}$	$\Delta\pi^{nc}_{BC}$
A. Monthly inflation measures							
1960(4)–2000(4)							
$\Delta\pi$	0.254						
$\Delta\pi^{c}_{CT}$	0.415	0.029					
$\Delta\pi^{nc}_{CT}$	0.994	0.312	0.243				
$\Delta\pi^{c}_{NFE}$	0.361	0.198	0.353	0.237			
$\Delta\pi^{nc}_{NFE}$	0.605	0.209	0.606	−0.524	0.278		
1967(3)–2000(4)							
$\Delta\pi^{c}_{BC}$	0.423	0.180	0.420	0.391	0.078	0.170	
$\Delta\pi^{nc}_{BC}$	0.762	0.307	0.758	0.129	0.622	−0.264	0.237
B. Annual (12-month moving average) inflation measures							
1961(3)–2000(4)							
$\Delta\pi$	0.318						
$\Delta\pi^{c}_{CT}$	0.414	0.095					
$\Delta\pi^{nc}_{CT}$	0.955	0.127	0.292				
$\Delta\pi^{c}_{NFE}$	0.518	0.282	0.473	0.282			
$\Delta\pi^{nc}_{NFE}$	0.583	0.177	0.578	−0.394	0.296		
1968(2)–2000(4)							
$\Delta\pi^{c}_{BC}$	0.632	0.284	0.596	0.732	−0.013	0.251	
$\Delta\pi^{nc}_{BC}$	0.656	0.253	0.632	−0.002	0.740	−0.171	0.258

$\Delta\pi$ denotes the first difference of the measured CPI inflation rate; $\Delta\pi^{c}_{CT}$, $\Delta\pi^{c}_{BC}$ and $\Delta\pi^{c}_{NFE}$ denote the first difference of the common trend (CT), the Bryan–Cecchetti (BC) and the CPI less food and energy (NFE) measures of core inflation rates; $\Delta\pi^{nc}_{CT}$, $\Delta\pi^{nc}_{BC}$ and $\Delta\pi^{nc}_{NFE}$ are the associated non-core inflation changes, defined as $\Delta\pi^{nc}_{x} = \Delta\pi - \Delta\pi^{c}_{x}$. The figures on the main diagonals are standard deviations in percentage points. The first observation available for $\Delta\pi^{c}_{BC}$ and $\Delta\pi^{nc}_{BC}$ in panel A is 1967(3). The corresponding 12-month lagged moving average series in panel B therefore begin in 1968(2).

on the diagonal. These latter statistics show that there is a remarkable difference in variability between the core and the non-core component: standard deviations are 0.03 and 0.24 for $\Delta\pi^{c}_{CT}$ and $\Delta\pi^{nc}_{CT}$ respectively in monthly data (0.10 and 0.29 in annual data), with a standard deviation of changes in the observed inflation rate of 0.25 (0.32). We also note the low positive correlation between core and non-core inflation changes (0.31 in monthly and 0.13 in annual data).

A second desirable property of a core inflation measure is the ability in forecasting future headline inflation rates. The forecasting power of our core inflation measure is warranted, since it is estimated as the long-run conditional forecast of inflation. This property can be formally assessed by means of a bivariate VAR system including the observed inflation rate and core inflation π^{c}_{CT}. As argued by Freeman (1998), the integration and cointegration properties of the inflation series require an error-correction representation to perform appropriate Granger-causality tests. [7] In fact, both π and π^{c}_{CT} are non-stationary, $I(1)$ series, whereas the associated non-core component π^{nc}_{CT} displays stationarity, which may be interpreted as evidence of cointegration between the core inflation measure and the actual inflation rate, since $\pi^{nc}_{CT} \equiv \pi - \pi^{c}_{CT}$. The specification of the bivariate system is then the following:

