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Evidence from the European Union countries

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Contents

Abstract

Whether excluding food and energy components from overall price indices produces a useful indicator for monetary policy purposes is widely debated. The proposals of model based measures of underlying inflation are scarce and the evidence on their performance is limited. In the paper the multidimensional performance of exclusion and model based core inflation indicators is compared in the period of persistently low inflation and interest rates. Providing new measures of underlying inflation we look for specific features of such indices as: tracking trend, appropriate smoothing, unbiasedness with respect to the cost-of-living index, good approximation of the demand pressure, and good short- to medium-term forecasting abilities. To this end, we extract permanent and transitory components of headline HICP and core inflation in the sample of 26 European Union countries for the period 2002-2016 using bivariate unobserved correlated components model and maximum likelihood estimator. We construct an aggregate performance measure, named Core Inflation Score, to capture different dimensions of underlying inflation indicators which could be of interest in monetary policy analysis.

JEL classification: E31, E52, F62.

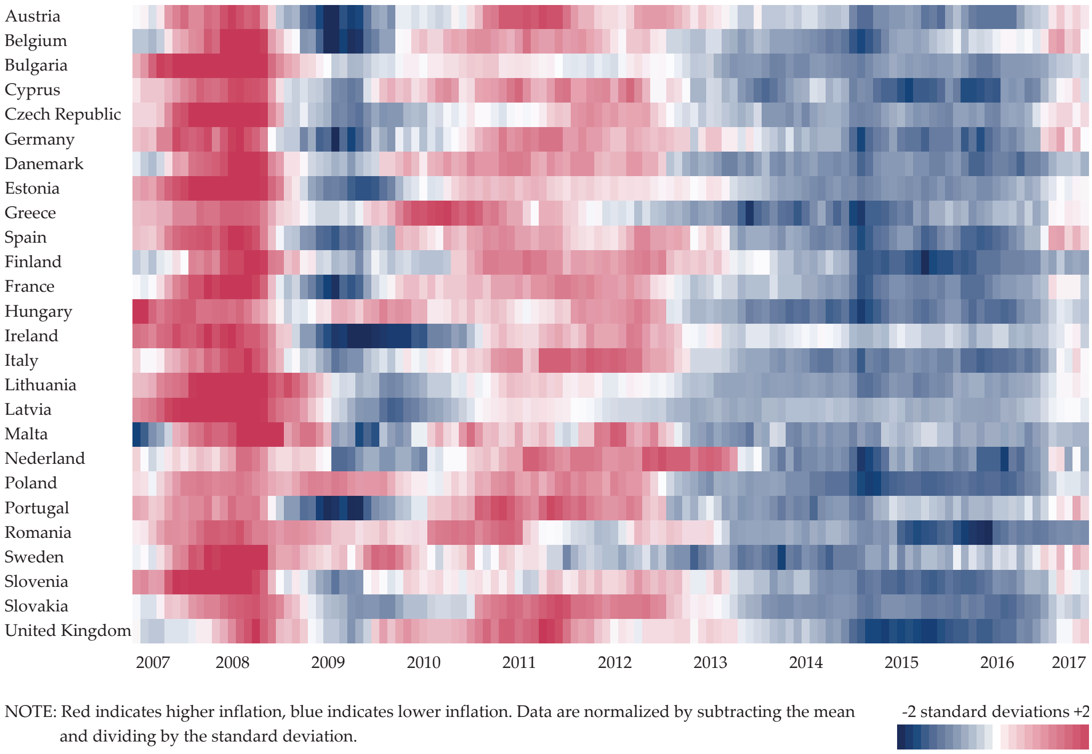
Keywords: core inflation, unobserved correlated components model, forecasting inflation.

1. Introduction

Deep and long-lasting decreases in consumer inflation to the levels markedly below inflation targets have been observed since 2013 throughout the European Union economies (see Fig. 1). In many countries they were accompanied with historically low records of core inflation indicators, which together with slowdown in economic activity raised concerns about bringing secular deflation in the euro area (Crafts 2014). The period of persistently ultra-low inflation, nicknamed as *lowflation*, is also a challenging time to a monetary policy. Any further deflationary shocks in the economy would be hard to offset with a conventional monetary policy, when there is zero lower bound on nominal interest rates accompanied with borrowing constraints and downward nominal wage rigidity (see Iossifov, Podpiera 2015, Casiraghi, Ferrero, 2015).

The discussion on common origins of shocks to domestic inflation in a sample of advanced economies developed into a strand of literature on global inflation (among others: Borio and Filardo 2007, Hakkio 2009, Ciccarelli and Mojon 2010). Whether globalization in tradable goods, food and energy commodities may lead to a deflation spiral and worsen economic conditions worldwide is questioned, however. These deflation concerns are disregarded as *lowflation* is perceived to be mainly driven by persistent supply shocks (ECB, 2016). These are technological innovations, like e.g. explorations of unconventional sources of oil and gas, which might have decreased tensions on global energy market prices in a prolonged way without hampering economic growth in the long run. Although the headline inflation is currently rebounding after the episode of 'missing inflation' (observed in the euro area after 2012), a decline in the underlying inflation trend persists. This shift in an underlying inflation could be also a result of a de-anchoring of inflation expectations (Łyziak, Pallovita 2016) or demographic factors (Bobeica et al. 2017) but the evidence in the euro area is not very conclusive (Ciccarelli and Osbat 2017). The persistence is puzzling unless there is a significant transmission of shocks from food and energy prices to non-energy goods and services. First, ex food and energy inflation in the euro area lag behind headline inflation at least by 6 months (see ECB, 2016) suggesting that the delayed transmission mechanism is in action. Second, the former would cast into doubt the usefulness of exclusion-based inflation indicators in forecasting future price developments. In the paper we ask whether studying common trends in core and headline inflation across EU countries is still informative for extracting an underlying inflation trend.

Fig. 1 Heat maps of monthly data on inflation rates (HICP all-items, year over year) across EU countries.



NOTE: Red indicates higher inflation, blue indicates lower inflation. Data are normalized by subtracting the mean and dividing by the standard deviation.

Source: Own calculations.

In the literature there is no single definition of an underlying inflation. Two concepts are prevailing. According to the first one, an underlying inflation is supposed to be an auxiliary tool in conducting monetary policy. It should be an indicator of monetary inflation influencing or being influenced by other variables which are relevant for monetary policy analysis (output gap, monetary aggregates, exchange rates inter alia). Historically, core inflation is aggregated from these prices which adequately reflect the changes in the domestic demand pressure (see Roger 1997).

Policymakers often refer to inflation excluding unprocessed food and energy prices as a demand-driven inflation indicator. They assume that shocks to food and energy prices are to a large extent supply driven and exogenous, hence they are neither under policy control nor persistent (Labonte 2008). Contrarily, many components of the ex food and energy inflation are insensitive to the domestic output gap – see evidence for Poland by Hałka and Kotłowski (2014) and for euro zone by Fröhling and Lommatzsch (2011). At the same time there are food and energy components of CPI that are sensitive to the cyclical position of the domestic economy (e.g. bread and cereals or heat energy). Contemporarily, the issues of open economy (global inflation literature) and price stickiness (Wynne 2008) are more and more important for discussing how to measure underlying inflation.

According to the second concept core inflation should approximate, with some precision, the underlying trend in general inflation (Bryan et al., 1997, Cecchetti, 1997, Clark, 2001). Good trend evaluation is a starting point for making reliable predictions of inflation which is important from the perspective of forward-looking monetary policy (Bryan and Cecchetti, 1994). In this vein core inflation is the inflation filtered out from the short-term and transitory shocks (Stock and Watson, 2016). While extracting signal (permanent component) from the noise (temporary component) two approaches are the most popular.

First one is based on the measures that exclude some categories of a consumption basket, usually the goods with the most volatile prices changes like food and energy goods (Wynne 2008). In more advanced versions of the exclusion-based approach only a central part of cross-sectional distribution of price dynamics is averaged with tail dynamics being discarded (e.g. trimmed mean approach or weighted median of Bryan and Cecchetti, 1994). There is also a related research on core inflation indicators, in which some components of CPI basket are down-weighed instead of being excluded (Laflèche and Armour, 2006). What is important these indicators tend to meet most of the criteria stated by Wynne (1999). They are computable in a real time, they have a track record and most of all they are not revised when new data arrive. Additionally, the most basic exclusion-based core inflation indicators are easy to understand by the public what helps to anchor inflation expectations. Being useful in communication and tracking inflation does not necessarily mean good forecasting power as the excluded or down-weighted components may be under influence of persistent shocks. Several research show that exclusion-based measures are rather poor in predicting future price changes (Bryan and Cecchetti, 1994, Freeman, 1998, Clark, 2001, Cutler, 2001, Rich and Steindel, 2005, Bermingham, 2010, Detmeister, 2012, Smith 2012, Garnier et al., 2015). Bradley et al (2015) using unobserved component models provide the evidence that permanent component of the US inflation is not well captured by a popular core inflation measure (i.e. inflation excluding energy and food prices). They find that standard core inflation measures may temporally overstate or understate the true permanent component of inflation. Moreover, the excluded items (food and energy prices) contain not only transitory but permanent components as well. Similar findings, on low volatility of US food inflation in 2000s, by Stock and Watson (2016) support inclusion of food prices into their multivariate trend inflation index.

The second decomposition approach is based on time-series smoothing methods which is the focus of our paper. On one hand they rely on applying statistical filters to time series of inflation. The trend figures approximated by centered moving averages (in time domain) or by Baxter-King filter (in frequency domain) are not available in real time and hence they are not very useful in predictions. The simple moving averages (Bryan and Cecchetti, 1994) and unobserved components models (Stock and Watson 2016) proved to be useful in subsamples mimicking the real-time behaviour of these series. On the other hand economists apply cross-section filters (e.g. factor models) to average price dynamics across many goods and sectors (Reis and Watson, 2010) or they broadly project underlying inflation trend from a large cross-sections of inflation-related variables in the vein of dynamic common factor model (Cristadoro et al. 2005, Amstad, Potter, Rich, 2014). However, these model based core inflation measures are subject to a critique because of data mining concerns, the influence of revisions as new data arrive and are hard to be verified by the general public.

Accepting the policy makers' point of view the construction of ideal core inflation measure should be based on what they want this index to illustrate. Generally, monetary policy authorities should focus rather on the persistent movements in prices and not react to the transitory ones. Wynne (2008) additionally points out that it would be preferable that core inflation could approximate future price developments as the main task of the central banks is to target the future inflation. This view is articulated by many economists and several central banks conduct research on the forecasting performance of the core inflation measures (to mention among others Amstad et al., 2014, Detmeister 2012, Khan et al., 2015). These diversified concepts may lead to very different core inflation indicators out of which some of them may be biased estimates of the headline inflation. Therefore, when inflation is close to its historical bounds and monetary authorities frequently use core inflation in communicating their policy decisions (Draghi, 2015, Narodowy Bank Polski, 2016), it would be desirable for the core inflation to be quite close approximation of the cost of living as well.

Taking monetary policy perspective many of these features would be desirable. As it is not possible to guarantee them in one measure, many central banks publish more than one core inflation indicator, often including measure that discards a volatile part of the CPI basket. Therefore, we focus in our work on the measures published by the Eurostat for all European Union countries i.e. HICPs excluding unprocessed food and energy (XUFE, henceforth). We aim to check whether they measure trend inflation or whether the exclusion based approach drops a significant part of the permanent inflation component. Moreover, we investigate to what extent this core inflation indicator reflects the changes in demand pressure and we check its forecasting performance.

To answer these questions we concentrate on the HICP excluding unprocessed food and energy for European Union economies and in addition, in case this measure fails to meet the desired criteria, we analyse permanent components of headline and XUFE inflation from the bivariate unobserved component model. In the decomposition of both indicators into permanent and temporary components we follow the method proposed by Bradley et al (2015). First, we evaluate whether selected core inflation index measures the underlying trend in a precise manner and whether it does not deviate persistently from the cost of living. Second, to address the question of the role of core inflation as demand factor we test whether domestic and foreign output gaps are relevant in description of inflation-output nexus motivated by the Phillips curve. As the New Keynesian Phillips Curve may not match data well (e.g. Rudd and Whelan, 2007) and because different specifications are suitable for different countries we decided to test backward-looking equations with domestic and foreign output gaps as well as with several popular control variables. Third, we check the forecasting performance of the selected measures using the method proposed by Bermingham (2010). Finally, to summarize the results we propose a synthetic measure Core Inflation Score which evaluates different core inflation indices in a simple and comprehensive way. Similar research was performed by Roger (1997), Wynne (1999), Clark (2001), Marques et al. (2003), and Silver (2007).

