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The Recent Performance of the Traditional Measure of Core Inflation in G7 Countries

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Abstract. In this paper we undertake an empirical investigation concerning the performance of the traditional measure of core inflation in recent years. We consider the group of G7 countries and explore both the high-frequency and the low-frequency relations between overall inflation and core inflation. We find that the traditional core measure, obtained by subtracting from the overall index those components which exhibit high volatility and which are responsible for the short-run variability of inflation, is a reliable indicator of trend inflation for a group of countries including the USA, Canada and Japan. The innovation accounting shows that for the three countries the transitory shock, *i.e.* the total inflation shock, has limited persistence and hence there is a relatively quick convergence of overall inflation to its trend component.

JEL Classification: C43, E31;

Keywords: Core Inflation Indicator; Structural Cointegrated VARs; Permanent-transitory Decompositions;

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1. Introduction

It is well known that the main central banks have different views about the reliability of core inflation measures interpreted as indicators of low-frequency movements of overall inflation. In particular, the US Federal Reserve System, the Bank of Japan and the Bank of Canada devote a certain attention to such measures. Instead, the European Central Bank (ECB) and the Bank of England share a more skeptical position.

However, the recent evolution of the rate of inflation in industrialized countries, with a generalized and strong increase in 2007-2008 and a subsequent, sharp decline between the fall of 2008 and the summer of 2009, has confirmed the importance of selecting good indicators of the long-run trend of total inflation.

Given the objective of price stability, which is pursued by monetary authorities over the medium run, this question is then strictly related to monetary policy choices by central banks since, also as a consequence of the lags which characterize the transmission of monetary policy decisions, the setting of policy interest rates should abstract from short-run movements of inflation.

In recent years, the traditional measure of core inflation has shown a remarkable stability when compared to the rate of change of the overall consumer price index (CPI). This traditional measure is typically obtained by subtracting from the overall index food and energy components. More generally, the goal in building such indicators is to remove from the CPI index those components affected by high volatility and identified as the dominant source of the short-run movements of total inflation.

It is worth stressing that in the course of 2008 both in the US economy and in the Euro area the underlying inflation did not signal the risk of future growing inflation, staying stable around 2 percent, while during the last 12 months, in the presence of a quick decrease of the inflation rate towards a negative territory, core inflation has gravitated around 1.5 percent.

Indeed, the unfortunate increase of the short-term nominal interest rate in the Euro area decided by the European Central Bank in July 2008 in a context of an already recessionary economy and hence with looming downward pressures on prices, is emblematic of the general difficulty faced by central banks in properly separating short-run movements of inflation from its medium-run tendency.

In this paper we study both the short-run and the long-run interaction between core inflation and total inflation in the G7 countries. In particular, we concentrate on the more recent decades.

We build on a strand of the literature which tests for core inflation indicators in the framework of cointegrated VARs (see *e.g.*, Freeman 1998, Marques *et al.* 2002, Ribba 2003).

Quah and Vahey (1995) introduced the Structural VAR methodology in this area of research. The authors obtain a measure of core inflation for UK from a bivariate VAR model in which the structural shocks are recovered by imposing long-run neutrality restrictions. Also based on a strategy of imposing neutrality restrictions but in the presence of identified cointegration relations is the research by Bagliano and Morana (2003) who aim to obtain an estimation of core inflation for the US economy.

Another recent application of the structural VAR methodology is provided by Martel (2008). The author builds a trend inflation indicator for Canada by estimating a trivariate VAR model which includes oil-price growth, output growth and inflation.

In the present research, we adopt a cointegration framework and estimate bivariate cointegrated VARs for the G7 countries in order to investigate: (a) the existence of a stable one-for-one long-run relation between core inflation and overall inflation; (b) the presence of unidirectional long-run causality running from core to overall inflation.

As shown in the literature (see *e.g.*, Ribba 2003), if both the conditions (a) and (b) are satisfied, then using the core measure as an indicator of the trend inflation is legitimate. Moreover, under conditions (a) and (b), a Wold causal ordering, with core inflation ordered first, implies that shocks affecting total inflation have only transitory effects on both the variables. In this context, the innovation accounting also provides useful information on the speed of adjustment to long-run relation, *i.e.* on the persistence of a diverging path between core and overall inflation.

The last question has a practical relevance. For, suppose that the core measure explains all the variability in overall inflation at horizons, say, of thirty or forty years: although this would, without doubt, be an interesting result it probably would not exert any practical influence on the effective conduct of monetary policy by central banks.

In the empirical investigation conducted in the present paper we obtain mixed results: there is a group of countries including USA, Canada and Japan in which the selected core inflation measure seems to satisfy the set of conditions that allows the identification of a trend inflation indicator. Moreover, and not less important, the impulse-response functions reveal that the transitory, total inflation shock does not exert persistent effects. This is a common conclusion for the three countries since the convergence of overall inflation to the core component, in the sample periods considered in this empirical investigation, has a range comprised between 9 months (the case of Japan) and 18 months (for the United States).

Instead, in the other countries and areas, it is not possible to draw a similar conclusion since the dynamic behaviour of the series at very low frequencies is consistent with a representation in first differences but, nonetheless, the existence of equilibrium long-run relations is rejected by data.

However, it is worth emphasizing that for the Euro area economy the sample period covers around ten years and hence a possible explanation of the rejection of cointegration, which is after all a long-run phenomenon, rests on the limited sample horizon of the investigation.

The paper is organized as follows. In section 2 we motivate the specific choice of the core inflation indicators for the countries included in the investigation. Of course, for those countries in which the central bank monitors a particular index we test for the reliability of such index. We also briefly discuss the different positions expressed by central banks about the validity of trend inflation indicatiors.

In section 3 the econometric approach of the paper, based on structural cointegrated VARs, is presented. Section 4 shows the main empirical results concerning the G7 countries and, for those countries in which the presence of a long-run equilibrium relation between core and total inflation is not rejected by data, we give particular attention to the analysis of the cointegration space. In section 5 we study the dynamic effects of structural shocks affecting overall and core inflation for those countries in which the underlying inflation reveals itself to be a reliable indicator of the permanent component of total inflation.

Section 6 concludes.

2. Core inflation indicators and central banks

In recent decades central banks have put more emphasis, with respect to the first decades of the postwar period, on the goal of price stability. This goal is pursued over the medium-run horizon. Moreover, it seems that the main central banks substantially converge to an operational quantification of price stability defined in terms of the year-on-year rate of change of the Consumer Price Index around two percent. For example, this indication is explicit for the ECB's choice of a rate of inflation "below but close to 2 percent"[1]

Where, instead, central banks diverge is on the opportunity of adopting core inflation indicators as a guidance for the underlying trend in total inflation.

