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Michal Franta, Ivan Sutóris*

Abstract

We decompose the Czech inflation time series into the trend and short-lived deviations from the trend by means of an unobserved component stochastic volatility model. We then carry out a regression analysis to interpret the two inflation components. The results indicate a fall in the inflation trend since the start of the sample (1998) which coincides with the introduction of the inflation targeting regime and with subsequent changes to the inflation target pursued by the Czech National Bank. Moreover, the regression analysis suggests that inflation expectations play a dominant role in the evolution of the trend. The behavior of the deviations from the trend exhibits features of an open-economy Phillips curve.

Abstrakt

Za použití modelu nepozorovaných komponent se stochastickou volatilitou rozkládáme časovou řadu české inflace na trend a krátkodobé odchylky od tohoto trendu. Na základě provedené regresní analýzy jsou pak tyto dvě komponenty inflace dále interpretovány. Výsledky ukazují pokles inflačního trendu od počátku vzorku dat (1998), který se kryje se zavedením režimu cílování inflace a s následnými změnami inflačního cíle České národní banky. Regresní analýza navíc naznačuje, že dominantní roli ve změnách trendu hrají inflační očekávání. Chování odchylek inflace od trendu pak vykazuje vlastnosti Phillipsovy křivky v otevřené ekonomice.

JEL Codes: E5, E31.

Keywords: Czech inflation, inflation trend, Phillips curve, UCSV.

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1. Introduction

Several issues regarding observed inflation have arisen recently. The main concern relates to the unprecedented fall in economic activity after 2008 which, however, was not accompanied by a corresponding fall in prices, and the lack of inflation when economic activity subsequently revived. The question is, then, whether the macroeconomic relationship between inflation and the degree of economic slack has weakened. A branch of literature suggests that, for some countries, the inflation process can be well described by a univariate framework decomposing inflation into a stochastic trend and a transitory component, with economic slack no longer dominating the price dynamics (Cecchetti et al., 2017).

Other papers have confirmed the existence of the standard Phillip's relationship but have noted that other phenomena play a role, for example, the influence of oil prices on inflation expectations (Coibion and Gorodnichenko, 2015) and problems with variables capturing economic activity (Baxa et al., 2017). If the standard Phillips curve cannot describe the basic macroeconomic relationship between the real economy and prices accurately any more, the modeling and forecasting frameworks should be adjusted accordingly. At the same time, the importance of changes in the trend component of the inflation process has also become increasingly clear, as reviewed, for example, by Ascari and Sbordone (2014).

The questions about the inflation process which have appeared recently, however, pertain mainly to the most advanced economies. For economies such as the Czech one, the inflation process seems to behave more along the lines of standard economic relationships (Baxa et al., 2015). This paper attempts to examine the recent inflation movements in the Czech Republic through the lens of the new approaches that have arisen in the literature recently. To that end, we closely follow analyses that have been conducted for the US (Cecchetti et al., 2017) and the UK (Forbes et al., 2019). In the first step, a variant of the seminal Stock and Watson (2007) unobserved component stochastic volatility model is employed to decompose the inflation series into a slow-moving component (the trend) and short-lived deviations from this trend (the cyclical component). Next, based on simple linear regressions, the components are interpreted with respect to the usual potential drivers of the inflation process.

The framework employed to decompose the inflation series is simple yet flexible enough. The simple univariate framework is an advantage in the presence of regime changes and thus in the presence of short time series related to a single policy regime as seen in the Czech Republic. Next, the framework still allows for economic interpretation of the filtered components. Using even simpler filters (such as the HP filter), proper economic interpretation of the decomposed series would be more difficult to do even though the in-sample fit might be superior.[1] Finally, note that this paper seems to be the first study that tries to answer the recent questions related to the evolution of Czech prices. Baxa et al. (2013) appears to be the only study drawing on the approach of Stock and Watson (2007) in the context of Czech macroeconomic time series. It, however, focuses on economic activity and its ability to explain the dynamics of inflation. Our research question is much broader and involves all possible inflation drivers.