Our paper adds in three ways to the existing literature. Firstly, our research encompasses wider range of countries than in Bradley et al. (2015). It covers countries different in terms of economic development (advanced and emerging markets) as well as trade openness and vulnerability to external shocks. Secondly, to overcome a problem of overlapping samples in year-over-year monthly indices we conduct the analysis on month-over-month basis obtaining results similar to the findings of Bradley et al. (2015). Thirdly, we account for possible breaks in HICP indices located usually after the outbreak of the global financial crisis or sovereign debt crisis and being strongly supported by the data. The breaks in the mean of inflation rates may reflect the decrease in mean of long-run inflation highly debated in the literature (cf. Ciccarelli and Osbat, 2017). This empirical observation is an issue of monetary policy relevance as it may be a symptom of downward movement in inflation expectations (so called 'de-anchoring').[1] Lastly, we propose a synthetic measure (Core Inflation Score) to evaluate the usefulness of exclusion based and model based measures of underlying inflation. The rest of the paper is organized as follows: Section 2 describes the data, Section 3 characterise the model used for inflation decomposition, Section 4 contains the discussion of the results and Section 5 concludes.

[1] There is also a hypothesis of increased inflation persistence as the most popular alternative to the hypothesis of decline in trend inflation. Both hypotheses are constructed to explain the determinants of 'missing disinflation' and 'missing inflation' observed in the euro area (Ciccarelli and Osbat, 2017).

2. Data

In the study we examine co-movements of two consumption price indices (HICP all items and HICP excluding energy and unprocessed food, XUFE) across 26 EU countries (see Tab. 1, column 1). The monthly data on all-items HICP (2015=100) and XUFE span the period from January 2001 to November 2016.[2] The data for analysis are seasonally adjusted if necessary and monthly inflation rates are approximated by logarithmic changes.

We focus on one of the exclusion-based measures of core inflation published by Eurostat (i.e. XUFE) for three reasons. Firstly, exclusion based measures are easy to communicate and understand by the public. Secondly, unprocessed food and energy components apart from being volatile are often hit by the shocks that are unexpected, transitory and of high magnitude. These are shocks from commodity markets which influence energy prices and unexpected changes in agrometeorological conditions which influence food supply and unprocessed food prices. Thirdly, the choice of XUFE in the policy analysis reflects the preferences of policy makers and central banks not only to focus on headline inflation but also on price indices which are directly influenced by factors under control of monetary policy (e.g. ECB, 2016, NBP, 2016). The prices of many XUFE items include substantial labour costs and hence they are demand driven. The researches on the output sensitivity of inflation's components (e.g. Fröhling and Lommatzsch (2011) for EMU and Hałka and Kotłowski (2014) for Poland) indicate that a non-negligible part of the processed food is output sensitive, too. Therefore, differently from other studies, we do not omit processed food from core inflation basket as these goods become less and less

[2] The beginning of the sample is determined by HICP data availability for most of the countries under analysis (2001), except for Hungary and Romania, where XUFE indices are available from 2002. We do not include Croatia (HICP excluding unprocessed food and energy is available from 2016) and Luxemburg.

dependent on the costs of agricultural inputs and their distribution channels resemble those of other (non-food) industrial goods.

We start with a battery of statistical tests for each of price indices to reveal the important features of their data generating process (dgp), separately. We test whether price changes follow univariate unit-root process against very general hypothesis of stationarity[3] (as in GLS-detrended Dickey-Fuller test) and additionally against an alternative hypothesis with a single break in a linear trend (Zivot-Andrews break test). The evidence whether consumption prices are integrated of order one (inflation is stationary then) or stationary around a deterministic trend is ambiguous. It depends on the country (emerging markets vs. advanced economies), monetary policy regime (e.g. before and after EMU accession), and timespan covered (including or not including non-stationary disinflation period of 1990s for the transition EU countries). In 9 countries (BG, CY, DK, FR, IE, LV, MT, SE, and UK) a stationary process around a constant or a linear trend is preferred for both monthly log-price changes of headline inflation see results of DF-GLS and KPSS tests in Tab. 1) and XUFE (Tab. 4 in Appendix), which means that seasonally adjusted price indices are difference or trend stationary. In most of other cases we find that each of analysed indicators (HICP and XUFE) is covariance stationary if one properly accounts for a single break in their dgps.[4] The breaks, we find in both price indices (Tab. 1 and Tab. 4, columns 4), may result from purely statistical reasons (data collection and methodological changes), or they may be an outcome of changes in economic processes per se. In fact in many cases single breaks indicated by Zivot-Andrews tests are located near global financial crisis (2008/2009) or soon after the outbreak of sovereign debt crisis in EU zone (2011/2012).

[3] Additionally, we use KPSS test for trend stationarity to confirm the results of unit-root tests.
[4] The exceptions are headline HICP in Romania (inflation being close to I(1) process) and XUFE in Netherlands (KPSS test indicates I(0) in price level).

Tab. 1 Unit-root and stationarity tests for headline HICP inflation series (logarithmic changes in price indices)

	HICP all-items						
	DF-GLS		**KPSS**		**Zivot-Andrews**		**conclusions**
country	t-stat	I(1)	LM	I(0)	break	pval	on stationarity
(1)	*(2)*		*(3)*		*(4)*		*(5)*
AT	-11,60	***	0,12		2012m9	<0.01	I(0)
BE	-1,04		0,13		2008m6	<0.01	break
BG	-4,92	***	0,89	***	2008m7	<0.01	I(0)
CY	-9,66	***	0,89	***	2013m1	<0.01	I(0)
CZ	-1,93	*	0,18		2008m1	<0.01	break
DE	-0,75		0,31		2007m11	<0.01	break
DK	-3,87	***	0,66	**	2012m8	<0.01	I(0)
EE	-4,21	***	0,31		2008m1	<0.01	I(0)
EL	-3,17		1,15	***	2011m11	<0.01	break
ES	-1,32		0,91	***	2012m9	<0.01	break
FI	-3,44	***	0,19		2008m5	<0.01	I(0)
FR	-2,34	**	0,60		2008m11	<0.01	I(0)
HU	-0,94		1,01	***	2012m1	<0.01	break
IE	-2,43	**	0,99	***	2008m6	<0.01	I(0)
IT	-6,86	***	0,80	***	2013m2	<0.01	I(0)
LT	-1,20		0,26		2004m5	<0.01	break
LV	-2,02	**	0,46	**	2009m2	0,02	I(0)
MT	-15,47	***	0,35	*	2008m7	<0.01	I(0)
NL	0,08		0,76	***	2009m7	<0.01	break
PL	-7,81	***	0,49	**	2012m6	<0.01	I(0)
PT	-1,14		0,86	***	2012m1	<0.01	break
RO	0,37		1,41	***	2010m9	0,24	no
SE	-7,81	***	0,58	**	2003m2	<0.01	I(0)
SI	-5,34	***	1,22	***	2008m11	<0.01	I(0)
SK	-0,19		1,26	***	2004m1	<0.01	break
UK	-2,87	***	0,31		2011m1	<0.01	I(0)

*Note: DF-GLS, KPSS and Zivot-Andrews stand for the results of univariate stationarity tests for inflation series: t-stat for unit-root test of Elliot, Rothenberg, Stock (1996) with I(1) as an alternative hypothesis, LM statistics for stationarity test of Kwiatkowski, Phillips, Schmidt and Shin (1992) with I(0) as an alternative, and p-value (pval) of unit-root test of Zivot and Andrews (1992) with a single break I(0) as an alternative, respectively. In the results of DF-GLS and KPSS tests the following stars; ***, **, *, correspond to 1%, 5% and 10% significance level of rejecting the null hypotheses in these tests. Description "I(0)" in conclusions (column 5) indicates that the tests provide the evidence of stationarity of log-price changes. Description "break" indicates stationarity of log-price changes after allowing for a single break at the date point indicated by Zivot-Andrews test. Description "no" indicates inconclusive outcomes of stationarity tests.*

Source: Own calculations.

Summing up, the inflation rates in EU countries in 2000s happen to be quite persistent (permanent shocks to prices dominate over transitory components) but still if we allow for breaks in the mean of the long-run trend inflation they tend to be stationary. In the empirical analysis presented in the next section we check whether these breaks individually selected by univariate tests of Zivot-Andrews are still statistically significant in bivariate models.

3. Bivariate unobserved correlated components model

To separate trend components of inflation series from the temporary components we estimate a bivariate unobserved correlated components (bUCC) model described by equations (1) to (4), separately for each of the country samples. In this framework we take into account correlations between shocks to transitory and permanent components and across indices as in Bradley et al (2015). In a baseline specification, following evidence from unit-root tests (cf. Data section), we assume that (logarithms of) consumer prices are $I(1)$ but not cointegrated and that there are breaks[5] in random-walk drifts of a long-run inflation rates:

$$p_{it} = \tau_{it} + c_{it} \tag{1}$$

where p_{it} $(i = 1,2)$ are log-transformed one-base (2015=100) seasonally adjusted monthly HICP indices: headline HICP $(i = 1)$ and XUFE $(i = 2)$. Permanent components, being potentially unit-root processes with a drift, and transitory ones, being AR(2) weak-stationary processes, are described by the following linear equations:

$$\tau_{it} = \mu_i + \mu_{i,t_{i,br}} I_t(t_{i,br}) + \tau_{it-1} + \eta_{it}, \tag{2}$$

$$c_{it} = \phi_{i1} c_{it-1} + \phi_{i2} c_{it-2} + \varepsilon_{it}. \tag{3}$$

Parameters μ_i represent deterministic trends in a permanent part of inflation series, $\pi_{it} = 100 * (p_{it} - p_{it-1})$, $\mu_{i,t_{br}}$ are single breaks in these trends after period $t_{i,br}$, and $I_t(\quad)$ are indicator variables such as $I_t(t_{i,br}) = 1$ for $t \geq t_{i,br}$ and 0 otherwise. Finally, $\eta_{it}, \varepsilon_{it}$ are jointly normally distributed and correlated innovations with zero mean and a constant covariance matrix:

[5] The potential break dates are motivated by univariate break point tests but the final selection is based on the results of likelihood ratio tests in unobserved component model.

$$Var\left(\begin{bmatrix}\eta_{1t}\\ \eta_{2t}\\ \varepsilon_{1t}\\ \varepsilon_{2t}\end{bmatrix}\right)=\begin{bmatrix}\sigma^2_{\eta 1} & \sigma_{\eta 1,\eta 2} & \sigma_{\eta 1,\varepsilon 1} & \sigma_{\eta 1,\varepsilon 2}\\ \sigma_{\eta 1,\eta 2} & \sigma^2_{\eta 2} & \sigma_{\eta 2,\varepsilon 1} & \sigma_{\eta 2,\varepsilon 2}\\ \sigma_{\eta 1,\varepsilon 1} & \sigma_{\eta 2,\varepsilon 1} & \sigma^2_{\varepsilon 1} & \sigma_{\varepsilon 1,\varepsilon 2}\\ \sigma_{\eta 1,\varepsilon 2} & \sigma_{\eta 2,\varepsilon 2} & \sigma_{\varepsilon 1,\varepsilon 2} & \sigma^2_{\varepsilon 2}\end{bmatrix}. \quad (4)$$

In the bUCC model (1)-(4) long-run component of inflation (being inflation expectations for the period $t+1$ formulated at t is $E_t(\tau_{it+1})-\tau_{it}=\mu_i+\mu_{i,t_{i,br}}I_{t+1}$, parameters ϕ_{i1} and ϕ_{i2} are bounded by weak-stationarity assumption. All these parameters (including those in a matrix of covariance innovations) are estimated by maximizing likelihood function with Kalman filtering and smoothing steps. After smoothing we obtain the decomposition of inflation indicators into permanent (i.e. stationary possibly with a single break) and transitory components:

$$\pi_{it}=100*[(\tau_{it}-\tau_{it-1})+(c_{it}-c_{it-1})], \quad (5)$$

In this setup shocks to headline and XUFE inflation are possibly correlated between the series and between permanent and transitory (i.e. mean-reverting) components. The bUCC model without linear restrictions on cointegration could potentially lead to divergent long-term dynamics of HICP and XUFE inflation unless their transitory components negatively interact with permanent ones. This offsetting effects has been found by Bradley, et al. (2013) in the dynamics of US inflation over the period 1984-2012. The effect comprises of an adverse shock in the transitory components after a one-off shock in a permanent component and it preserves headline and core inflation series from long-run divergence. We search for a preferred model decomposition among specifications with or without cointegration[6], and with or without a single break in each of univariate processes. In the case of structural breaks in the long-run inflation dynamics allowing for correlation between innovations to permanent and transitory components is even more important for equilibrium restoring mechanism than in the case of no cointegration.

[6] Cointegration in bUCC model is defined by two linear restrictions: $\eta_{1t}=\gamma\eta_{2t}$ and $\mu_1=\mu_2$.

4. Results

To take general insights from the analysis we approach each country inflation rates with an individually selected bUCC model. We follow such a modelling strategy. First, we estimate bUCC model (1)-(4) with cointegration constraints ($\eta_{1t} = \gamma\eta_{2t}$ and $\mu_1 = \mu_2$) and no breaks ($\mu_{1,t_{br}} = \mu_{2t_{br}} = 0$), which altogether denote common trends in permanent components of headline inflation and XUFE. If there is a cointegration the constrained model of common trends (with 11 parameters overall) should be strongly preferred in likelihood ratio tests over the model without cointegration and breaks ($\mu_{1,t_{1,br}} = \mu_{2t_{2,br}} = 0$, with 18 parameters). As in Bradley et al. (2015) analysis on US inflation rates for the period 1993-2012 we do not find any support for common trends (i.e. cointegration) between core and headline inflation in the EU countries[7] (second column in Tab. 5 in Appendix). Consequently, headline and XUFE inflation may diverge from each other for years without any error-correcting mechanism which would attract headline inflation to the long run underlying trend.