A traditional measure of core inflation is obtained by subtracting from the overall CPI index some components characterized by high volatility and interpreted as responsible for the short-run movements in total inflation.

Typically, the components subtracted are food and energy but in recent years some central banks have developed further refinements on core indicators[2].

The conduct of monetary policy and the settings of nominal short-term interest rates should, ideally, focus on the persistent component of inflation, both for the variable lags by which monetary policy is trasmitted to the economic system and in order to avoid risks of instability induced by an excessive monetary policy activism.

Moreover, the economic interpretation underlying the logic of removing food and energy components from the total index rests on the consideration that these components are driven by factors which are often external to the domestic economy and hence outside the control, at least the direct control, of the central bank.

[1]The indication is contained in a publication of May 2003. The document contains a revision in the operational definition of price stability by the ECB, previously indicated in terms of an annual inflation below 2 percent. In this revision the ECB has made clear that risks of deflation need to be considered in the design of the monetary policy strategy (ECB Press Release, May 8, 2003, "The ECB's monetary policy strategy).

[2]A recent and exhaustive survey of concepts and issues related to the selection of core inflation indicators is provided by Wynne (2008).

Further, in an epoch in which central banks attribute great importance to trasparency and thus to commmunication with the public, the simplicity of the traditional core inflation indicators is a good point in support of their adoption.

On the other hand, for those central banks more reluctant to the adoption of core indicators, as for instance is the case of the European Central Bank, the risk of losing control of inflationary expectations has been perhaps the most important indication against the reliability of core inflation measures.

Not by mere chance, in the ascendant phase of the oil price during 2007-2008 and of the related increase of total inflation, President Trichet has repeatedly emphasized that the ECB was worried and vigilant about the risks of the so-called second-round effects running from current, total inflation to expectations on future inflation. The possible transmission of these effects to nominal wage growth and hence the potential risk of a spiral between prices and wages induced the ECB to repeated threats of interest rate increases, particularly during the spring of 2008.

However, although there is no official measure of trend inflation monitored by the ECB, in recent months the European Central Bank has shown a greater attention to core inflation indicators[3].

Indeed, the core measure given by the overall HCPI index ex food, energy, alcool and tobacco components, has had a good performance in the last 24 months. For it is worth noting that, according to this indicator, in the first part of 2008 there had been no signal of an increasing future inflation in the Euro area.[4] Instead, the evolution of expected inflation, as measured by the periodical Survey of Professional Forecasters had pointed towards an increase of inflation, over a one year horizon, above the 2 percent level.

Unfortunately, this increase in the inflationary expectations may have played an important role in the central bank's decision to raise the short-term interest rate in July 2008.

A skeptical view on core inflation also seems to characterize the more recent monetary policy conduct by the Bank of England. For, starting from 2003, the central bank has abandoned the target defined in terms of the Retail Price Index ex mortgage interests (RPIX) and, currently, the inflation target of 2 percent is expressed by referring to the overall CPI[5].

Nevertheless, the Bank of England has recently reconsidered the possibility of introducing an intermediate target based on core inflation (cf. Bank of England, 2008).

Instead, within the group of countries and central banks included in this

[3]See, for example, the section: "Recent developments in selected measures of underlying inflation in Euro area", Monthly Bulletin, June 2009, 54-59. As stressed by Wynne (2008), the ECB in its monthly publication provides detailed informations on the evolution of various rates of inflation in the Euro area which are based on price indices from which some volatile components are removed.

[4]In the light of the good performance of this indicator in the last months, we have chosen this core measure for the Euro area in the empirical investigation of section 4.

[5]Although as from 1997 the Bank of England has an operational independence, the inflation target is established by the Government. Moreover, four out of nine members of the Bank's Monetary Policy Committee are appointed by the Chancellor of the Exchequer.

study, the Federal Reserve System, the Bank of Japan and the Bank of Canada attibute an explicit role to core inflation indicators.

The Federal Reserve System does not quantify the notion of price stability nor has it defined an explicit framework for its monetary policy strategy. However, there is an established tradition within the Open Market Comittee (FOMC) of attention to core inflation indicators.

In recent years the preferred measure has shifted from the CPI ex food and energy towards the core Personal Consumption Expenditures (PCE) inflation.

As emphasized by Bernanke and Mishkin (1997), monetary policy should neglect short-run movements in the price series, and hence adopting core inflation indicators might help to identify the persistent component of inflation. Moreover, by focusing on this persistent component the central bank may also achieve the goal of stabilizing inflationary expectations [6].

It is worth stressing that the Federal Reserve System's mandate has a dual character since the goal of price stability coexists with the goal of stabilization of the real economy. This is of course an important, formal difference with respect to the mandate of ECB which, enshrined in the Mastricht Treaty of 1992, gives a hierarchical priority to price stability.

It should be added that on the basis of the first ten years of experience of ECB's activity, and abstracting from a certain communicative rhetoric of ECB members on price stability, it is hard to identify such relevant differences between the two central banks in the effective conduct of monetary policy. For, although with variable lags, both restrictive and expansionary cycles in the settings of short-term nominal interest rates in the Euro area have systematically followed monetary policy decisions made by the US Fed.

The Bank of Canada has an inflation target of 2 percent based on the year-on-year change in the overall CPI but with a flexible range of 1 - 3 percent. Among the central banks considered in this investigation, the Bank of Canada attibutes the greater importance to measures of core inflation interpreted as the long-run trend of headline inflation. Moreover, starting from 2001 a new measure of core inflation was adopted, the CPIX, which excludes the eight most volatile components of the overall index. The excluded components are: fruits, vegetables, gasoline, fuel oil, natural gas, mortgage interest, inter-city transportation and tobacco products.

Since 1991 the Bank of Canada has been embracing inflation-targeting and in the first years of this monetary regime the selected core inflation measure was the CPI ex food and energy. Nevertheless, further research conducted within the central bank led to to the choice of CPIX as a more reliable indicator[7].

As far as the Bank of Japan is concerned, one important peculiarity of its monetary policy strategy framework is that the notion of price stability is revised by the Policy Board every year. This high-frequency revision in the

[6]The reference to these authors is not casual since both of them have contributed to shape the effectual conduct of monetary policy in the United States, not only with their researches but also with their role within the FOMC

[7]An evaluation of alternative core inflation measures concerning the Canadian economy is provided by Armour (2006).

quantification of price stability by the central bank might be a consequence of the particular experience of Japan in the second part of the 90s, when the economy was hit by a persistent deflation.