$$\begin{aligned}\Delta\pi_t &= \delta_{10} + \sum_{i=1}^{2}\delta_{11}(i)\Delta\pi_{t-i} + \sum_{i=1}^{2}\delta_{12}(i)\Delta\pi^{c}_{CT,t-i} + \rho_\pi(\pi - \pi^{c}_{CT})_{t-1} + u_{1t},\\ \Delta\pi^{c}_{CT,t} &= \delta_{20} + \sum_{i=1}^{2}\delta_{21}(i)\Delta\pi_{t-i} + \sum_{i=1}^{2}\delta_{22}(i)\Delta\pi^{c}_{CT,t-i} + \rho_{CT}(\pi - \pi^{c}_{CT})_{t-1} + u_{2t},\end{aligned} \tag{13}$$

where two lags are sufficient to eliminate residual serial correlation. The upper part of Table 4 reports the results of the F-tests on each block of lagged regressors and the coefficient estimates of the error-correction coefficients ρ_π and ρ_{CT}. As expected, π^{c}_{CT} has strong additional predictive power for the actual inflation rate, with the error-correction coefficient on $\Delta\pi$ (−0.45) showing a tendency of actual inflation to adjust to the core component, whereas no adjustment is detected in the behavior of π^{c}_{CT}.

A desirable measure of core inflation should also have some theoretical foundations. Though not derived from a full-fledged theoretical model of the economy, the core inflation measure estimated here is based on a fairly general view of the long-run determinants of inflation: in particular, the relationship between nominal money growth and inflation may be motivated by long-run quantity theory considerations as in Hallman et al. (1991). [8] However, the common trends core inflation

[7] In addition to Granger causality tests, Le Bihan and Sédillot (2000) apply out-of-sample tests to evaluate the forecasting accuracy of various measures of core inflation for France.

[8] Wehinger (2000) derives a measure of core inflation from a fully specified open-economy macro model, imposing a set of long-run restrictions to identify several sources of shocks. The core inflation rate is then constructed by considering only demand-side influences (eliminating the effects of all supply-side disturbances). Unlike the common trends measure of the present paper, the resulting core inflation cannot bear the interpretation of a long-run inflation forecast.

Table 4
Results from bivariate VAR systems

Equation for:	F test (p-value) on 2 lags of:		Coefficient estimate on:
	$\Delta\pi$	$\Delta\pi^{c}_{CT}$	$(\pi - \pi^{c}_{CT})_{t-1}$
System: π, π^{c}_{CT} sample: 1960(6)–2000(4)			
$\Delta\pi$	0.004***	0.004***	−0.448*** (0.059)
$\Delta\pi^{c}_{CT}$	0.995	0.972	0.00001 (0.0083)
System: π, π^{c}_{BC} sample: 1967(5)–2000(4)			
	$\Delta\pi$	$\Delta\pi^{c}_{BC}$	$(\pi - \pi^{c}_{BC})_{t-1}$
$\Delta\pi$	0.169	0.014*	−0.611*** (0.090)
$\Delta\pi^{c}_{BC}$	0.195	0.000***	0.173*** (0.066)
System: π, π^{c}_{NFE} sample: 1960(6)–2000(4)			
	$\Delta\pi$	$\Delta\pi^{c}_{NFE}$	$(\pi - \pi^{c}_{NFE})_{t-1}$
$\Delta\pi$	0.008**	0.000***	−0.496*** (0.078)
$\Delta\pi^{c}_{NFE}$	0.071	0.000***	0.339*** (0.068)

$\Delta\pi$ denotes the first difference of the measured CPI inflation rate; $\Delta\pi^{c}_{CT}$, $\Delta\pi^{c}_{BC}$ and $\Delta\pi^{c}_{NFE}$ denote the first difference of the common trends (CT), the Bryan–Cecchetti (BC) and the CPI less food and energy (NFE) measures of core inflation rates. *, ** and *** denote statistical significance at the 5%, 1% and 0.01% levels respectively. Two lags of the dependent variable are used in all equations. When $\Delta\pi^{c}_{BC}$ is used the first observation available, after allowing for two lags, is 1967(5).

series does not possess other desirable properties. For example, as in the case of all measures derived from econometric procedures, new observations may entail changes of past core inflation figures, adding to the difficulties of using this measure as a part of the public communication strategy of the monetary authorities. Nevertheless, these considerations do not rule out the potential usefulness of the common trends methodology as an "internal" tool of inflation analysis for monetary policy purposes.