Breaks

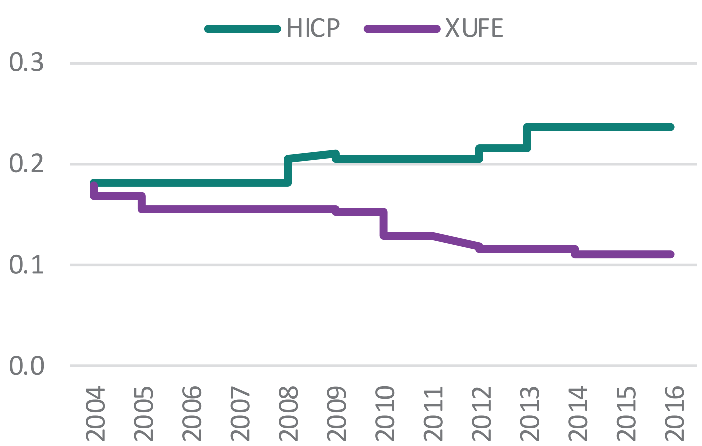
After rejecting common trends hypotheses in an unrestricted model we test the following single-break restrictions in the long-run equations of bUCC model $\mu_{1,t_{1,br}} = \mu_{2t_{2,br}} = 0$ (Tab. 5, left panel). One can obtain different long-run dynamics of inflation rates depending on the date of the breaks, whether they are common ($t_{1,br} = t_{2,br}$) for the two series or specific ones ($t_{1,br} \neq t_{2,br}$). Firstly, we introduce a common break in January 2009 and check (with LR test) whether the difference in the maximum likelihoods between the model with and without common break is statistically significant (Tab. 5, third column). Then, we also apply LR inference to the model of country specific breaks which are based on the outcomes of Zivot-Andrews tests. The finally selected (individual or common) breaks in the trends of headline and XUFE

[7] The result for Romania is non-conclusive. In the model with cointegration and no breaks we failed to obtain numerically stable maximum of log-likelihood function.

indices are presented in the right panel of Tab. 5. We find that individually specified dates are preferable over other alternatives in bivariate price dynamics of 10 countries, common breaks in January 2009 are acceptable in 10 countries, and in the rest of the EU countries no significant breaks in bivariate dgps are found.

To present synthetically information on individually selected breaks long-run trend components of HICP and XUFE ($\mu_i + \mu_{i,t_{i,br}} I_t$) are separately aggregated across bUCC country models (with median) into a monthly time series. We find that after the outbreak of the global financial and sovereign debt crises in EU the divergence between HICP and core inflation widened on average (see Fig. 2). The divergence may support the hypothesis that either the structural change (e.g. growing importance of the e-commerce) in the relation of these two indices occurred or that shocks that hit inflation in recent years were not temporary in nature (cf. Ciccarelli and Osbat, 2017).

Fig. 2 The divergence in the long-run dynamics of headline and XUFE



Note: The divergence is estimated according to equations (1)-(4) as a difference between cross-sectional medians of the long-run levels ($\mu_i + \mu_{i,t_{i,br}} I_t(t_{i,br})$) of headline HICP and XUFE inflation rates.
Source: Own calculations from the bUCC model.

Model-based decomposition

Based on the results of the model decomposition we compare different characteristics of permanent components (pHICP, pXUFE) with a benchmark indicator of an underlying inflation (XUFE). and evaluate them from the point of view of monetary policy. Based on the rich literature (cf. Roger, 1997, Wynne, 1999, Clark, 2001, Marques et al., 2003, Silver, 2007) we take into account the following features of ideal core inflation measures:

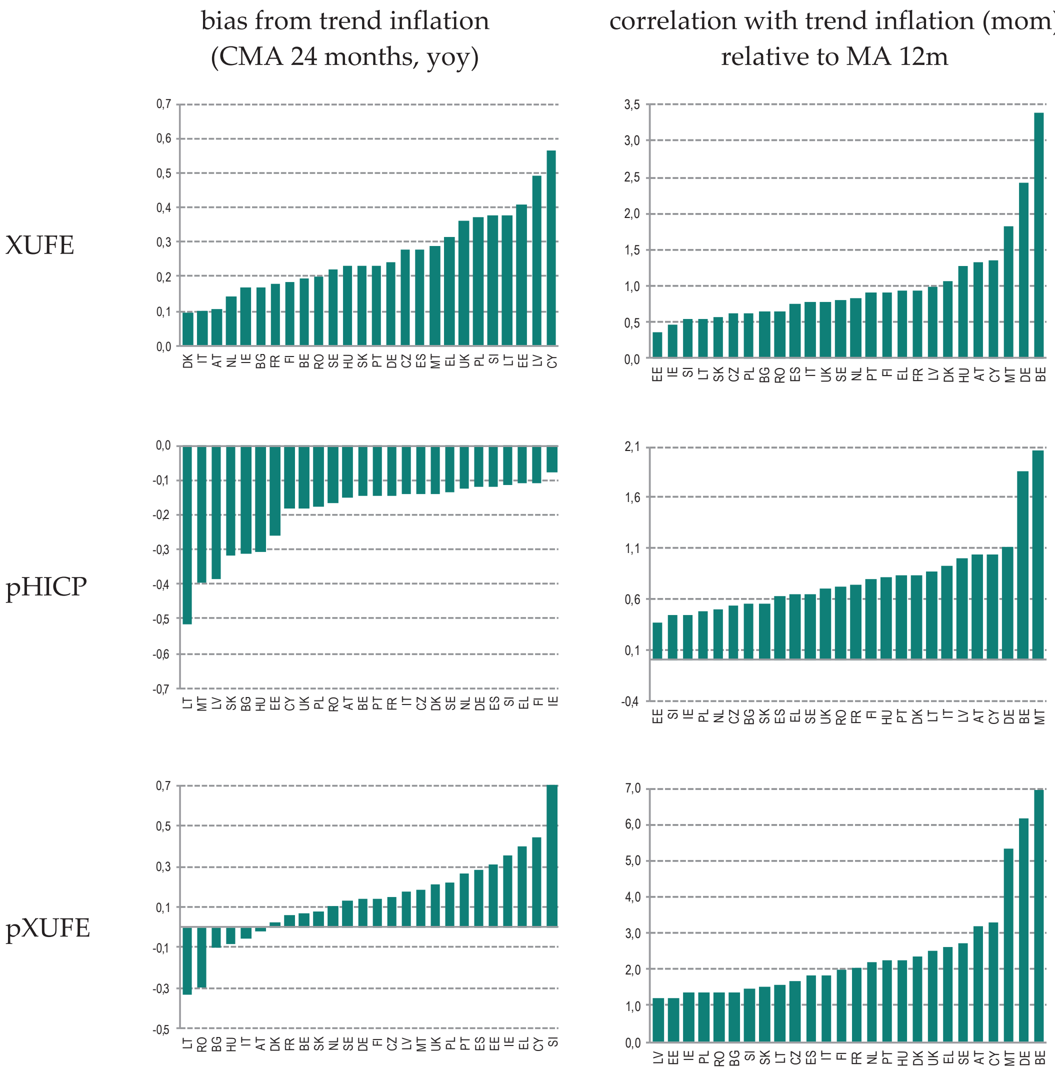
- tracking trend,
- appropriate volatility reduction,
- no systematic bias and no long-run divergence from the cost-of-living index (HICP all items),
- good indicator of demand pressure (e.g. index sensitive to the output gap in the Phillips curve equations),
- leading indicator for headline inflation i.e. good trend forecasting abilities (in comparison with random walk forecasts).

Then we compile these performance records into a synthetic indicator, named Core Inflation Score (CIS). Below we describe these criteria in more details, formulate the benchmarks for their quality and comment the results.

<u>Tracking trend</u>

Extreme shocks to observed inflation push it away from the equilibrium temporarily. Policy makers prefer to eliminate these shocks from aggregate price index to better track the smooth long-run component of inflation. The exclusion of considerable part of consumption basket, however, may systematically bias the inflation trend. Overstating or understating the true long-run inflation may provide spurious signals for monetary policy authorities. While excluding volatile components from the basket (e.g. energy prices) it is often the case that the price changes of excluded components diverge from headline inflation in a systematic way, particularly when positive shocks to oil prices prevail. Such an exclusion-based index may lack tracking trend quality.

Fig. 3 The bias and correlation of CMA trend with: XUFE, permanent HICP (pHICP), permanent XUFE (pXUFE)



Note: Left panel indicates the bias of the analysed measures from the approximation of trend inflation as measured with 24-months centered moving average. Positive (negative) number indicates underestimation (overestimation) of the trend inflation. Right panel indicates correlation of the selected measures with the same trend inflation measure.

Source: Own calculations.

Comparison of the inflation trend estimates (proxied by 24-months centred moving average, CMA 24m, cf. Bryan et al., 1997, Cecchetti, 1997, Clark, 2001), which are available *ex post* and pseudo real-time estimates of XUFE reveal that inflation index ex. energy and unprocessed food systematically underestimates trend in every EU country in the last 4 years under investigation (Fig. 3 left panel). The bias is moderate, on average below 0.3 pp. in y-o-y terms, it is the biggest for Cyprus and Baltic countries. As for the model-based candidates for underlying inflation, the permanent component of XUFE (pXUFE) underestimates trend in inflation for most of the EU countries, though this bias is twice smaller than in original XUFE. Contrarily, permanent component of HICP (pHICP) overestimates the trend in every EU country (0.20 pp. on average).

Trend tracking performance (CMA 24m) is also compared in terms of correlation to a simple *ex post* benchmark i.e. one-sided 12-months moving average (MA 12m). Permanent components of XUFE is again the best trend tracking tool in most of the countries (Fig. 3 right panel). The correlation of pXUFE with a trend for the analysed countries is 0.81 on average, with Malta and Belgium (correlations below 0.66) being the only outliers. Hence, the correlations of pXUFE with trend are on average more than twice bigger than correlations of the real-time benchmark (i.e. MA 12m). Even in the case of Malta and Belgium, where trend tracking with MA method is really poor, the increases in trend tracking performance are substantial (correlations with a trend more than 5 times bigger than the correlations of MA). Both XUFE and permanent component of HICP record much worse trend-tracking performance. In many countries a simple moving average performs better than the exclusion-based (XUFE) and model-based (permanent HICP) measures[8]. To sum up permanent

[8] The value below 1 in Fig. 3 right panel indicates that the correlation of the analysed index with the trend inflation is smaller than the correlation of the benchmark i.e. simple moving average.

component of XUFE is, on average, the best tool for tracking inflation trend in terms of correlation and unbiasedness.

Volatility reduction versus trend approximation

Another important feature of the ideal core inflation measure is a reduction in the volatility of the series when compared to the volatility of headline inflation. A common knowledge is that XUFE tends to dampen headline inflation volatility as unprocessed food and energy are relatively more volatile than other components. The evidence for XUFE indices across all EU countries shows that the average extent of inflation dampening in terms of standard deviation of monthly inflation is about 20% (Tab. 2 and Fig. 4 left panel). In most of the analysed countries (except for Malta) the volatility of the model-based permanent indicators is higher than the volatility exclusion-based inflation. Our comparative multi-country analysis indicates that on average the volatility of the permanent components of HICP and XUFE are bigger than the volatility of original XUFE, by 81% for pHICP and by 32% for pXUFE, respectively (Tab. 2). These are direct consequences of high and negative correlations between permanent and transitory components of XUFE (offsetting effect). Also permanent shocks to HICP-all items index and transitory shocks to XUFE are predominantly negatively correlated i.e. persistent shocks are followed by equilibrium restoring short-term fluctuations which offset initial shock. There is an analogous observation about the comovements in US inflation rates formulated by Bradley et al. (2015).

Reducing the volatility of observed inflation, underlying inflation measure should not at the same time diverge from the *ex post* measure of underlying trend inflation. Yet, all of the analysed indices (XUFE, and model-based pHICP and pXUFE) diverge from the trend (CMA24m) substantially (see Fig. 4, right panel). None of them has an advantage over the other in this respect.[9] The choice of less divergent optimal trend indicator vary between the countries and indicators without any regularities.