For the current year, an inflation rate comprised in the range between 0 and 2 percent is consistent with the view expressed by each member of of the Policy Board (cf. Bank of Japan, 2009).

Moreover, in order to identify the long-run trend of inflation, the central bank monitors a wide set of core inflation measures, including the CPI ex food and energy. The attention devoted in recent months to this price index is, of course, due to the great fluctuations of oil prices between 2007 and 2008 which has exerted an important short-run influence on overall inflation. However, the preferred core inflation indicator remains the CPI ex fresh food (see, for example, Shiratsuka, 2006).

3. The approach of the paper

Let us assume that the inflation rate, $\pi = (ln(P_t) - ln(P_{t-12})) * 100$, behaves as an I(1) variable. Any I(1) series can always be thought of as composed of the sum of a permanent and a transitory component, where these components follow, respectively, an I(1) and an I(0) stochastic process (cf. Quah, 1992).

If we assume that also the year-on-year core inflation measure, given by $\pi^* = (ln(P_t^*) - ln(P_{t-12}^*)) * 100$, is an I(1) variable it is then possible to establish a set of conditions under which it can be interpreted as the permanent component of inflation.

Given the assumptions on the difference-stationary behaviour of the two series, their joint dynamics has the following reduced-form Wold representation:

$$\begin{pmatrix} \Delta\pi_t^* \\ \Delta\pi_t \end{pmatrix} = \begin{pmatrix} C_{11}(L) & C_{12}(L) \\ C_{21}(L) & C_{22}(L) \end{pmatrix} \begin{pmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{pmatrix} \qquad [1]$$

where Δ is the first difference operator and L is the lag operator, with $C(0) = I$. $\epsilon_t = (\epsilon_{1t}, \epsilon_{2t})'$ is the (2x1) vector of reduced-form disturbances such that $E(\epsilon_t) = 0$ and $E(\epsilon_t \epsilon_t') = \Omega_\epsilon$.

The folllowing set of conditions (cf. Ribba, 2003) is required in order to identify π^* as the trend component of π:

(i) the matrix of long-run multipliers, C(1), has reduced rank 1, *i.e.* the core indicator and total inflation are cointegrated.

(ii) the cointegrated vector has the form $(1, -1)'$;

(iii) There is one-way causality at frequency zero running from core inflation to overall inflation. In the context of a cointegrated system this implies that core inflation does not adjust to long-run equilibrium.

It is not necessary to impose further restrictions on causality relationships at frequencies other than zero. For, let us suppose the presence of bidirectional Granger causality and hence that total inflation contains information for predicting short-run movements of the core component. Despite this possible short-run interaction, the established set of conditions allows the core indicator to be identified as the permanent (trend) component of total inflation.

In order to motivate this assertion, note that under the joint set of conditions from (i) to (iii), there is the following implication for the conditional expectation of $E_t(\pi_{t+h})$ for long forecast horizon, since we have:

$$\lim_{h\to\infty} \frac{\partial E_t(\pi_{t+h})}{\partial \pi_t^*} = \lim_{h\to\infty} \frac{\partial E_t(\pi_{t+h}^*)}{\partial \pi_t^*} \neq 0 \qquad [2]$$

$$\lim_{h\to\infty} \frac{\partial E_t(\pi_{t+h})}{\partial \pi_t} = \lim_{h\to\infty} \frac{\partial E_t(\pi_{t+h}^*)}{\partial \pi_t} = 0 \qquad [3]$$

In words. Let us assume an unexpected increase at date t of the core inflation measure. As a consequence, the long-run forecast of both total inflation and core inflation will then be revised upward. Instead, let us suppose that total inflation at date t is higher than expected. Although in this case an upward revision of both variables for short horizons is possible, the long-run forecast of inflation will be unchanged.

Remark 1. Freeman (1998) was the first to introduce the framework of cointegration in investigating the properties of trend inflation indicators. In his research Freeman states that the two variables should exhibit an equilibrium long-run relation with a cointegrated vector of form (1, -1). Besides, he suggested that a good core indicator should not be Granger-caused by inflation. In his empirical investigation, concerning the United States, the author does not find much support for the traditional measure of core inflation.

Marques *et al.* (2002) formally establish a set of conditions that a trend inflation measure should satisfy and, moreover, they evaluate a set of potential trend inflation measures for a group of industrialized countries. In particular, they add to the three conditions from (i) to (iii) a further restriction by maintaining that a condition of strong exogeneity for the core indicator is also required. Once again the empirical results gainsay the reliability of the rate of change of CPI ex food and energy as a trend inflation indicator.

It is worth noting that if the set of conditions from (i) to (iii) is satisfied, then there exists the following Error Correction Model representation:

$$\Delta\pi_t^* = A_{11}(L)\Delta\pi_t^* + A_{12}(L)\Delta\pi_{t-1} + \epsilon_{1t} \qquad [4]$$

$$\Delta\pi_t = A_{21}(L)\Delta\pi_{t-1}^* + A_{22}(L)\Delta\pi_{t-1} - \alpha_2(\pi_{t-1}^* - \pi_{t-1}) + \epsilon_{2t} \qquad [5]$$

Since the goal pursued consists in the selection of a long-run indicator for total inflation, then the further restriction of strong exogeneity, which requires $A_{12}(L) = 0$ and implies $C_{12}(L) = 0$, is unnecessary. For, unidirectional causality at frequency zero (cf. condition (iii)) ensures that $A_{12}(1) = 0$ and, equivalently, that $C_{12}(1) = 0$.

Hence by indicating with β the $(2x1)$ vector of cointegration coefficients and with α the $(2x1)$ vector of loadings, conditions (ii) and (iii) imply that $\beta = (1, -1)'$ and $\alpha = (0, \alpha_2)$.

Summing up this part: the condition of strong exogeneity for core inflation, sometimes advocated in the literature, would be satisfied in the presence of a random walk process for the core indicator; building on the literature on the permanent-transitory decompositions, we have argued in this section that this restriction on the dynamic shape of the permanent component of inflation is unnecessary[8].

Remark 2. In building core indicators, emphasis is usually put on a property of smoothness of the selected measure. In other words, the so-called exclusion measures of core inflation are built by removing the most volatile components from the overall index.

While smoothness is, of course, a reasonable and expected property of any trend indicator, it should be stressed that in our framework removing the most volatile components from the total price index does not necessarily ensure that the core measure be a good trend indicator of inflation. For, by assuming that the inflation rate exhibits a stochastic trend, a crucial condition to be satisfied in order to select an appropriate measure of trend inflation is that the dynamic process of the residual components subtracted from the overall index has no power at very low frequencies. In principle, the volatile components of the index do not necessarily satisfy this restriction.