As a final point in the discussion, we compare the behavior of the core inflation series estimated from the common trends model (π^{c}_{CT}) with two other widely used measures of core inflation: the Bryan–Cecchetti median inflation series (π^{c}_{BC}), regularly published by the Federal Reserve Bank of Cleveland and available on its web site, and the inflation rate obtained from the CPI index excluding food and energy (π^{c}_{NFE}), computed by the US Bureau of Labor Statistics.

We start in Fig. 2 by showing, in the upper panel, π^{c}_{CT}, π^{c}_{BC} and actual inflation π over the 1968–2000 period. [9] As before, the series are 12-month lagged moving averages of the respective annualized monthly inflation rates. Median inflation more closely follows the behavior of actual inflation, with substantial deviations limited to short periods of time (for example, 1986 and 1997–98), and displays larger swings than the common trends core inflation, especially in the seventies, when π^{c}_{BC} more closely tracks observed inflation and π^{c}_{CT} has smoother fluctuations. In 1970–71

[9] The first observation available for the median inflation series is February 1967. The 12-month lagged moving average of the series therefore begins in January 1968.

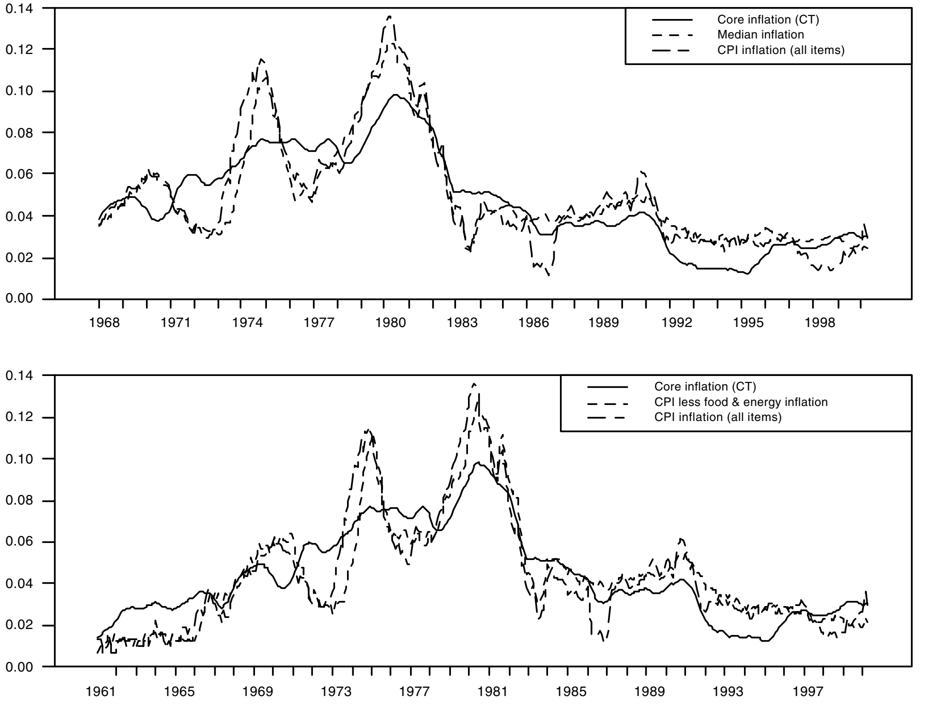

Fig. 2. Actual CPI (all-items) inflation, common trends core inflation, and alternative measures of core inflation (12-month moving averages).

and in 1975–76, the different behavior of the two core inflation measures is particularly evident. On the contrary, there are also important episodes in which the two series show more similar movements: for example, according to both measures the disinflation of 1981–82 (accompanied by a sharp monetary restriction) is mainly due to a decline in the core inflation component, whereas the low inflation of 1986 is almost entirely attributed to the transitory element. Overall, since the mid-'80s the median and the common trend core inflation measures display a more similar behavior, although from 1991 to 1995 the level of π^c_{CT} is substantially lower than π^c_{BC}. Very similar comments apply also to the inflation rate derived from the CPI excluding food and energy, π^c_{NFE}, shown in the lower panel of Fig. 2. Interestingly, from early 1999 π^c_{BC} and π^c_{NFE} display a similar behavior, with a steady decrease towards the 2.2–2.4% range, whereas π^c_{CT} remains close to 3%.