Tab. 2 Descriptive statistics of XUFE and two model-based underlying inflation measures: permanent HICP (pHICP), permanent XUFE (pXUFE), and their correlations with HICP and transitory components (tHICP, tXUFE) in percentage points

	mean	st. dev.	st dev. (vs. HICP)			correlations of shocks			
country	HICP (mom)		XUFE	pHICP	pXUFE	XUFE pXUFE	HICP pHICP	tXUFE pHICP	pXUFE tHICP
AT	0,15	0,18	0,78	1,43	0,92	0,26	0,90	-0,95	-0,95
BE	0,16	0,26	0,55	1,01	0,38	0,16	0,96	-0,96	-0,96
BG	0,29	0,55	0,84	3,02	2,97	0,70	0,84	-0,99	-0,99
CY	0,14	0,38	0,71	1,12	0,77	0,70	0,98	-0,20	-0,20
CZ	0,15	0,32	0,75	1,83	1,56	0,84	0,92	-0,99	-0,99
DE	0,12	0,20	0,72	1,31	0,63	0,39	0,95	-0,92	-0,92
DK	0,13	0,21	0,74	0,93	1,14	0,79	0,98	-0,43	-0,43
EE	0,27	0,39	0,75	2,37	1,91	0,65	0,87	-0,98	-0,98
EL	0,17	0,30	1,28	1,79	1,06	0,95	0,92	-0,88	-0,88
ES	0,17	0,23	0,79	1,61	0,86	0,59	0,90	-0,94	-0,94
FI	0,14	0,20	0,80	1,67	1,70	0,57	0,94	-0,93	-0,93
FR	0,13	0,18	0,62	1,09	0,85	0,46	0,94	-0,87	-0,87
HU	0,33	0,27	1,12	2,26	2,18	0,43	0,81	-0,93	-0,93
IE	0,13	0,26	0,87	1,93	2,28	0,76	0,85	-0,94	-0,94
IT	0,15	0,19	0,70	1,68	1,09	0,79	0,95	-0,99	-0,99
LT	0,21	0,43	0,79	2,61	1,96	0,75	0,89	-0,97	-0,97
LV	0,32	0,49	0,92	1,19	0,78	0,84	0,99	-0,39	-0,39
MT	0,17	0,38	0,92	0,58	0,03	-0,63	0,98	0,69	0,69
NL	0,14	0,21	0,75	1,55	1,37	0,69	0,93	-0,77	-0,77
PL	0,17	0,24	0,64	2,52	2,32	0,76	0,79	-0,96	-0,96
PT	0,16	0,26	0,88	1,13	1,23	0,49	0,98	-0,99	-0,99
RO	0,58	0,66	0,90	1,80	1,14	0,55	0,38	-0,98	-0,98
SE	0,12	0,22	0,67	0,84	0,83	0,96	0,99	-0,57	-0,57
SI	0,23	0,36	0,78	4,18	2,09	0,28	0,89	-0,66	-0,66
SK	0,24	0,52	0,57	3,82	1,34	0,82	0,96	-0,98	-0,98
UK	0,17	0,20	0,70	1,70	0,96	0,64	0,93	-0,99	-0,99

Source: Own calculations based on month-over-month indices and bUCC model.

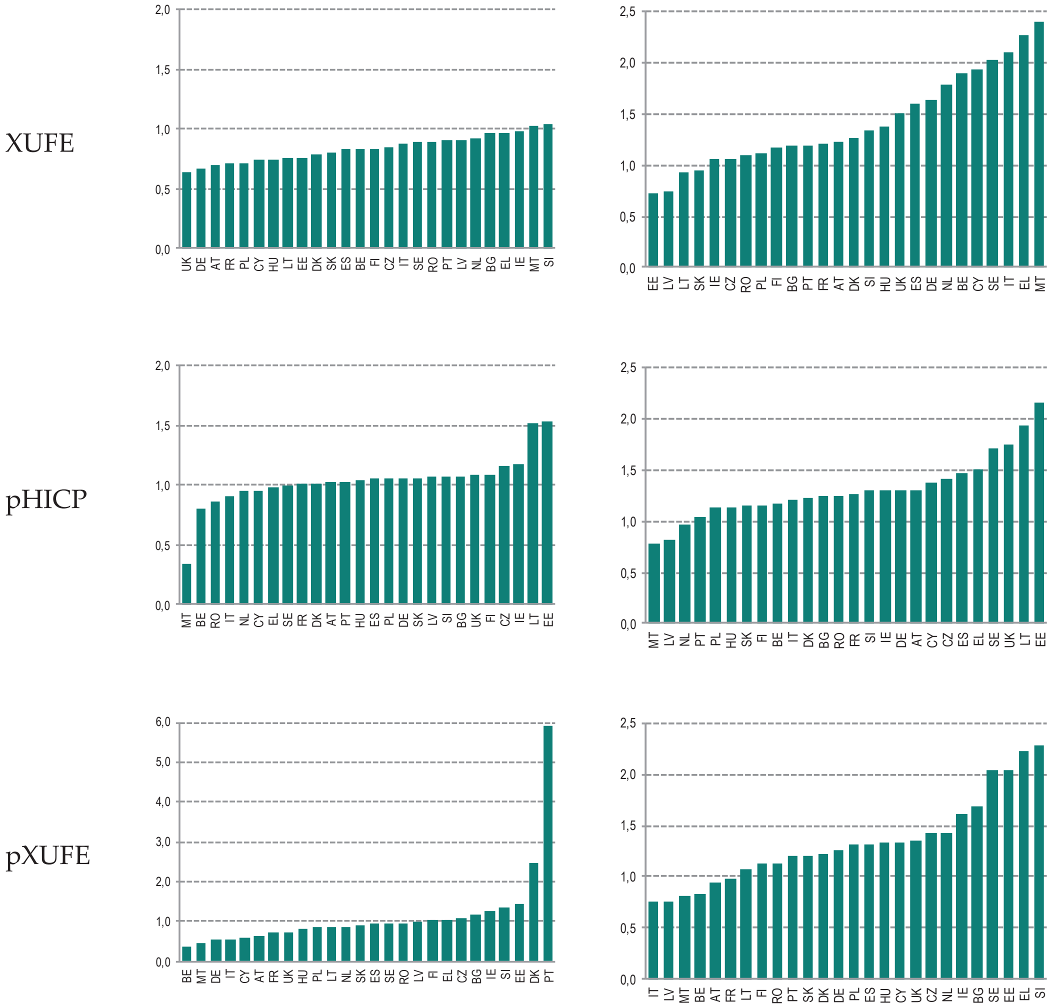
Summing up, an ideal core inflation measure should be as smooth as possible but not constant, as the volatility is an important input for approximating current trend and

[9] The divergence is measured in pp. as the root of the average squared differences of the indicator record from the underlying inflation trend approximation (CMA24) relative to a real-time (MA12) one-sided benchmark.

predicting future changes in trend inflation. It seems that by excluding unprocessed food and energy from HICP some part of the permanent component may also be excluded. Such outcome suggests that smoothing inflation by exclusion method may be excessive.

Fig. 4 The volatility reduction and divergence of core inflation measures from trend: XUFE, permanent HICP (pHICP), and permanent XUFE (pXUFE).

volatility reduction relative to HICP std(X)/std(HICP) in mom terms	divergence (RMSE) from trend (CMA 24 months) relative to MA 12m



Note: Left panel indicates the volatility reduction relative to the volatility of HICP. The numbers below 1 indicates lower volatility of the selected measure. Right panel indicates RMSE of the selected measures from the 24-months centered moving average in relation to the 12 month moving average.

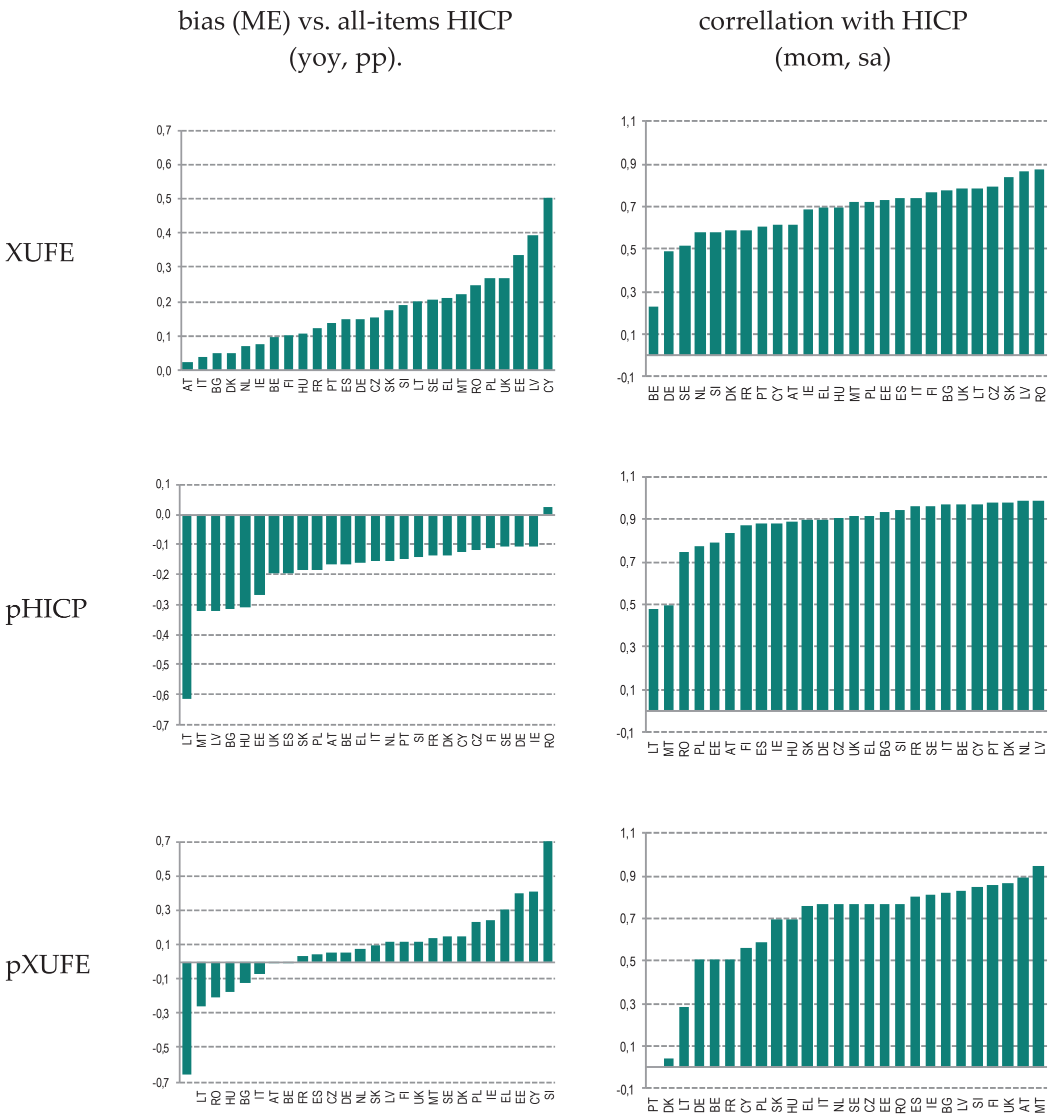
Source: Own calculations.

Cost of living approximation

The exclusion of considerable part of consumption basket (as it is made in exclusion based core inflation measures) may also lead to a systematic error in the evaluation of changes in the cost of living index. This bias is a disadvantage from a general viewpoint as the policy indicator should bring relevant information not only for economic models but also for the general public. Therefore, it would be desirable that proposed core inflation measure is on average not far from the closest approximation to the cost of living index (COLI) i.e. headline HICP. The significant positive correlation with COLI is also an important complementary requirement.

In the samples of EU countries the measure of core inflation with the smallest bias to COLI is pXUFE. Some of the country pXUFE indices tend to underestimate and some overestimate observed headline inflation (see Fig. 5, left panel). For most of the countries original XUFE indices underestimate HICP country inflations. Contrarily, the measures based on the permanent components of HICP (pHICP) tend to overestimate it. Yet, the bias for most of the countries is small (less than 0.3 pp.) and according to the t-ratio test it is not statistically significant. From the perspective of correlation with headline inflation (Fig. 5, right panel), pHICP indices seem to be the best choice for an approximation of COLI index with an average cross-country correlation at 0,88. The other two (XUFE and pXUFE) have comparable record – close to 0,68 on average. There are some countries with really low correlations of one of these measures with headline HICP, which suggests that 'one size fits all' strategy is not a best approach for the choice of COLI.

Fig. 5 The bias and correlation of HICP with XUFE, permanent HICP (pHICP), and permanent XUFE (pXUFE)



Note: Left panel indicates the bias of the selected measure (ME) from the all-items HICP index (in yoy terms). Right panel indicates correlation of the selected measures with the HICP-all-items (mom, sa).

Source: Own calculations.

Demand Pressure

To evaluate if core inflation indicator is a good approximation of the demand pressure we build regression equations based on the Phillips curve (PC) relationship for each country. We follow the specification of the backward-looking PC proposed by Stock and Watson (2009). However, as they analyse closed economy (US) they do not include exchange rate variables nor the foreign output gaps. In our case we analyse both closed and open economies versions of PC equations, the second ones also enhanced with exchange rate and global output gap[10] variables:

$$\pi_t^{core} = \alpha_1 * \pi_{t-1}^{core} + \alpha_2 * \pi_{t-2}^{core} + \alpha_3 * gap_t^{c|w|c,w} + \alpha_4 * EER_t^{nom|real} + \beta c_t + \varepsilon_t, \quad (6)$$

where: π_t^{core} denotes selected core inflation indicator (XUFE, pXUFE and pHICP),

gap_t^x is an output gap (deviations from HP-detrended log real GDP), either domestic ($x = c$) or global output gap measure ($x = w$), or both ($x = c, w$),

EER_t is an effective exchange rate (nominal or real),

c_t are different control variables (prices of energy, food commodities and raw materials).