Remark 3. We are also interested in recovering the structural disturbances which affect this bivariate dynamical system and, in particular, we want to separate a permanent from a transitory shock. In general, in VAR models including I(1) variables, with or without cointegrating relations among the variables, this separation is obtained by imposing long-run zero restrictions for the effects of some structural shocks. Nevertheless, in our empirical analysis, we impose a Wold causal form with the core inflation measure ordered first in the causal ordering since, given the assumption of weak exogeneity for core inflation, *i.e.* $\alpha_1 = 0$, a contemporaneous restriction is equivalent to a long-run restriction.

[8]As a further indication that the set of conditions that we require in order to obtain a permanent-transitory (P-T) decomposition of total inflation indeed allows the goal to be achieved, note that the restrictions from (i) to (iii) meet all the conditions required by the Gonzalo-Granger (1995) P-T decomposition in cointegrated systems. However, in the bivariate case considered in the present paper, the stochastic trend is given by a directly observable economic variable. The Gonzalo-Granger decomposition builds upon Quah's (1992) work. The authors develop a method to identify the long-memory component in cointegrated VAR models.

In order to briefly explain this equivalence result, let H(0) be the unique lower triangular matrix such that $H(0)H(0)' = \Omega_\epsilon$. We obtain the following structural representation:

$$\begin{pmatrix} \Delta\pi_t^* \\ \Delta\pi_t \end{pmatrix} = \begin{pmatrix} H_{11}(L) & H_{12}(L) \\ H_{21}(L) & H_{22}(L) \end{pmatrix} \begin{pmatrix} \eta_{1t} \\ \eta_{2t} \end{pmatrix} \qquad [6]$$

where $H(L) = C(L)H(0)$, $\eta_t = H(0)^{-1}\epsilon_t$ and $E(\eta_t\eta_t') = I$. $\eta_t = (\eta_{1t}, \eta_{2t})'$ is a $(2x1)$ vector of structural disturbances. Under this structural representation a change in total inflation does not exert a contemporaneous effect on core inflation, but an implied result is that such a change is also neutral in the long run and, by virtue of the cointegration relation, has only a transitory effect on both variables.

Recall that as a consequence of condition (iii) the error-correction term does not enter the core indicator equation and this implies that the second column of C(1) contains zero elements. Given orthonormal disturbances, the contemporaneous restriction, adopted in order to obtain exact identification of the model, imposes that H(0) be lower triangular, *i.e.* $H_{12}(0) = 0$.

However, since the matrix of structural long-run multiplier, H(1), is given by $H(1) = C(1)H(0)$ the important conclusion is that also the second column of H(1) contains zero elements, *i.e.* a shock to total inflation, with core inflation fixed, has only a transitory effect on both the variables.

Hence, this is a case of equivalence between short-run and long-run identifying restrictions in bivariate cointegrated VAR models (cf. Cochrane, 1994 and Ribba, 1997).

A recent article by Fisher and Huh (2007) proposes further equivalence results for cointegrated models with a number of variables greater than two. Moreover, in a very recent paper, Keating (2009) provides necessary and sufficient conditions under which a Wold causal ordering and a long-run recursive VAR give identical results in a more general multivariate context including both difference-stationary and trend-stationary variables.

4. An empirical investigation for the G7 countries

In this section we present the main results concerning the cointegration analysis of core measures and total inflation for the G7 countries. As for the European countries, we consider aggregate data for the Euro area and also analyze the case of Germany, France and Italy which are the largest economies of the European Monetary Union (EMU).

Both for the aggregate Euro area data and for the member countries of EMU the sample period considered in this investigation is 1999:1 - 2009:6. Hence, we aim to cover the period of activity of ECB.

As for the Euro area, we consider a bivariate VAR model including the annual inflation rate measured by the HICP and the core inflation measure obtained by the overall index ex food, energy, alcool and tobacco. Instead, as far as the three EMU countries are concerned, we utilize the more traditional core index ex food and energy.

As for United States and Japan the sample period considered is 1987:1 - 2009:6. For the first country the choice implies an assumption of continuity of the central bank's monetary policy strategy from Greespan's leadership to that of Bernanke. This assumption seems, indeed, quite reasonable.

The estimated model for the USA includes the year-on-year rate of inflation based on the overall PCE index and the core indicator obtained by subtracting food and energy components from the total index.

In the case of Japan, concerning the choice of the sample period, we believe it is important in the empirical investigation to cover also the deflationary period of the 1990s. The bivariate VAR model for this country includes the rate of change of the CPI and a core measure based on the exclusion of fresh food.

Instead, the sample period for Canada is 1996:1 - 2009:6. This choice is due to the fact that the data attaining the new core inflation measure monitored by the central bank, the CPIX, are available only from 1996.

Finally, for the UK the period under investigation is 1997:1 - 2009:6. The choice of the sample period rests on the important institutional change which happened in 1997, when the Government gave the the Bank of England operational independence in the setting of policy interest rates.

The variables included are, respectively, the inflation rate measured by the total CPI and the rate of change of the CPI ex food and energy.

In table 1 we summarizes the overall results obtained for the group of countries under investigation. As shown, the hypothesis of existence of an equilibrium long-run relation between core inflation and total inflation is rejected for the Euro area and the UK[9]. In particular, the results which are based on Johansen's tests on the rank of the long-run matrix imply that the joint dynamics of the two variables is consistent with a representation in first difference but without the presence of cointegration[10].

Insert table 1 about here

[9]The dynamic behaviour of both inflation and the selected core indicators is consistent with the assumption that the series are I(1). This conclusion is based on the battery of unit roots conducted. Although it is well known that these tests suffer from low power in small samples as against sufficiently close alternatives, the conclusion on the I(1) behaviour of the variables is also confirmed by the multivariate Johansen tests. Given the amount of tests and results conducted in this paper for the group of countries, we report only the principal results. Of course, all the output obtained is available on request.

[10]We have selected the lag length for each estimated VAR on the basis of Schwarz and Hannan-Quinn criteria. Moreover, we have also computed likelihood ratio tests for restrictions on the deterministic trend coefficients. For those countries in which the existence of cointegration is not rejected by data, the tests (not reported) support the zero restriction on the constant term.

Instead, as far as United States, Canada and Japan are concerned, there is evidence of an equilibrium long-run relation between the core indicator and overall inflation[11] and hence, in these cases, we go on to analyze the properties of the cointegration space, *i.e.* we test both for the existence of a cointegrating vector of form (1,-1)' and for the weak exogeneity of core inflation.