[1] The HP filter can filter out a deterministic trend but not necessarily a stochastic trend (Phillips and Jin, 2015). Our approach ensures that the statistical properties of the inflation components, i.e., non-stationarity of the trend and stationarity of the cyclical component, are taken into account.

From the policy perspective, the point of the paper is to provide a complementary analysis to general equilibrium models, where the inflation trend is pinned down by the inflation target assumed in the monetary policy rule and rational expectations are imposed. The statistical approach employed in this paper does not impose the nominal trend and rational expectations.

The estimation results indicate a fall in the inflation trend since the start of the sample (1998) which coincides with the introduction of the inflation targeting regime and with subsequent decreases of the inflation target set by the Czech National Bank. In addition, it turns out that the univariate framework replicates the basic inflation dynamics observed in the data more closely than the standard structural multivariate framework. This observation justifies the use of the unobserved component stochastic volatility model.

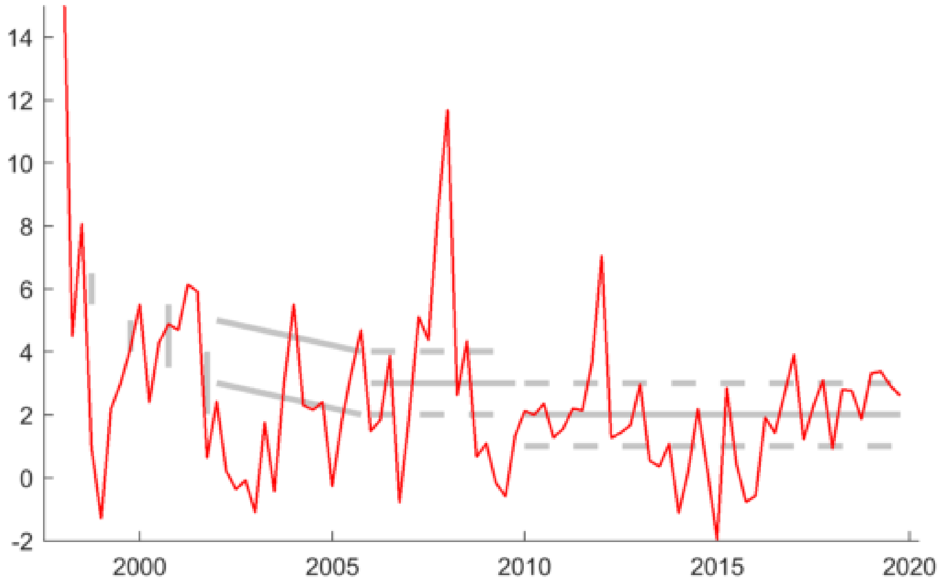
The regression results from the second step demonstrate that the behavior of the cyclical component broadly follows the logic of the open-economy Phillips curve. Conditional on the value of the trend, inflation is positively related to inflation expectations, foreign prices, and the output gap. In addition, the exchange rate and the credit spread display statistically significant coefficients. On the other hand, changes in the trend component seem to be related most strongly to changes in inflation expectations, with other variables playing a limited role.

The rest of the paper is organized as follows. Section 2 discusses the observed Czech inflation over the last two decades and provides the motivation for the model employed to decompose inflation. Section 3 presents the unobserved component stochastic volatility model and the decomposition of inflation into its permanent and cyclical components. Section 4 presents a regression analysis of the two components. Finally, Section 5 concludes. Additional results can be found in appendices.

2. Stylized Facts and Motivation

Figure 1 presents inflation observed in the Czech Republic over the period 1998Q1–2019Q4. The starting point coincides with the introduction of the inflation targeting regime. The beginning of the sample exhibits rather high levels of inflation. The Czech economy then went through a disinflation period and has been experiencing low and stable inflation since then. Several features of the inflation series can be observed. First, the decreasing trend observed immediately after the introduction of inflation targeting in 1998 and the medium-term inflation changes since then relate to changes to the inflation target as indicated by the grey lines in Figure 1. Changes to the inflation target presumably affect inflation expectations, and the modeling framework should be able to account for such changes in order to model the inflation process properly. Second, it seems that periods of lower and higher volatility can be distinguished. For example, the period 2006–2012 is the subject of higher inflation volatility than the period since 2013. Therefore, stochastic volatility should be allowed for in the modeling framework so that the inflation dynamics can be assessed appropriately.