The equation (6) may be further augmented with inflation expectations, better proxies for output gap (unit labour costs, unemployment rates, country specific external gap) and more elaborated dynamics of inflation and output gap. Yet, for the purpose of this cross-country study, and because of the comparability (different lag structures for each country), ease of exposition (IV with proper instruments would be necessary for forward looking versions of Phillips curve), the lack of comparable data for the whole analysed period (especially for late acceding EU countries and HICP inflation expectations) we have limited the analysis to the backward-looking version presented in equation (6). All in all we estimate 30 Phillips curves with domestic output gap and 30 with global output gap in a variety of specifications i.e., with or without the other output gap, with or without different effective exchange rates (nominal or real), as well as with or without additional control variables.

[10] Calculated as an output gap for all European OECD countries.

The outcomes of the cross-country study indicate that in two thirds of the euro zone countries and half outside the zone (the Czech Republic, Denmark, Poland and United Kingdom) XUFE is sensitive to a contemporaneous domestic demand factor. There is no common characteristic for the countries in which inflation is output insensitive and it seems that insensitiveness of inflation to the domestic output gap is rather country specific. For example, in Hungary it may stem from rather frequent changes in administered prices and tax rates. The possible explanation for Sweden may by its high openness and relative high importance of the global commodity shocks in shaping the inflation (see Hałka and Kotłowski, 2016). In case of e.g. Portugal it may arise from the deep economic crisis not followed by an adequate drop of prices (due to prise rigidities or inability of the internal devaluation of the currency). Altogether, the evidence from 30 different PC specifications, XUFE and pXUFE, compared to pHICP, are more sensitive to both domestic and global output gap measures (Tab. 2), however the performance of these gaps substantially differs between the countries. For example, in Slovenia, Czech Republic and Baltic countries global output gap is more important than domestic factors. Contrarily, in Poland, Greece, Slovakia, and Cyprus only domestic output gaps are significant real factors explaining inflation dynamics. In general, the results indicate surprisingly that XUFE is rather a good indicator of the demand pressure in the economy[11].

[11] In the literature one can find inflation indices that capture demand pressure better than the exclusion based core indicators – see Fröhling and Lommatzsch (2011) for euro area countries, and Hałka and Kotłowski (2014) for Poland.

Tab. 3 The percentage of statistically significant (p-value<0.05) coefficients at the output gap variable in 30 specifications of backward-looking Phillips curves – equation (6) with different inflation measures.

	country gap			OECD gap		
country	XUFE	perm XUFE	perm HICP	XUFE	perm XUFE	perm HICP
AT	70	60	30	60	30	30
BE	50	40	90	100	30	40
BG	0	0	0	0	10	50
CY	20	20	70	0	0	0
CZ	60	60	60	100	100	60
DE	60	60	30	60	50	20
DK	60	50	0	40	0	0
EE	60	60	10	100	100	100
EL	100	100	100	0	0	0
ES	0	0	0	0	0	0
FI	40	0	0	0	0	0
FR	0	0	0	0	0	0
HU	0	0	0	0	0	0
IE	0	0	0	0	50	20
IT	60	0	0	0	0	0
LT	60	60	0	100	100	100
LV	50	40	20	60	50	80
MT	0	60	0	0	60	0
NL	0	0	0	0	0	0
PL	100	80	70	0	0	0
PT	0	0	0	0	0	0
RO	0	0	0	0	0	0
SE	0	0	0	0	0	0
SI	0	0	0	40	90	60
SK	30	10	0	0	0	0
UK	10	10	10	10	10	10

Source: Own calculations.

Forecasting performance

To evaluate the forecasting performance of the underlying inflation indicators we use the approach of Bermingham (2010). The general idea of the simple test relies on the assumption that the gap between core and headline inflation has some forecasting power for future evolution of headline inflation. Bermingham test is based on direct forecasting rules for different forecast horizons ($h = 1,2, \dots, H$):

$$\hat{\pi}_{t+h} - \pi_t = \hat{\alpha}_h + \hat{\beta}_h \left(\pi_t^{core} - \pi_t\right) \quad (7)$$

where π_t is HICP inflation in year-over-year (yoy) terms and, $\hat{\pi}_{t+h}$ is its forecast with h-months ahead time horizon, and π_t^{core} stands for different core inflation indicators: XUFE , pXUFE and pHICP (yoy). The out-of-sample forecasts are obtained using different pairs of recursively estimated parameters ($\hat{\alpha}_h, \hat{\beta}_h$) in 12 forecasting models each for any of forecast horizons up to 12 months ahead ($h = 1,2, \dots, 12$). The estimates $0 < \hat{\beta}_h < 1$ describe the rate of convergence of headline inflation (in a forecast horizon h) toward underlying inflation indicator.

The outcomes of the forecasting exercise for different indices are compared with each other as well as with the benchmark model with $\hat{\beta}_h = 0$ (being a random walk with a drift model) by using statistical test of equal forecast accuracy by Harvey et al. (1997) and standard t-ratio test for the forecast bias. The forecast errors are calculated and evaluated over the 48-months out-of-sample verification window.

The results of the forecasting exercise (Tab. 6 and Tab. 7 in Appendix) indicate that, as expected, XUFE is a weak predictor of headline inflation. Both permanent components of HICP and XUFE perform better in this respect. With an exception of three countries (Bulgaria, the Czech Republic, and Romania) the permanent component of HICP helps to forecast inflation in the short run (up to 3 months ahead), but only in 15 country cases the difference in RMSFE is statistically significant. The results of 3-months ahead forecasting exercise on pXUFE are slightly worse. Short-run forecasts of headline inflation with pXUFE are worse than the benchmark in 7 countries and only in 6 countries the difference in RMSFE is statistically significant.

For longer time horizons (including 12 months ahead) we can observe that forecasts based on permanent component of XUFE tend to perform better than others. In seven out of the EU countries the improvement over RW benchmark in RMSFE terms is statistically significant. Neither XUFE nor pXUFE provide significant forecast improvements. Finally, the benchmark model and XUFE are worst in forecasting inflation both in the short and in the medium term.

For monetary policy purposes the most important forecasting horizon is medium to long term. In our case it will be the last forecasted one – 12 months. The models based on the permanent components of HICP and XUFE for Sweden and Ireland give particularly good results, with a RMSFE below 1.0. For Austria, Germany, France and Denmark the RMSFE is still below 1.5. The worst outcomes are for the Hungary[12] with the RMSE more than three times bigger than for the Sweden – over 3 and also for Poland, Estonia and Bulgaria (over 2.5). For the rest countries RMSFE is between 1.5 and 2.5. The outcomes indicate that both model-based measures predict headline inflation with some precision but the difference between errors in many cases is statistically insignificant. Forecasts using XUFE generally give worst results of all examined core inflation indicators.

For most of the EU countries the least biased forecasts are forecasts based on the HICP permanent component with an exception of six small open economies belonging to the euro zone and Sweden. For the shortest horizons the forecasts biases are more or less uniformly distributed around zero. But when the horizon becomes longer the bias becomes negative for each country and of a bigger magnitude. One of the possible explanation is that evaluation period covers the time span with unexpected drops of oil prices what caused overestimation of the forecasts, especially in the longer horizon. However, the best models in terms of RMSFE are not always the best models in terms of bias as the less biased model for the longest horizon is for Greece (0.02) followed by Sweden, Belgium and Portugal (bias below 0.5). Still there are forecasts with bias over 1.5 – Estonia, Cyprus, Latvia and Finland.

To conclude Bermingham (2010) forecasts based on pHICP give both smaller RMSFE and smaller bias in comparison to other models in short-term horizons. However, the discrepancies between RMSFE and bias for the analysed countries may be substantial. For the longer term the forecasts based on pXUFE are better, however being also considerably upward biased.

[12] The reason why for Hungary all models have the highest error may be twofold. Firstly Hungary in recent years faced several administrative decisions that influenced the price dynamics like changes in the VAT as well as in regulated prices. Secondly, due to unexpected drop of the oil prices in 2014 and 2015 the error amplify.

Core Inflation Score

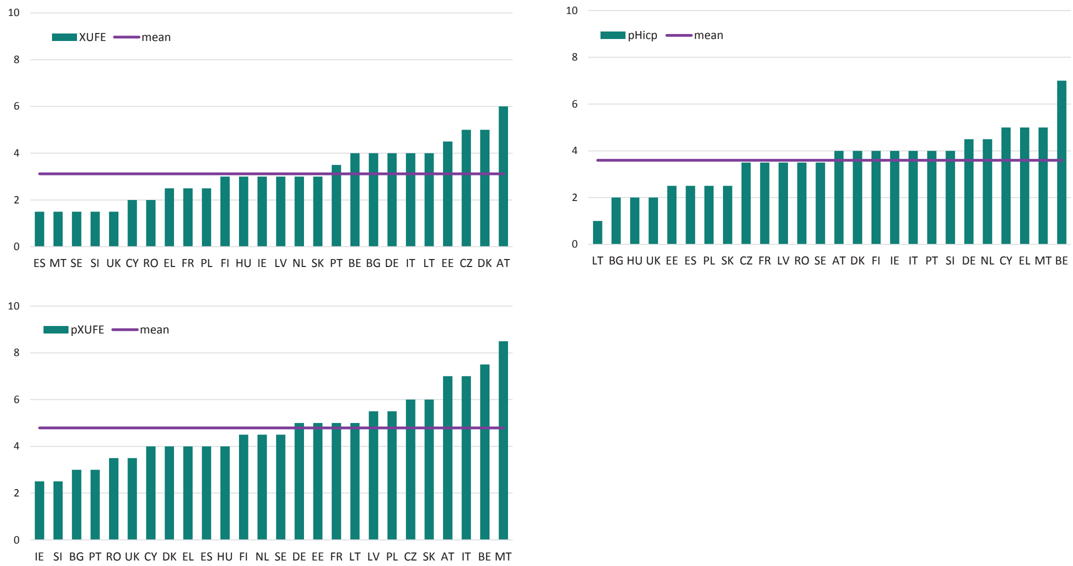
Lastly, we propose a synthetic measure CIS (Core Inflation Score) to evaluate the usefulness of the measures of underlying inflation based on the five criteria proposed in the paper. Depending on the index and proportionally spread thresholds we assign 0, 0.5, 1, 1.5 or 2 points for each criteria. Hence, the ideal CIS could gain maximally 10 points.

The analysis of these criteria shows that XUFE perform the worst as an underlying inflation measure both in terms of average CIS (3.1 across countries, see Fig. 6) and in terms of the number of countries where it is a preferred measure (two countries: Denmark and Bulgaria, see Fig. 7 in Appendix). The average CIS of pHICP (3.6) and pXUFE (4.8) across all countries is considerably bigger than XUFE (Fig. 6). Following these criteria pXUFE index is as good as or better than any other core inflation measure in 19 countries, and pHICP in 6 countries. Still none of the underlying inflation measures gets on average more than 50% of the CIS points. The CIS of pXUFE is equal or bigger than six in 6 out of 26 countries (Austria, Belgium, the Czech Republic, Italy, Malta, Slovakia), whereas CIS of pHICP and XUFE is equal or bigger than 6 only for one country (Belgium and Austria, respectively).

On average CIS is higher for the euro zone countries. This outcome is especially visible for two criteria: volatility reduction and forecasting performance. When the selected 5 criteria indicate that pXUFE is the best measure of core inflation when it comes to the trend tracking and volatility reduction. pHICP has better forecasting performance and it is the best approximation of the COLI index. Finally, the only advantage of XUFE refers to the demand pressure approximation, however we have to bear in mind that, as other research shows, it is not the best demand tracking measure.

Summing up, none of the selected indicators is the ideal core inflation measures and therefore for the monetary policy authorities it is important to focus on bunch of different indicators based on exclusion methods and models instead on one selected measure.

Fig. 6 The Core Inflation Score (CIS) for XUFE, permanent HICP and permanent XUFE component



Source: Own calculations.

5. Conclusions

There are some important insights from our analysis for the monetary policy and future research in the topic.

Firstly, after the outbreak of the global financial crisis the divergence between HICP-all items and XUFE widened, thus XUFE no longer reflects long term HICP. Breaks in the bivariate dynamics may be, among others, a result of the structural changes on the world energy market.

Secondly, although the energy and unprocessed food prices are volatile they also include part of the permanent component of inflation. Therefore, the measure that excludes these components is not the best approximation of an underlying inflation trend, in particular in the period of the long-lasting commodity shocks.

Thirdly, despite of the economists' opinion that the core inflation measures should reflect the trend inflation rather than the fluctuation of the demand pressure we find a positive correlation between XUFE and output gap for many countries. Moreover, for a number of countries XUFE is a better indicator of the demand pressure in the estimated backward-looking Phillips curve than of the trend inflation.

Fourthly, in short-term forecasting exercise across EU countries the permanent component of HICP may be a better and less biased predictor of headline inflation (HICP) than any other core inflation indicator. At the same time for longer horizons the permanent components of XUFE often overperform other indicators.