Another interesting question concerns the possible presence of Granger causality running from total inflation to core inflation. The last column of table 1 shows that only in the case of Japan are both weak and strong exogeneity detected in the data. Indeed, for this country, the dynamic behaviour of underlying inflation is consistent with a random walk process. As for United States and Canada, there is, instead, evidence of short-run predictability of total inflation with respect to the core indicator and hence the data do not reject the hypothesis of bidirectional Granger causality.

Thus our results, at least for USA, Canada and Japan support the choice made by the monetary authorities of these countries to rely on core measures as indicators of the long-run evolution of total inflation. This conclusion is partially in contrast with the results presented in Freeman (1998) and Marques *et al.* (2002) since the authors find evidence against the reliability of the traditional measure of core inflation for the United States and a group of European countries[12].

It is important to stress that if we had considered the presence of unidirectional Granger-causality as a key condition for validity of the core measure, we should have had to draw a positive conclusion only for the case of Japan.

On the other hand, we have repeatedly emphasized in this paper that this potential short-run bidirectional influence between the two variables does not prevent an interpretation of the core indicator as the permanent, long-run component of total inflation.

Summing up the results of this part: as shown in tables from 2 to 4, the analysis of the cointegration space for this group of countries reveals that there is evidence in favor of core inflation as an indicator of the underlying trend in total inflation.

Insert table 2 about here

Insert table 3 about here

Insert table 4 about here

[11]Since we are considering small samples, we have also used the small sample correction suggested by Cheung and Lai, 1993. There are only little changes in the statistics obtained and hence the results supporting the presence of cointegration are confirmed for all the three countries.

[12]However it should be added that the sample period considered in the present paper is different, since the investigation focuses on the more recent decades and, clearly, this adds another possible source for the differences in the results obtained. Moreover, in the paper by Marques *et al.* Japan and Canada are not included in the empirical investigation and hence the contrasting results concern essentially the most important economic area, *i.e.* the United States.

5. The dynamic effects of structural shocks and the speed of convergence to long-run equilibrium

Figures from 1 to 6 show the impulse-response functions and the forecast error variance decomposition analysis (FEV) for each of the countries in which the empirical investigation reveals a good performance of the core measure monitored by the national central bank[13].

For the US economy (see figure 1), a transitory shock, *i.e.* an unexpected increase in total inflation, causes a temporary increase in core inflation and hence there is a significant short-run influence. However, the transitory shock does not exhibit high persistence since the effects vanish after around 18 months.

This (relatively) limited persistence of the overall inflation shock has, of course, important implications for the usefulness of the core indicator since the selected measure of trend inflation can also be viewed as a reliable operational guide for the central bank's monetary policy conduct.

The innovations accounting for the US is completed by the presentation of the FEV decomposition (see figure 2).

There is an important feature of this analysis that is worth stressing: the permanent, core inflation shock explains almost all of the variability of the core PCE inflation, whereas the transitory shock explains almost all of the variability of total inflation at high frequencies. Nevertheless, the role of the permanent shock becomes dominant over the medium-low frequencies.

Insert figure 1 about here

Insert figure 2 about here

The conclusion deriving from the innovations accounting concerning the Canadian economy are not very different from those of US economy, since there is a reasonable speed of convergence to the long-run equilibrium: after around 12 months the effects of the transitory, total inflation shock becomes negligible (see figure 3) and the movements in overall inflation are mainly driven by shocks to the core inflation.

A similar picture emerges from the FEV decomposition (cf. figure 4). The short-run variability of overall inflation is largely dominated by unexpected changes in total inflation but from horizons of 12-24 months the role of changes in core inflation become increasingly important and, ultimately, dominant.

Insert figure 3 about here

[13] We plot the impulse-response functions with the 90 percent confidence bounds based on the analytical formulae developed by Amisano and Giannini (1997).

Insert figure 4 about here

Some important differences instead characterize the dynamic responses of the two variables to structural disturbances in the case of Japan. For, the response of core inflation to a transitory shock is not significant at (almost) all horizons (cf. figure 5).

Indeed, this dynamic response of core inflation to the total inflation shock is consistent with a random walk behaviour of the dynamic process of the trend inflation indicator.

These results also show up in the FEV analysis: total inflation shocks have no power in explaining the variability of the core component at all horizons whereas the core indicator has a pre-eminent role in explaing the variability of overall inflation, not only at medium-low frequencies but also at business cycle frequencies.

Insert figure 5 about here

Insert figure 6 about here

7. Concluding remarks

Central banks do not agree on the relevance of core inflation indicators which are obtained by removing from the total CPI index those components affected by high short-run variability. Nevertheless, some among the central banks of the group of G7 countries and, precisely, the Federal Reserve System, the Bank of Canada and the Bank of Japan adopt such indicators as a guide for anticipating the medium-run evolution of total inflation.

In this paper we have tried to test for the reliability of the core measures adopted by these three central banks and, moreover, we have also investigated the low-frequency properties of the relation between the underlying inflation and the total inflation characterizing the Euro area, the UK and the three largest economies of EMU.

The results, concerning the most recent decades, reveal that for the United States, Canada and Japan the traditional measures of core inflation meet some important properties which legitimate the interpretation of core inflation as an indicator of the long-run trend in overall inflation. These properties essentially refer to the existence of a one-for-one long-run relation with total inflation and to the presence of unidirectional causality at frequency zero running from core to total inflation.

Another important conclusion deriving from the innovations accounting for this group of countries is that the convergence of total inflation to its trend indicator requires no more than two years. In this sense, the indicator not only has a sound empirical basis, it also has a practical reliability as a reference for monetary policy choices.

Instead, the positive conclusions concerning the reliability of core inflation measures do not extend to the European countries. For in this area, although behaving as I(1) series, the core indicator and total inflation do not exhibit a long-run equilibrium relation.

However, even in the light of the good performance of core inflation measures in the Euro area in the last two years, one may not exclude that for this area we need a longer span of data to validate the use of core inflation indicators by means of cointegration techniques.

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DATA APPENDIX

For all countries we use monthly data. In the empirical investigation conducted in section 4 of the paper the inflation rate is always defined as the year-on-year rate of change of the underlying index, $\pi = (ln(P_t) - ln(P_{t-12}) * 100$.

Consequently, also the core inflation rate is defined in terms of the year-on-year rate of change of the selected index.

U.S.

The price indices used in the empirical investigation are the Personal Consumption Expenditure index (PCE) and the PCE ex food and energy.