Figure 1: Quarter-to-Quarter Percentage Change in CPI, Seasonally Adjusted (X12), Annualized Rate



Note: Thick lines indicate the inflation target (the target range). Note that net inflation (i.e., inflation excluding changes in indirect taxes and regulated prices) was targeted until 2001. Since 2002, headline inflation has been targeted.

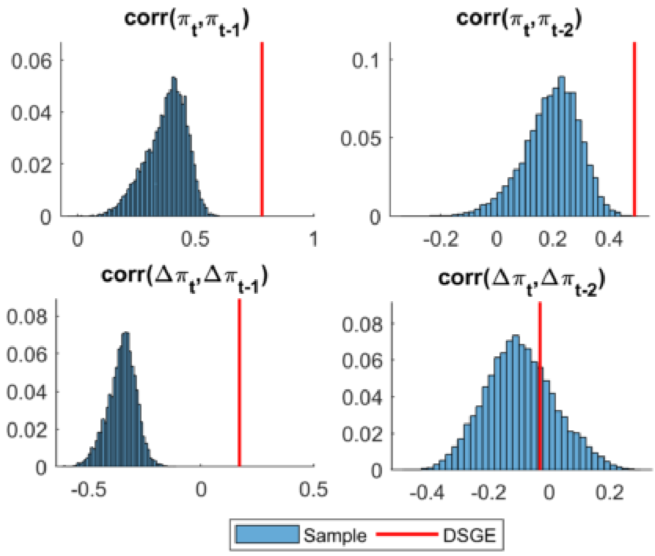
Testing the stationarity of the inflation series provides mixed results. While the Augmented Dickey Fuller test rejects the hypothesis of a unit root at the 1% level of significance, the KPSS test rejects the null of stationarity at the 10% level of significance. However, the changes to the inflation target and the resulting likely changes in inflation expectations suggest long-lasting changes in the trend component of inflation, which should be taken into account in the modeling framework.

Apart from the motivation of the modeling approach based on observed features of the inflation series, the motivation based on the deficiencies of other modeling approaches can be also discussed. Figure 2 compares the sample first-order autocorrelation of CPI inflation and its simulated counterpart from a standard New Keynesian DSGE model.[2] It turns out that the DSGE model implies significantly higher inflation persistence.[3] Moreover, the DSGE model considers inflation changes to be only slightly (positively) autocorrelated. Our data sample, however, suggests that an inflation change in one period is followed by the opposite movement in the next period. The dynamics of inflation imposed by the DSGE model differ significantly from those observed in the data. Cecchetti et al. (2007) point out that replicating inflation dynamics presents a challenge for standard New Keynesian models in general.

[2] The DSGE model used in this exercise is "g3" – the model used as the main forecasting and policy analysis tool at the Czech National Bank. For a description of the model, see Andrle et al. (2009). The model is calibrated, so the simulated autocorrelations are represented by numbers.

[3] The lower inflation persistence observed in the data sample could be a consequence of monetary policy conduct superior to the interest rate rule assumed in the DSGE model. Another possible reason for the difference in inflation persistence is related to the model structure and calibration. For our purposes, however, looking for the reason is not a purpose of the paper. We simply aim to find a model that replicates inflation moments in the data more accurately.