Fifthly, none of the proposed indices seem to be an ideal measure of the core inflation. Each of the selected indicators provides valuable insights into how the inflation is evolving. XUFE is the best when it comes to the demand pressure (with the reservation that in literature we can find better measure of the demand pressure), pHICP offers the best forecasting performance and best approximation of the cost of living index. On the other hand, and pXUFE is the best core measure when trend tracking and volatility reduction is concerned. Therefore, monetary policy authorities rather than concentrating on one favourite measure should look at wide spectrum of different price indices to better understand the inflation behaviour.

Finally, Core Inflation Score (CIS) indicates that model-based permanent component of XUFE is the preferred overall measure of core inflation in the most of EU countries in terms of analysed the multiple criteria for the ideal core inflation indicator.

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Appendix

Tab. 4 Unit-root and stationarity tests for XUFE inflation series (logarithmic changes in price indices)

HICP ex unprocessed food and energy (XUFE)						
DF-GLS		KPSS		Zivot-Andrews		conclusions
t-stat	I(1)	LM	I(0)	break	p(val)	on stationarity
(1)		(2)		(3)		(4)
-0,54		0,11		2005m4	<0.01	break
-2,56	**	0,06		2010m7	<0.01	I(0)
-3,88	***	0,78	***	2010m4	<0.01	I(0)
-12,86	***	0,76	***	2003m1	<0.01	I(0)
-3,17	***	0,12		2008m1	<0.01	I(0)
-1,73	*	0,08		2010m4	<0.01	break
-12,39	***	0,62	**	2012m1	<0.01	I(0)
-2,20		0,24		2008m1	<0.01	break
-1,63	*	1,51	***	2011m10	<0.01	break
-1,10		1,28	***	2009m3	<0.01	break
-0,60		0,17		2007m12	<0.01	break
-2,83	***	1,01	***	2014m2	<0.01	I(0)
-0,43		0,44	*	2009m7	<0.01	break
-2,47	**	0,90	***	2008m2	<0.01	I(0)
-1,86	*	1,22	***	2012m6	<0.01	break
-5,67	***	0,19		2008m4	<0.01	I(0)
-2,30	**	0,39	*	2009m3	<0.01	I(0)
-15,82	***	0,20		2004m11	<0.01	I(0)
-0,53		0,30		2012m10	0,18	no
-1,29		1,24	***	2011m12	0,02	break
-0,93		1,24	***	2004m6	<0.01	break
-0,97		1,29	***	2004m2	<0.01	break
-15,54	***	0,28		2009m7	<0.01	I(0)
-1,20		1,09	***	2003m6	<0.01	break
-1,65	*	1,08	***	2004m1	<0.01	break
-3,81	***	0,36	*	2008m12	<0.01	I(0)

Note: The same remarks as in Tab. 1 apply.

Source: Own calculations.

Tab. 5 Statistical results of selection procedure of the bUCC models for EU countries: tests for cointegration restrictions, common and specific breaks, and the time and size of breaks in the long-run mean of HICP and XUFE series

	LR tests in UCC model				Breaks specification			
country	H0: cointegration		H1: I(0)+break		headline	$\mu_{1,br}$	XUFE	$\mu_{2,br}$
AT	541,70	***	6,87	**	2012m9	0,09	2005m4	-0,06
BE	617,84	***	8,99	**	2008m6	0,11	2010m7	-0,08
BG	291,59	***	2,61			0,00		0,00
CY	626,26	***	6,29	**	2009m1	-0,47	2009m1	-0,14
CZ	483,60	***	0,17			0,00		0,00
DE	698,66	***	2,15			0,00		0,00
DK	413,75	***	12,79	***	2012m8	0,06	2012m1	-0,07
EE	528,41	***	6,68	**	2009m1	-0,05	2009m1	-0,33
EL	234,08	***	29,60	***	2009m1	-0,89	2009m1	-0,17
ES	772,77	***	8,27	**	2009m1	0,11	2009m1	-0,16
FI	430,99	***	2,41			0,00		0,00
FR	765,33	***	11,79	***	2008m11	0,10	2014m2	-0,10
HU	28,18	***	36,51	***	2012m1	-0,11	2009m7	-0,13
IE	621,99	***	12,47	***	2008m6	-0,07	2008m2	-0,13
IT	518,81	***	13,16	***	2013m2	0,07	2012m6	-0,11
LT	2 917,09	***	1,58			0,00		0,00
LV	622,62	***	50,72	***	2009m2	-0,23	2009m3	-0,28
MT	337,38	***	13,68	***	2009m1	0,13	2009m1	-0,10
NL	499,76	***	4,20			0,00		0,00
PL	700,91	***	10,03	***	2012m6	-0,14	2011m12	0,00
PT	367,97	***	26,83	***	2012m1	0,22	2004m6	-0,19
RO	NA		188,47	***	2010m9	0,00	2004m2	0,13
SE	940,58	***	14,49	***	2009m1	0,34	2009m1	-0,07
SI	489,26	***	7,47	**	2009m1	-0,09	2009m1	0,03
SK	574,82	***	12,40	***	2009m1	0,20	2009m1	-0,06
UK	672,53	***	11,81	***	2009m1	0,15	2009m1	0,04

*Note: LR tests stand for statistics of likelihood ratio tests for cointegration (without breaks) and breaks restrictions. The benchmark (comparison) model is always bUCC model without cointegration and breaks. ***, **, * denote respectively 1%, 5% and 10% significance level of rejecting the null hypotheses.*

Source: Own calculations.

Tab. 6 RMSFE from the Bermingham (2010) type forecasts of headline inflation rates using 3 underlying inflation indicators (XUFE, pHICP, pXUFE) compared with a random walk (RW) benchmark

	benchmark RW		Forecasts from Bermingham type model (horizon h)					
			XUFE		pHICP		pXUFE	
	h=3	h=12	h=3	h=12	h=3	h=12	h=3	h=12
AT	0.92	1.28	0.93	1.62	0.72 ***	1.42	0.85	1.38
BE	1.90	2.28	1.91	2.58	1.21 ***	2.34	1.51 ***	1.80 ***
BG	1.24	2.47	1.25	2.36 ***	1.24	2.55	1.22	2.75
CY	2.02	2.35	2.04	2.57	1.48 ***	2.39	1.71	2.32
CZ	0.72	1.94	0.72	1.71	0.72	1.97	0.69	1.85
DE	0.70	1.23	0.72	1.43	0.61 **	1.28	0.70	1.29
DK	0.64	1.12	0.65	1.27	0.55 **	1.18	0.64	1.11
EE	1.10	2.72	1.12	2.70	1.04	2.79	0.94	2.46 **
EL	1.95	2.43	1.98	2.39	1.51 ***	2.52	1.34 ***	1.86 **
ES	1.57	2.20	1.61	2.40	1.29 ***	2.38	1.56	2.20
FI	0.65	1.78	0.65	1.87	0.55	1.78	0.57 **	1.79
FR	0.79	1.09	0.80	1.21	0.63 ***	1.32	0.70	1.03
HU	1.23	3.12	1.27	3.22	1.10 **	3.14	1.25	3.18
IE	0.81	0.95	0.87	1.00	0.80	0.94	0.79	0.95
IT	1.73	2.28	1.71	2.24	1.19 ***	2.37	1.27 ***	1.96 ***
LT	0.98	2.10	1.03	2.69	0.97	2.11	1.05	1.99
LV	1.14	1.20	1.14	1.61	1.12	1.84	1.20	1.17
MT	2.59	1.95	2.54	2.40	1.56 ***	1.94	1.69 ***	1.81 **
NL	1.20	2.02	1.21	2.06	1.03	2.03	1.13	2.04
PL	0.70	2.08	0.88	2.39	0.63 *	2.06	0.72	1.93
PT	1.13	1.75	1.13	1.51	0.91 **	1.84	0.93 **	1.50 ***
RO	1.55	1.85	1.47	1.95	1.58	1.77	1.46	1.57
SE	0.58	0.90	0.59	0.89	0.48	0.91	0.51	0.78
SI	1.00	1.82	1.00	1.86	0.98 **	1.95	0.98	1.78 ***
SK	0.66	2.02	0.70	2.21	0.64	2.18	0.59	2.03
UK	0.70	1.70	0.79	2.19	0.63 **	1.85	0.78	1.79

*Note: Stars, *,**, and ***, indicate that Harvey, Leybourne, and Newbold (1997) equal accuracy hypothesis of the Bermingham-type forecasts vs. RW is rejected in favour of Bermingham model, at the significance level of 10%, 5%, and 1%, respectively. The evaluation period is the last 48 months window. The selected forecast horizons are 3 and 12 months.*

Source: Own calculations.

Tab. 7 Bias (ME) of random walk (RW) and the Bermingham (2010) type forecasts of headline annual inflation rates using 3 underlying inflation indicators (XUFE, pHICP, pXUFE)

	benchmark		Forecasts from Bermingham type model (horizon h)					
	RW		XUFE		pHICP		pXUFE	
	h=3	h=12	h=3	h=12	h=3	h=12	h=3	h=12
AT	-0.18	-0.86	-0.19	-1.25	-0.11	-0.87	-0.46	-1.14
BE	0.02	-0.35	0.08	-1.23	0.11	-0.36	-0.51	-0.84
BG	-0.24	-1.12	-0.23	-1.24	-0.18	-1.50	-0.23	-1.47
CY	-0.15	-1.71	-0.19	-2.25 *	0.12	-1.72	-0.82	-1.89
CZ	-0.08	-1.21	-0.29	-1.48 *	-0.03	-1.24	-0.25	-1.24
DE	-0.10	-0.94	-0.17	-1.19	-0.02	-0.97	-0.35	-1.12 *
DK	-0.11	-0.68	-0.14	-0.98	-0.01	-0.73	-0.28	-0.69
EE	-0.17	-1.88	-0.17	-2.50 **	-0.06	-2.00	-0.36	-1.76
EL	0.07	0.04	0.10	0.19	0.38	-0.01	-0.13	0.25
ES	-0.25	-1.28	-0.05	-1.55	0.01	-1.45	-0.72	-1.28
FI	-0.22	-1.59 **	-0.17	-1.74 **	-0.04	-1.58 *	-0.22	-1.60 **
FR	-0.05	-0.72	-0.11	-0.96	0.03	-0.83	-0.17	-0.73
HU	-0.20	-1.70	-0.28	-1.39	0.01	-1.73	-0.20	-1.48
IE	-0.08	-0.55	0.05	-0.65	-0.06	-0.55	-0.18	-0.55
IT	-0.24	-1.19	-0.03	-1.25	0.07	-1.22	-0.41	-1.10
LT	-0.19	-1.34	-0.32	-2.43 **	-0.08	-1.34	-0.62	-1.49
LV	0.08	-0.27	0.09	-1.18	0.13	-1.09	0.17	-0.23
MT	-0.12	-0.54	0.22	-1.30	-0.10	-0.54	-0.76	-0.65
NL	-0.25	-1.46	-0.15	-1.54	-0.19	-1.48	-0.67	-1.48
PL	-0.14	-1.47	-0.53	-2.05	-0.04	-1.42	-0.45	-1.54
PT	0.03	-0.39	0.08	-0.55	0.12	-0.47	-0.12	-0.39
RO	-0.22	-1.41	-0.35	-1.64	-0.24	-1.33	-0.26	-1.19
SE	0.06	0.21	-0.05	-0.15	0.12	0.21	-0.25	0.04
SI	-0.18	-1.38	-0.23	-1.56	-0.14	-1.46	-0.23	-1.34
SK	-0.22	-1.33	-0.33	-1.80	-0.17	-1.78	-0.25	-1.31
UK	-0.19	-1.41	-0.40	-2.01 **	-0.10	-1.54	-0.54	-1.53

*Note:. Stars, * and **, indicate that the RW or model forecasts of headline HICP are biased at the significance level of 10%, and 5%, respectively. The evaluation period is the last 48 months window. The selected forecast horizons are 3 and 12 months.*

Source: Own calculations.

Fig. 7 Core Inflation Score (CIS) of: HICP excluding unprocessed food and energy (XUFE), permanent HICP (pHICP) and permanent XUFE (pXUFE) components

Source: Own calculations.