Source: FRED database; research.stlouisfed.org

CANADA

As for Canada the inflation rate is based on the Consumer Price Index (CPI). The core inflation measure adopted by the central bank is the CPIX. The CPIX removes from the total index the eight most volatile components: fruits, vegetables, gasoline, fuel oil, natural gas, mortgage interest, inter-city transportation and tobacco products.

Source: Bank of Canada; www.bankofcanada.ca

JAPAN

Total inflation is calculated on the basis of the CPI index.

Source: OECD; www.oecd.org

The core inflation measure is based on the CPI ex fresh food. This is the preferred core indicator monitored by the central bank.

Source: Japan Statistics Bureau; www.stat.go.jp

EURO AREA

Overall inflation for the Euro area is obtained from the Harmonized Consumer Price Index (HCPI).

The core indicator used in the empirical analysis of section 4 is the HCPI ex food, energy, alcool and tobacco.

Source: Eurostat; ec.europa.eu/eurostat

In the case of Germany, France and Italy we use the CPI and the measure of underlying inflation is built by removing from the overall index food and energy components.

Source: Eurostat

UK

Also in the case of United Kingdom total inflation is based on the CPI index while the core indicator is obtained by excluding food and energy components.

Source: OECD; www.oecd.org

Table 1. Summary of the results for the short-run and long-run interaction between the core measure and overall inflation for the G7 countries

Country	π^* and π are I(1)	rank C(1) = 1	$\beta = (1, -1)'$	$\alpha_1 = 0$	$C_{12}(L) = 0$
Euro Area	yes	no			
Germany	yes	no			
France	yes	no			
Italy	yes	no			
UK	yes	no			
USA	yes	yes	yes	yes	yes
Canada	yes	yes	yes	yes	yes
Japan	yes	yes	yes	yes	no

Notes: The notation is based on equations from (1) to (3) of section 3. A reduced-form bivariate VAR model, including core inflation and total inflation, was estimated for each country. The rank of the long-run matrix, C(1), and the restrictions on the cointegrating vectors and on the vectors of loading are based on the Johansen (1991) framework. The null of a vector of adjustment coefficients $(0, \alpha_2)$ implies testing for unidirectional long-run causality. In the last column, the null of $C_{12}(L) = 0$ tests for unidirectional causality at all frequencies.

Table 2.
Parameter Estimates and Analysis of the Cointegration Space: USA

	core inflation π^*	total inflation π
Normalized cointegration vector	1	-0.977 (0.057)
Loading coefficients	0.001 (0.0016)	0.081 (0.029)
H_0: The cointegration space contains		
the cointegrating vector	1	-1
and		
the loading coefficients	0	α_2
$\chi^2_{(2)} = 0.167$ P-value 0.92		
H_0: $\Delta\pi$ does not Granger cause $\Delta\pi^*$		
$\chi^2_{(2)} = 10.40$ P-value 0.006		

Notes: The results are based on an estimated vector error-correction model with 2 lags in differences. For all the estimated models the lag length is selected on the basis of the Schwarz and the Hannan-Quinn criteria. The sample period is 1987:1 - 2009:6. The numbers in parentheses are standard errors. The null of a cointegrating vector $(1, -1)'$ and of unidirectional long-run causality, (0, α) is a joint test for conditions (ii) and (iii) of Section 3. Johansen's likelihood ratio test of restrictions on the cointegrating vectors and on the loading coefficients is distributed as a chi-squared with degrees of freedom equal to the number of restrictions tested.

Table 3.
Parameter Estimates and Analysis of the Cointegration Space: Canada

	core inflation π^*	total inflation π
Normalized cointegration vector	1	-0.944 (0.097)
Loading coefficients	-0.019 (0.0025)	0.132 (0.051)
H_0: The cointegration space contains		
the cointegrating vector	1	-1
and		
the loading coefficients	0	α_2
$\chi^2_{(2)} = 0.704$ P-value 0.703		
H_0: $\Delta\pi$ does not Granger-cause $\Delta\pi^*$		
$\chi^2_{(1)} = 5.43$ P-value 0.021		

Notes: The results are based on an estimated vector error-correction model with 1 lag in differences. The sample period is 1996:1 - 2009:6. See also notes in table 2.

Table 4.
Parameter Estimates and Analysis of the Cointegration Space: Japan

	core inflation π^*	total inflation π
Normalized cointegration vector	1	-0.972 (0.033)
Loading coefficients	-0.049 (0.043)	0.291 (0.062)
H_0: The cointegration space contains		
the cointegrating vector	1	-1
and		
the loading coefficients	0	α_2
$\chi^2_{(2)} = 1.522$ P-value 0.467		
H_0: $\Delta\pi$ does not Granger-cause $\Delta\pi^*$		
$\chi^2_{(2)} = .844$ P-value 0.358		

Notes: The results are based on an estimated vector error-correction model with 1 lag in differences. The sample period is 1987:1 - 2009:6. See also notes in table 2.