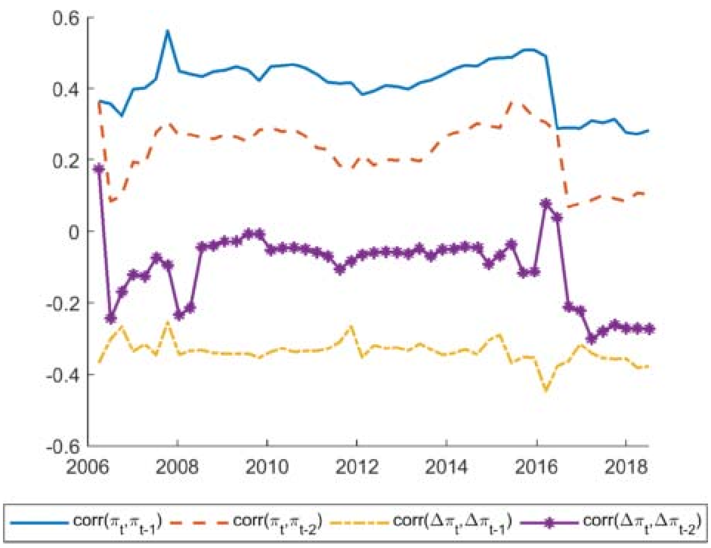
Figure 2: Sample Autocorrelation (Blue Bars) and Autocorrelation from the DSGE Model (Red Line) for CPI Inflation



Note: The empirical distributions of the sample autocorrelation are estimated employing a block bootstrap of window size equal to 9.

The sample autocorrelation of inflation is influenced by the fact that the inflation target (and the targeted inflation index) has changed several times since 1998 – as reported in Figure 1. The DSGE model imposes a nominal trend/inflation target equal to 2 percent. So, to avoid confusing inflation persistence with a long-lasting change in the inflation target, the first-order autocorrelation of inflation and differenced inflation on a rolling window of size 40 quarters is reported in Figure 3. It follows that the autocorrelations do not change much and the low persistence of inflation and the negative serial correlation of inflation changes are probably not a consequence of changes in inflation targets, but an inherent feature of the inflation process.

Figure 3: Autocorrelations of CPI Inflation and Differenced CPI Inflation over a Rolling Window of 40 Observations



Note: The x-axis indicates the last quarter included in the window used for the estimation.

Taking into account the discussion above, the model for inflation should be able to deal with long-lasting changes in the inflation trend and should exhibit moderate inflation persistence. Furthermore, innovations to inflation around the trend should be noisy and not too persistent, as suggested by the autocorrelations dealing with the first differences of inflation.

3. Model

We employ a variant of the unobserved component stochastic volatility model of Stock and Watson (2007). Observed inflation π_t is decomposed into its trend, τ_t, and its cyclical component, $\pi_t^c \equiv \pi_t - \tau_t$. The trend represents the slow-moving component of the inflation process. The specification allows for some persistence in the cyclical component of inflation. From the economic point of view, such persistence is justified by the price-setting process itself (the "intrinsic" inflation persistence). The persistence of the cyclical component suggests how easily inflation can be expected to reach its target. The rate of convergence captures the relationship between the monetary policy rule and private sector behavior (Cogley et al., 2010).

The autoregressive unobserved component stochastic volatility model (AR-UCSV) is formulated as in Chan et al. (2013):[4]

$$(\pi_t - \tau_t) = \rho_t(\pi_{t-1} - \tau_{t-1}) + \varepsilon_t exp\left(\frac{h_t}{2}\right)$$
$$\tau_t = \tau_{t-1} + \varepsilon_t^\tau$$
$$h_t = h_{t-1} + \varepsilon_t^h$$
$$\rho_t = \rho_{t-1} + \varepsilon_t^\rho$$

where $\varepsilon_t \sim N(0,1)$, $\varepsilon_t^\tau \sim N(0, \sigma_\tau^2)$, and $\varepsilon_t^h \sim N(0, \sigma_h^2)$. Using the truncated normal distribution in the form

$$\varepsilon_t^\rho \sim TN\big(0 - \rho_{t-1}, 1 - \rho_{t-1}; 0, \sigma_\rho^2\big),$$

the persistence of the cyclical component is bounded in the interval $0 < \rho_t < 1$.

The cyclical component follows an AR(1) process and the trend is modeled as a random walk, i.e., it exhibits a stochastic trend. Based on a statistical test carried out on a simpler version of the model described in Appendix A, we assume stochastic volatility in the cyclical component only, not in the trend component.[5]

Note that the AR-UCSV specification implies an interpretation of expected inflation as the weighted average of the current trend and observed inflation, with the weight on current inflation equal to the AR parameter:

$$E_t[\pi_{t+1}] = \rho_t\pi_t \mp (1 - \rho_t)\tau_t.$$

The interpretation is that a fraction of agents ρ_t form their expectations adaptively and a fraction $(1 - \rho_t)$ base their expectations on the inflation trend. At this point, we deviate again from the

[4] We use MATLAB code from Chan et al. (2013) for this exercise.