The correlation statistics reported in Table 3 suggest several further comments. First, unlike the common trends estimate of the core and transitory inflation components, $\Delta\pi^c_{BC}$ and $\Delta\pi^c_{NFE}$ display only a slightly lower standard deviation than the associated non-core inflation series in monthly data and an almost identical variability in annual data. Second, $\Delta\pi^c_{CT}$ has a similar correlation with changes in observed inflation ($\Delta\pi$) as the other core inflation measures in monthly data and a lower correlation in annual data (0.41 against 0.63 and 0.52 for $\Delta\pi^c_{BC}$ and $\Delta\pi^c_{NFE}$ respectively).

On the contrary, changes in the common trends, non-core inflation component, $\Delta\pi^{nc}_{CT}$, display higher correlation with changes in measured inflation than the corresponding measures $\Delta\pi^{nc}_{BC}$ and $\Delta\pi^{nc}_{NFE}$ (e.g. 0.95 against 0.66 and 0.58 in annual data). Finally, $\Delta\pi^{c}_{CT}$ has low correlation with changes in the other measures of core inflation (around 0.20–0.30) whereas $\Delta\pi^{c}_{BC}$ is more correlated with changes in the non-core inflation component estimated from the common trends model (0.42 in monthly and 0.60 in annual data). A similar pattern of correlation holds for the CPI excluding food and energy measure. Overall, the correlation patterns show that π^{c}_{CT} has very different properties from π^{c}_{BC} and π^{c}_{NFE}, whereas the latter two core inflation measures share several similar features.

Using the bivariate system (13) with π^{c}_{BC} and π^{c}_{NFE} in turn in the place of π^{c}_{CT} we also evaluated the forecasting power of the two alternative measures of core inflation. Table 4 shows that, as for the common trends core inflation, also the median and CPI less food and energy measures have additional predictive power for the observed inflation rate. However, in the latter two cases the error-correction coefficients are strongly significant in *both* equations. The negative estimate of ρ_π captures the error-correcting behavior of π following a transitory movement in the inflation rate, but the positive estimates of ρ_{BC} and ρ_{NFE} suggest that past values of the inflation rate above the core component cause an increase in the core rate itself, reflecting the transmission of transitory shocks to the measure of the permanent component of inflation. Overall, the above results suggest that those alternative measures of core inflation do not fully reflect permanent inflation as captured by the common trends estimate in our small-scale macroeconomic system. The transitory components do affect the subsequent behavior of the core measures, casting some doubt on their ability to capture the long-run inflation trend.

4. Conclusions

A common trends model has been used to estimate the underlying, "core" inflation behavior for the US over the last four decades. In this framework core inflation is interpreted as the long-run forecast of inflation conditional on the information contained in nominal money growth, output fluctuations and movements in the oil price. We argue that this measure of core inflation may be useful for monetary policy purposes since it embodies long-run economic restrictions strongly supported by the data and bears the interpretation of a long-run forecast, which should be the relevant target variable for monetary policy.

The core inflation series derived from the common trends model has also been compared with other commonly used core inflation measures, such as the median inflation and the CPI inflation excluding food and energy. The latter series display substantially different time-series properties in terms of variability and correlations with the observed inflation rate. Moreover, they are not completely successful in separating the permanent from the transitory component of observed inflation. The core inflation rate derived from a common trends model can then provide useful additional information in evaluating the trend behavior of inflation for monetary policy purposes.

Of course, such core inflation measure will depend 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 a long-run link between inflation and monetary growth, seems an useful benchmark to evaluate the properties of other measures of US core inflation currently used in the monetary policy debate.

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