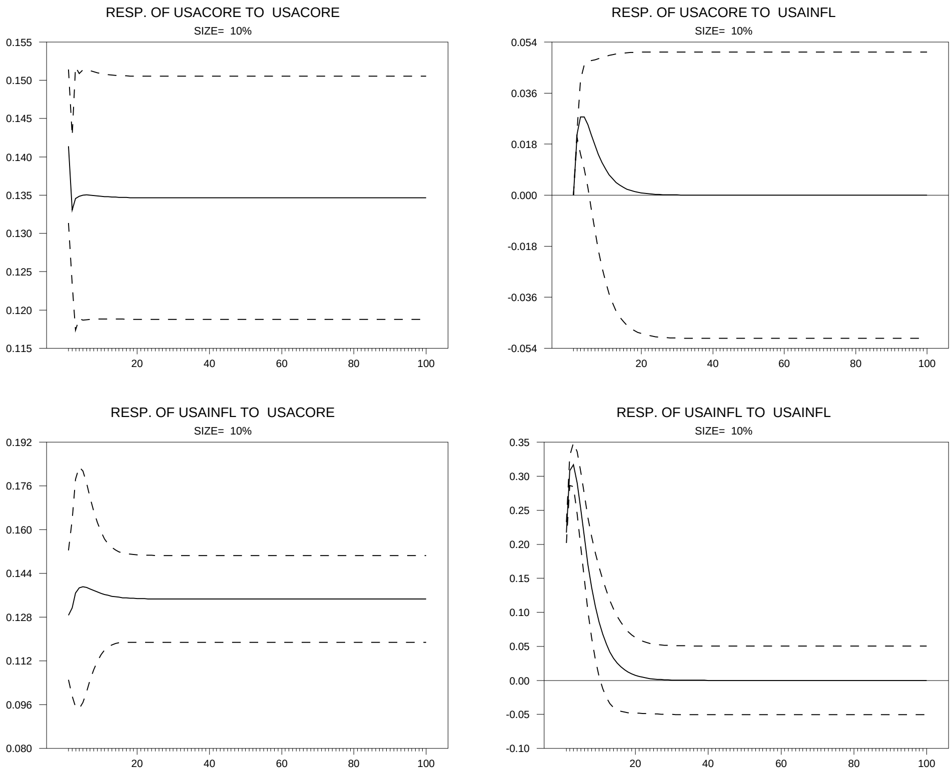

Figure 1 Impulse Response Functions: United States

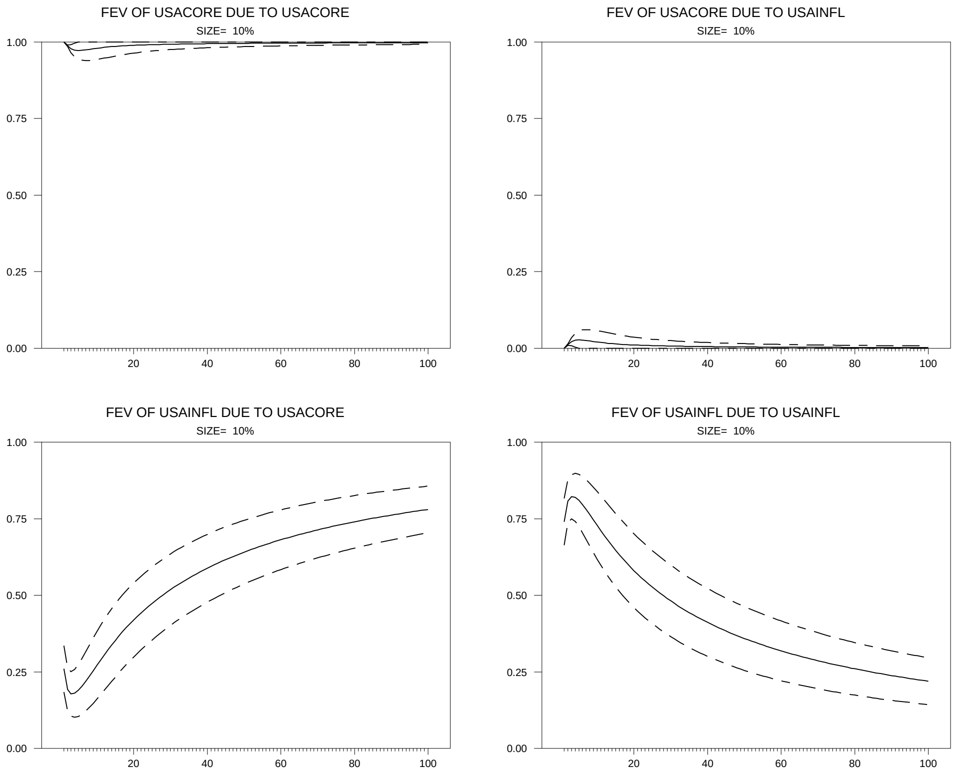

Figure 2 Forecast Error Variance (FEV) Decomposition: United States

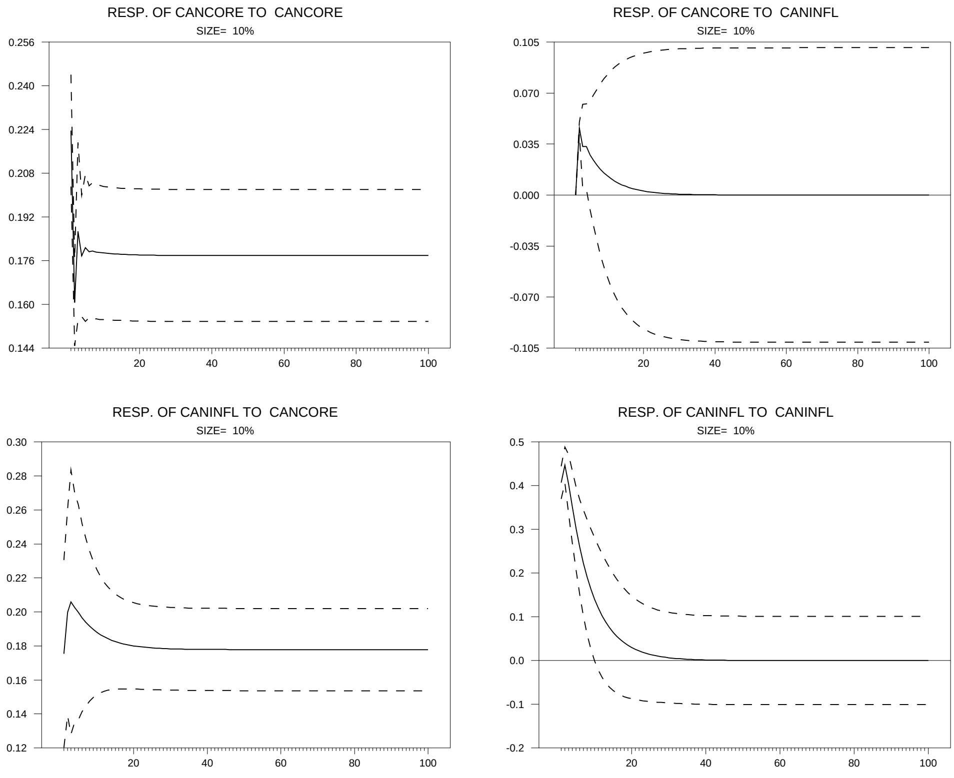

Figure 3 Impulse Response Functions: Canada

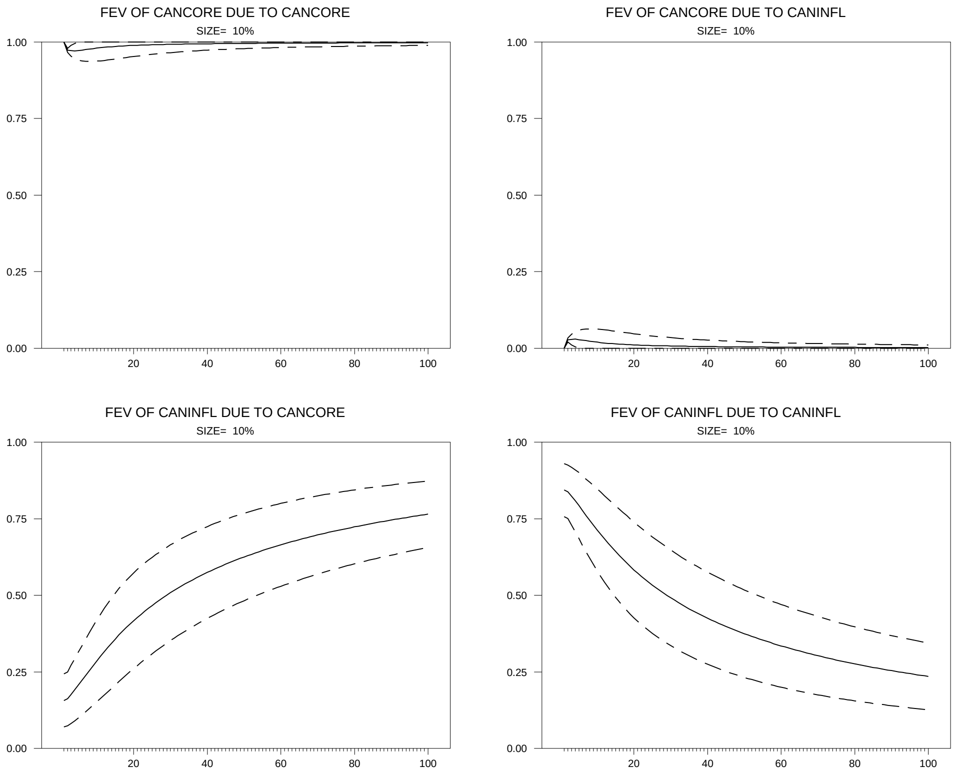