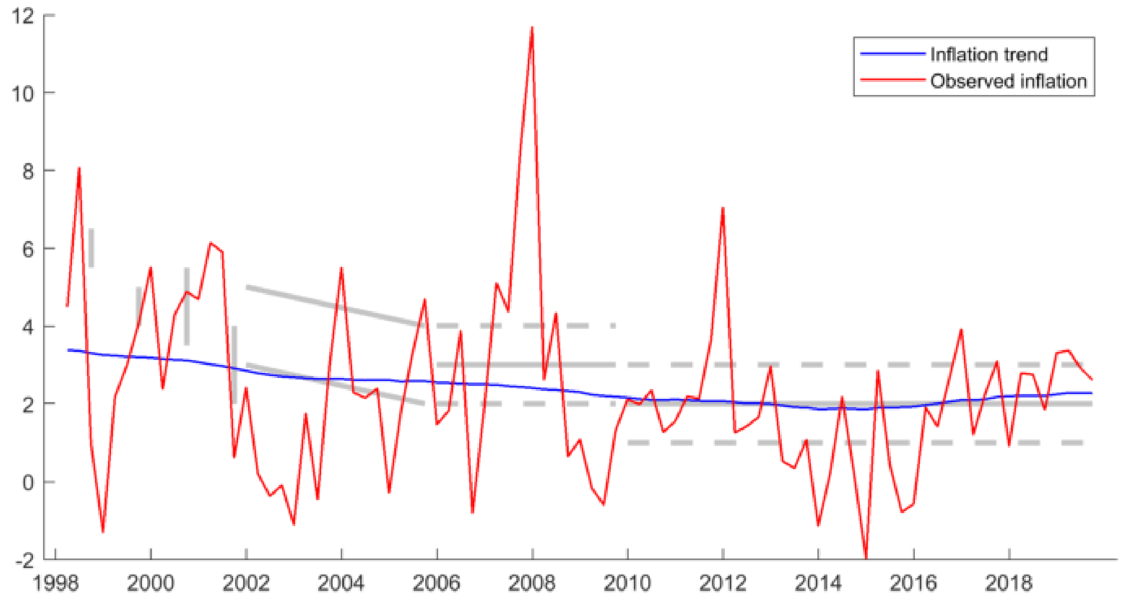
[5] Note that the two papers we follow assume no stochastic volatility in the case of the US (Cecchetti et al., 2017) and stochastic volatility in both components for the UK (Forbes et al., 2019).

traditional structural modeling approach, which assumes rational agents with expectations anchored around a constant inflation trend.

The estimation procedure is a Markov chain Monte Carlo (MCMC) procedure. The important point is the presence of an inequality constraint for ρ_t that leads to posterior sampling in a nonlinear state-space model. To sample from the conditional density $p(\rho|y,\tau,h)$, Gaussian approximation is used as the proposal density for the accept-reject Metropolis-Hastings (ARMH) step. The rest of the parameters are drawn from the respective conditional distributions using the MH algorithm or direct draws from a known distribution. For details, see Appendix A in Chan (2013). The dataset employed in the estimation procedure is of quarterly frequency. The reason we do not base our exercise on higher frequency (e.g., monthly) data is that the examined phenomena are of business cycle frequency (the cyclical component) or even lower frequency (the trend component). Moreover, some variables in the subsequent analyses are available at quarterly frequency only.