Fig. 8 Decomposition of HICP (left column) and XUFE (right column) inflation rates into permanent and transitory components from Jan 2002 to Nov 2016 (% in year-over-year terms)

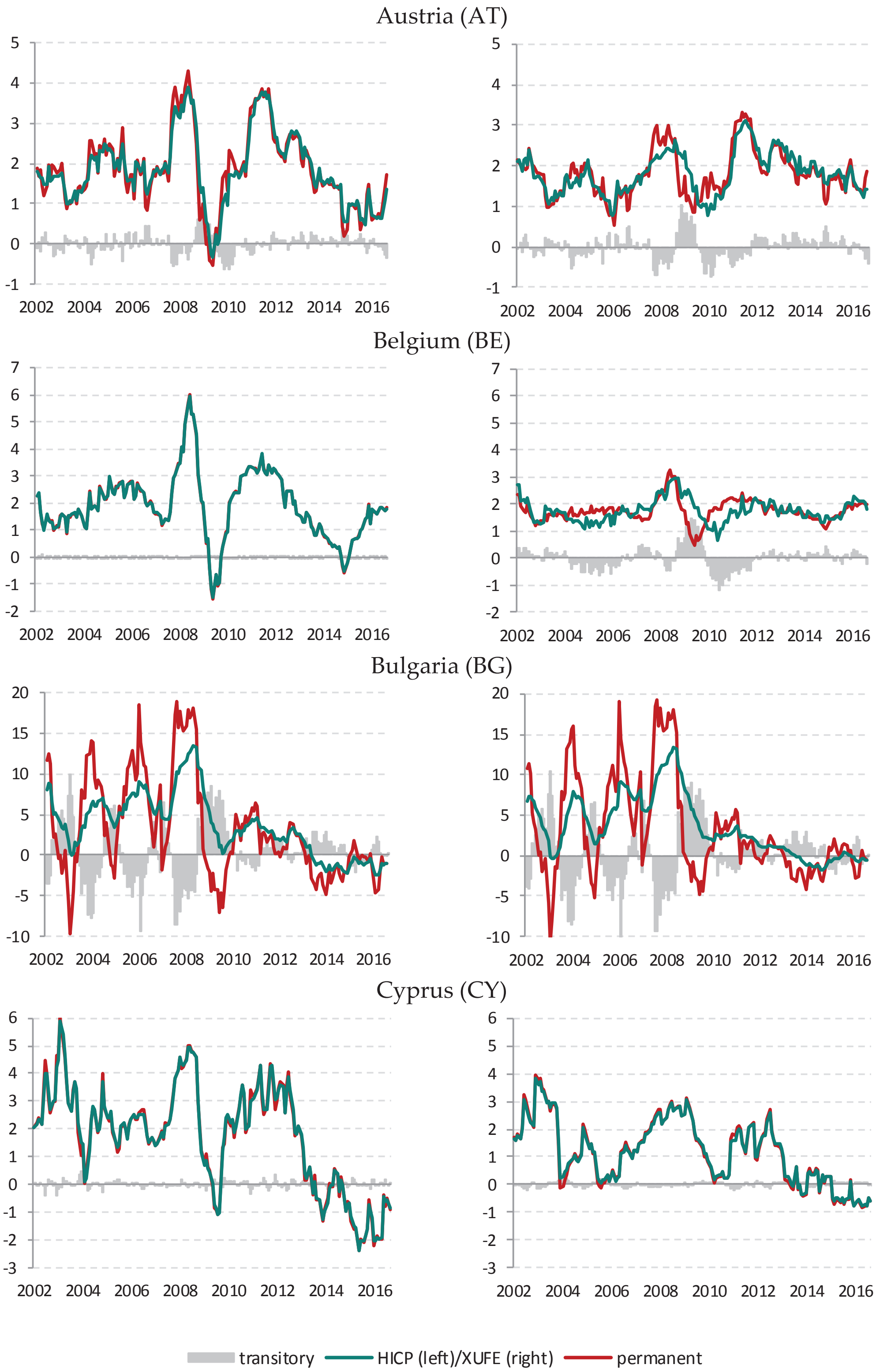
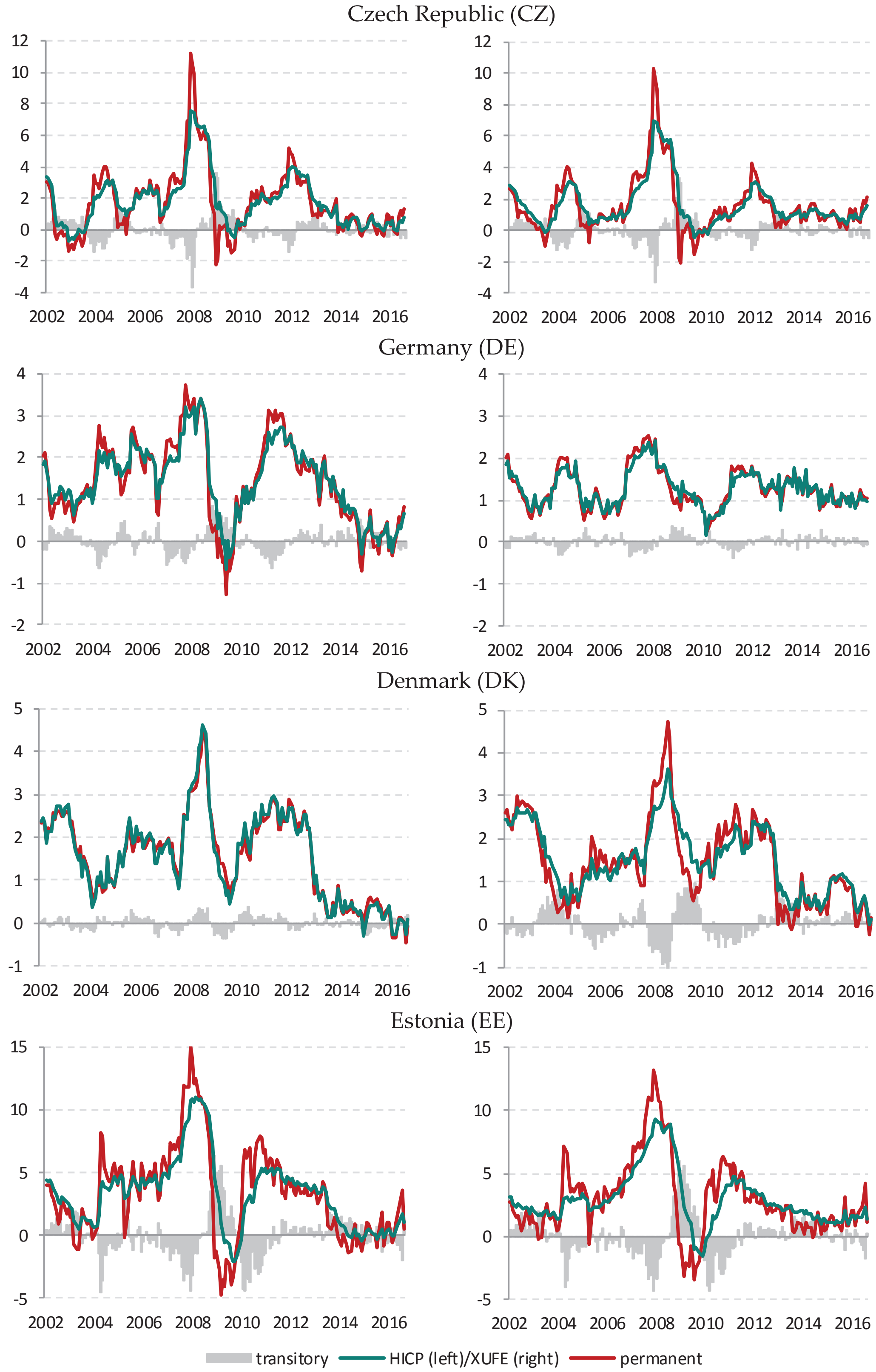


Fig. 8 cont.



Czech Republic (CZ)
Germany (DE)
Denmark (DK)
Estonia (EE)
12
10
8
6
4
2
0
-2
-4
2002
2004
2006
2008
2010
2012
2014
2016
transitory
HICP (left)/XUFE (right)
permanent

Fig. 8 cont.

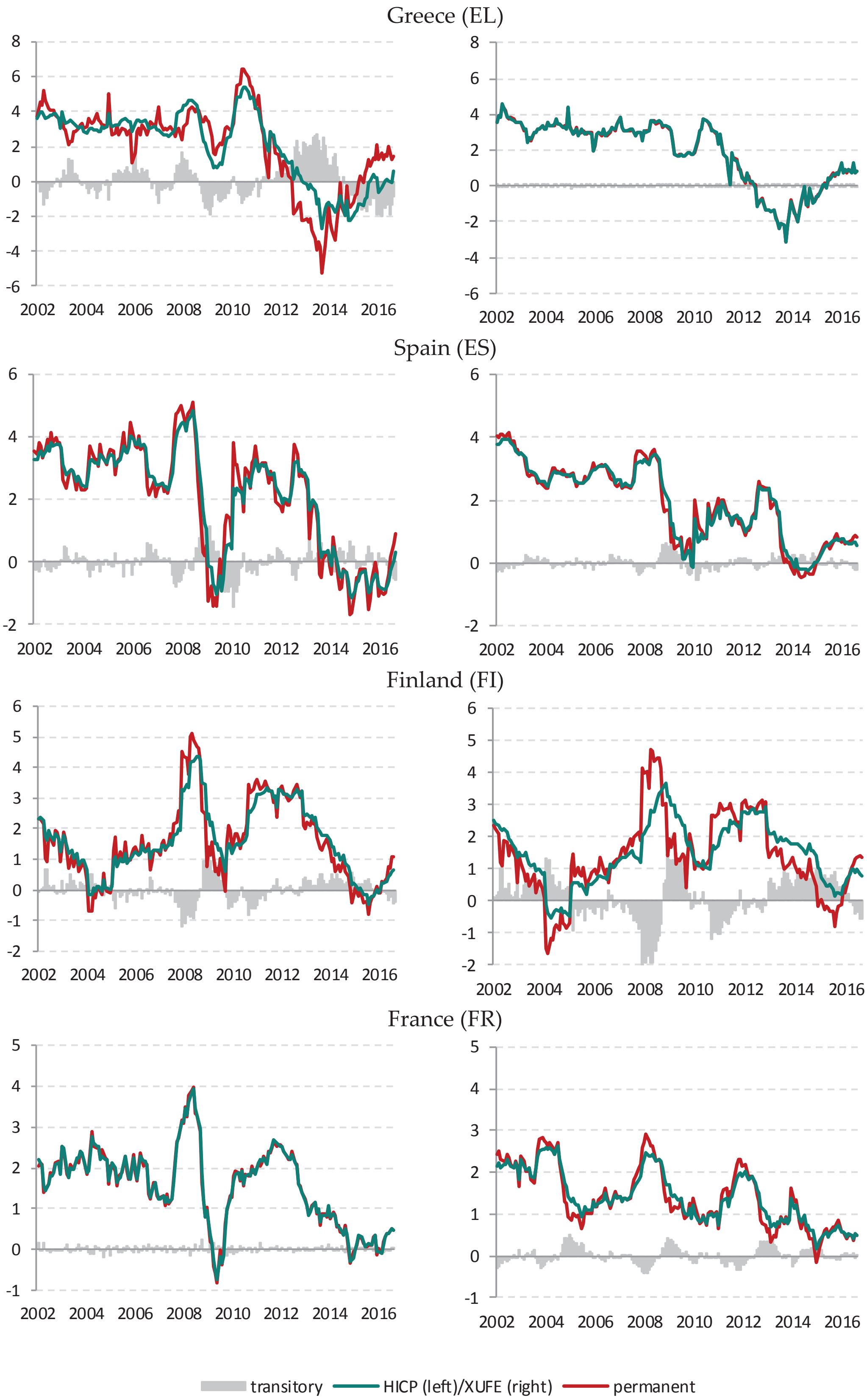

Greece (EL)
Spain (ES)
Finland (FI)
France (FR)
2002 2004 2006 2008 2010 2012 2014 2016
transitory
HICP (left)/XUFE (right)
permanent

Fig. 8 cont.

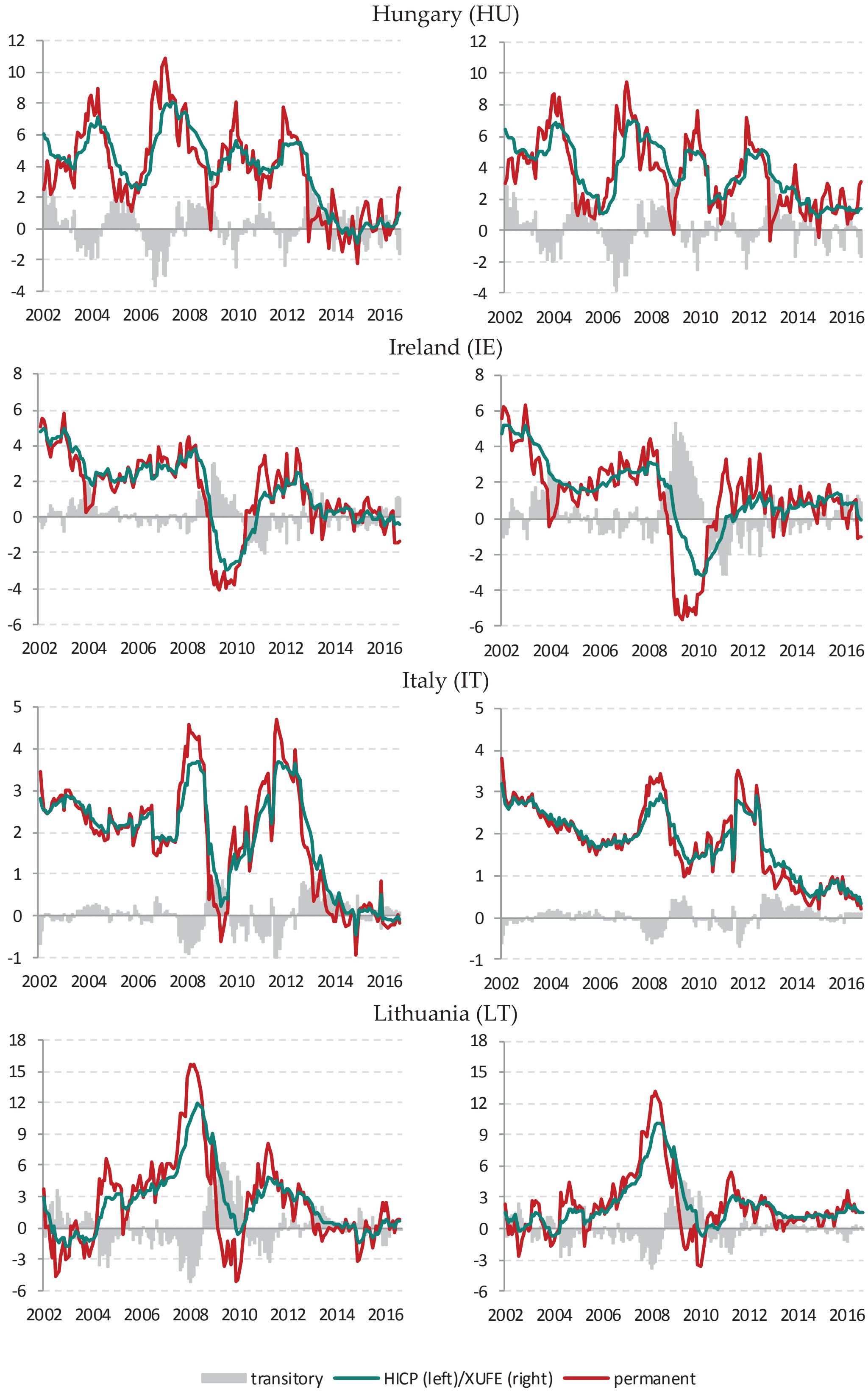

Hungary (HU)
Ireland (IE)
Italy (IT)
Lithuania (LT)
2002 2004 2006 2008 2010 2012 2014 2016
transitory
HICP (left)/XUFE (right)
permanent

Fig. 8 cont.