Figure 4 Forecast Error Variance (FEV) Decomposition: Canada

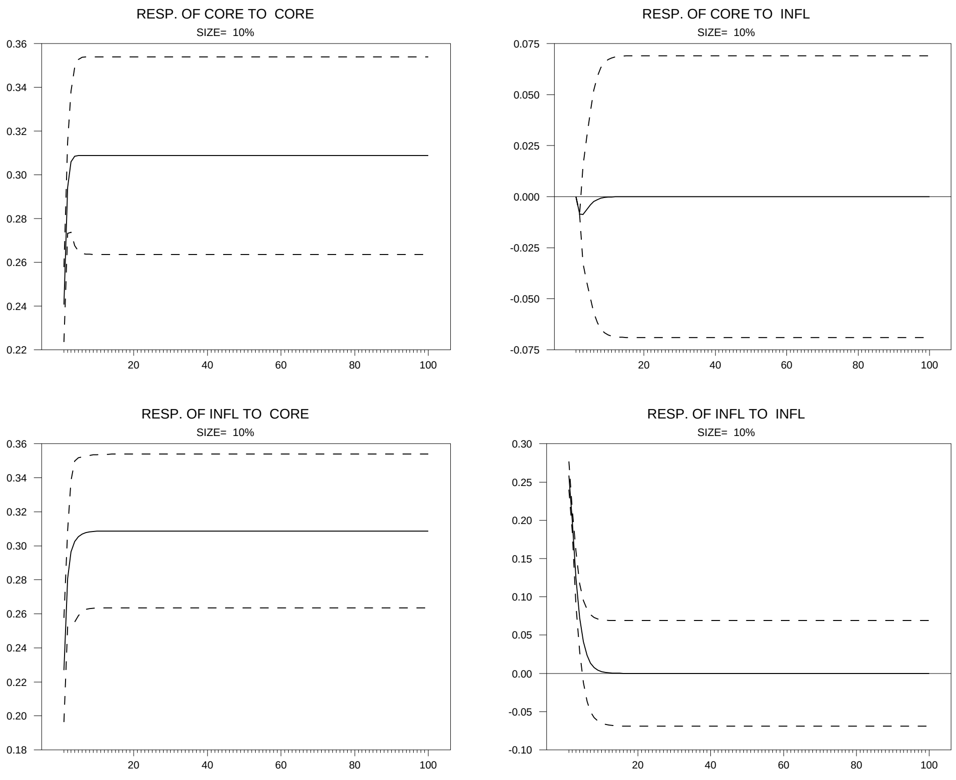

Figure 5 Impulse Response Functions: Japan

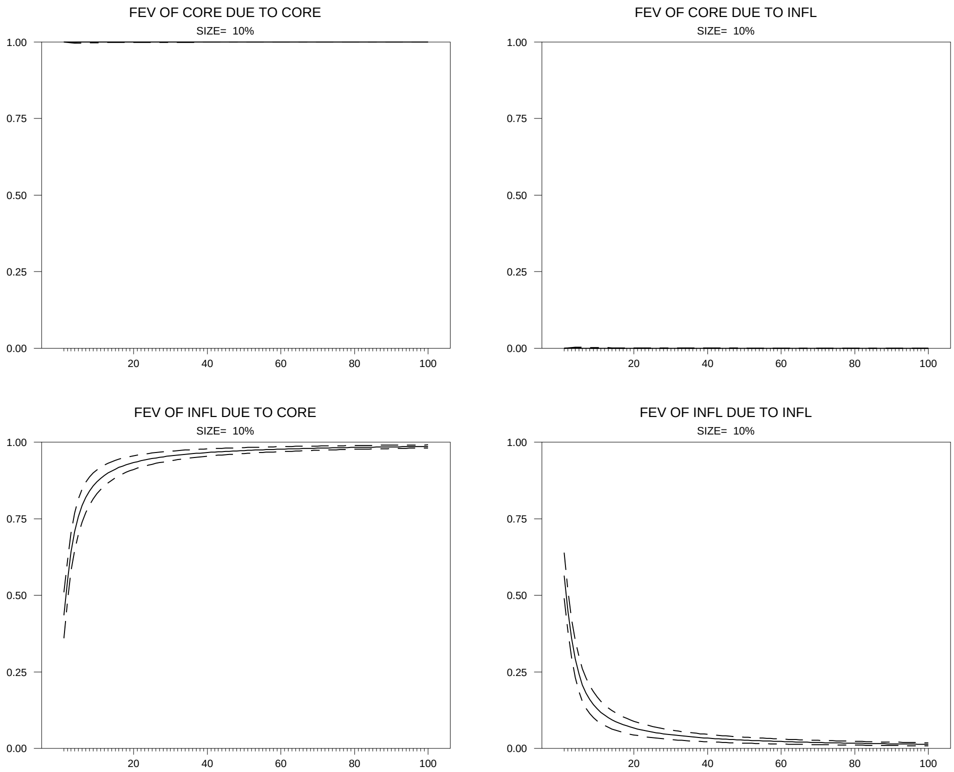

Figure 6 Forecast Error Variance (FEV) Decomposition: Japan

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