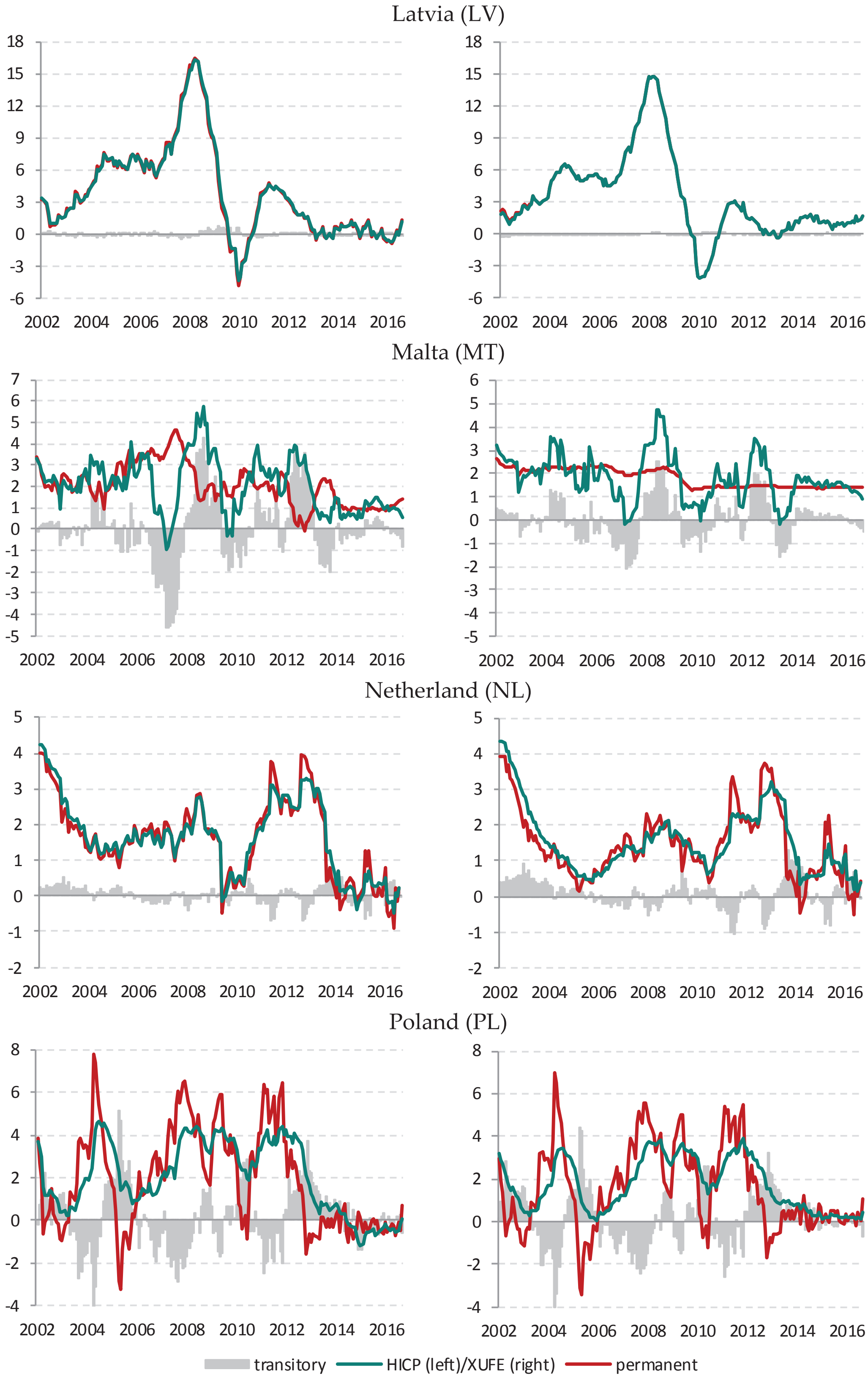
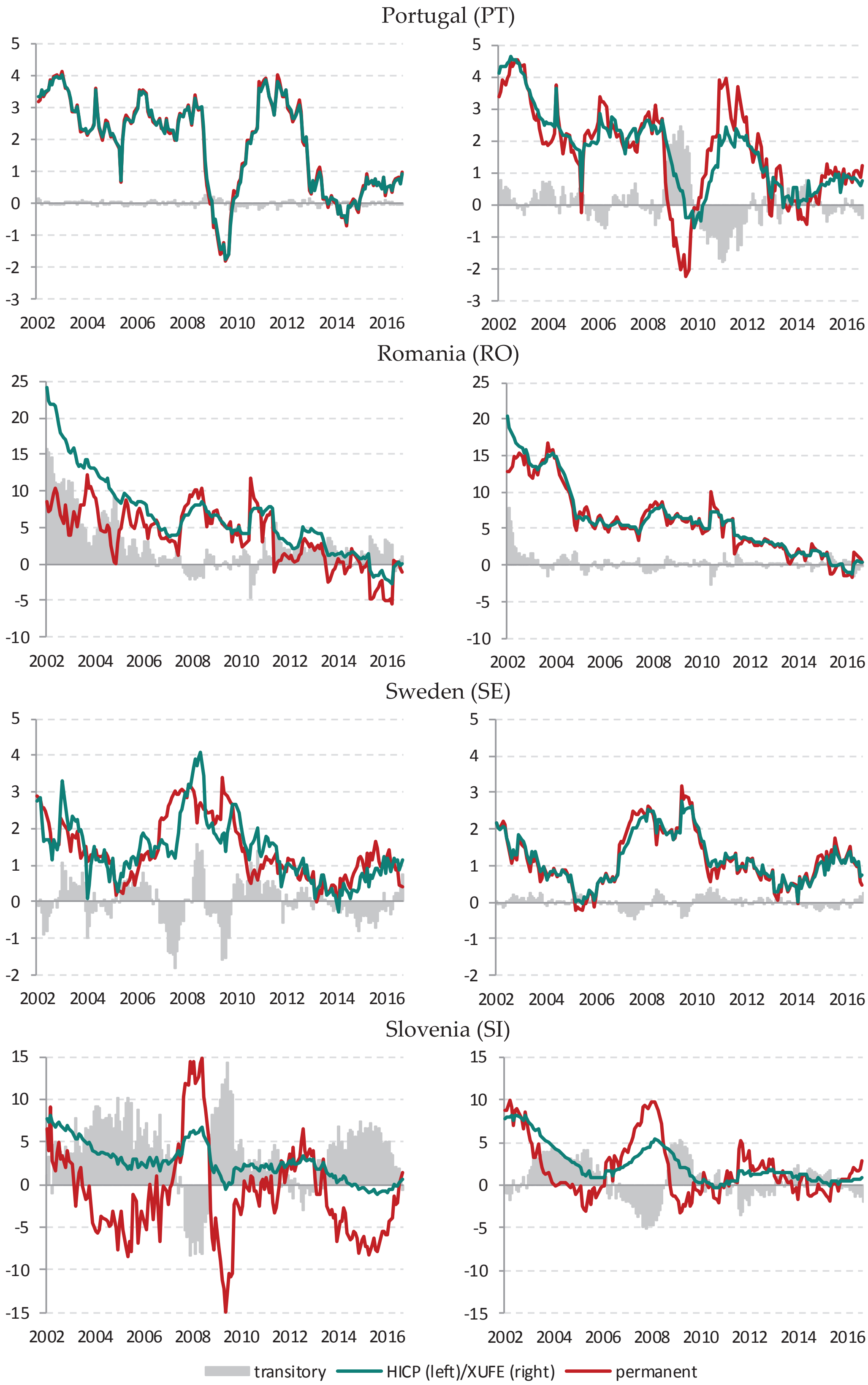
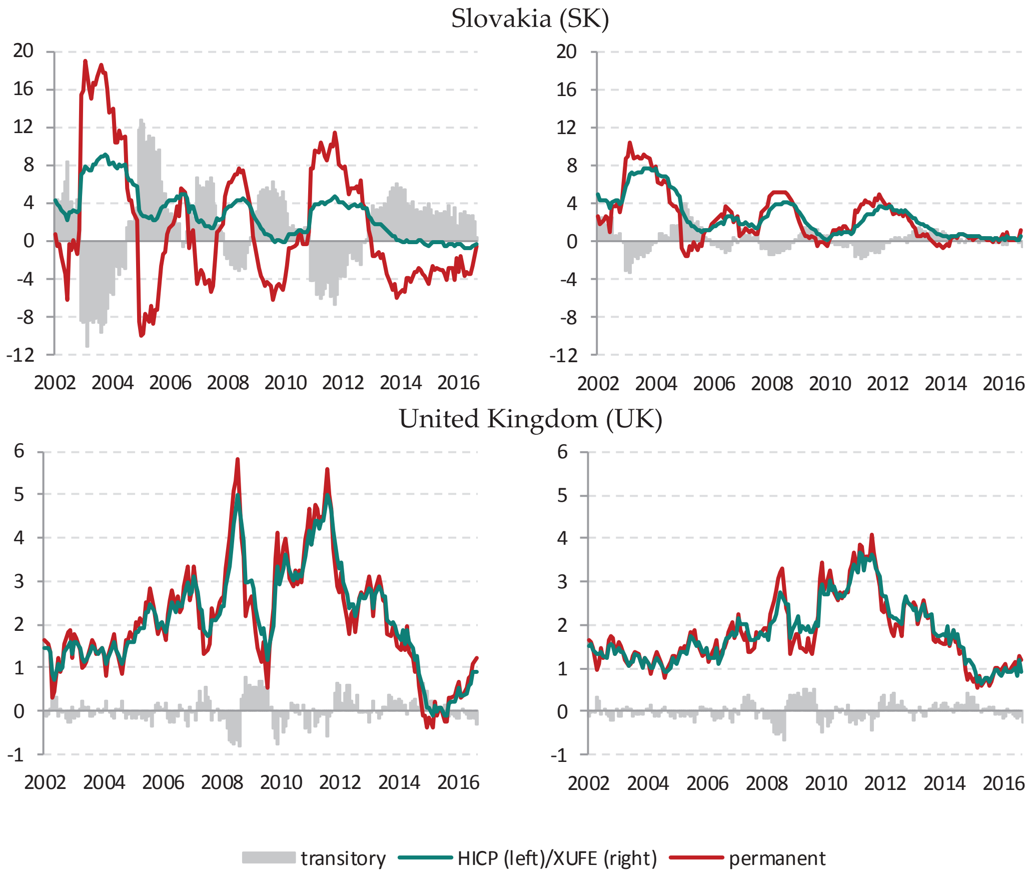

Latvia (LV)
18
15
12
9
6
3
0
-3
-6
2002 2004 2006 2008 2010 2012 2014 2016
Malta (MT)
7
6
5
4
3
2
1
0
-1
-2
-3
-4
-5
Netherland (NL)
5
4
3
2
1
0
-1
-2
Poland (PL)
8
6
4
2
0
-2
-4
transitory
HICP (left)/XUFE (right)
permanent

Fig. 8 cont.



Portugal (PT)
Romania (RO)
Sweden (SE)
Slovenia (SI)
2002 2004 2006 2008 2010 2012 2014 2016
transitory
HICP (left)/XUFE (right)
permanent

Fig. 8 cont.



Source: Eurostat, own calculations.

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