To initialize the trend, we assume $\tau_1 \sim N(6,1)$, reflecting the fact of heightened inflation expectations at the beginning of the sample, which enters the estimation in the form of an informative prior. Next, following Chan et al. (2013) the other states are initialized as follows: $h_1 \sim N(0,5)$ and $\rho_1 \sim TN(0,1;0,1)$. So, with the exception of the trend, the other states are initialized using noninformative distributions. The priors on σ_τ^2, σ_h^2, and σ_ρ^2 follow an inverse-Gamma distribution with degrees of freedom equal to 10 and scale parameters equal to 0.18, 0.009, and 0.45, respectively.[6] The inference is based on 35,000 draws with a burn-in period of 5,000 draws.

Figure 4: Observed Inflation and the Estimated Trend in the AR-UCSV Model



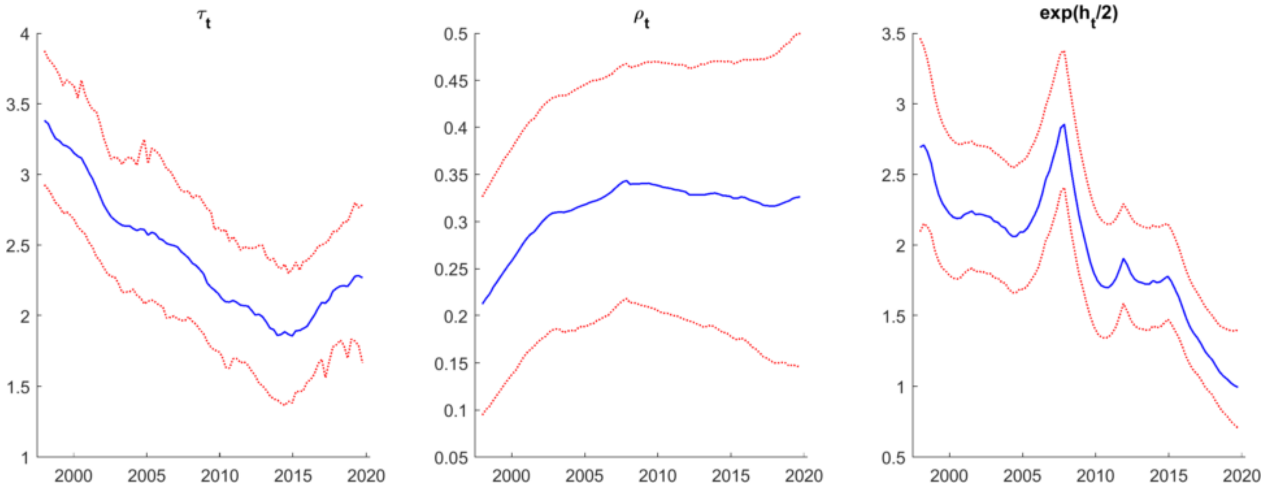
Note: Thick lines indicate the inflation target (the target range). Note that net inflation was targeted until 2001. Since 2002, headline inflation has been targeted.

[6] As noted in Chan et al. (2013), the inverse-Gamma priors are set to be relatively noninformative. For example, in the case of σ_τ^2, $IG(10{,}0.18)$ implies that with high probability the difference $\tau_t - \tau_{t-1}$ lies between -0.3 and 0.3.

Figure 4 presents the estimated trend component together with the observed inflation series. It can be seen that the trend component was moving mostly around the lower bounds of the inflation target in the first half of the sample. This observation can be explained by policymakers' preference for below-target inflation when inflation targeting was introduced (CNB, 2008) or by public understanding of inflation targeting as a regime of stable prices, not steadily growing ones. In addition, some role could be played by the strong influence of exchange rate appreciation on prices, because target undershooting can be seen in periods of strong appreciation of the Czech koruna.

Furthermore, the trend component started to move below the 2% target in 2012. This fact supports the launch of the exchange rate commitment, which was used as a tool for further easing monetary policy when the interest rate was stuck at the zero lower bound and deflationary pressures started to spread throughout the economy (Franta et al., 2014). The estimated trend suggests the existence of a danger that inflation expectations were no longer anchored to the inflation target. The exchange rate commitment was launched in November 2013. Similarly, a switch of the trend component close to the 2% inflation target can be observed in 2017, which corresponds to the date of termination of the exchange rate commitment on April 6, 2017.

Figure 5: The Estimated Trend, the AR Parameter, and the Standard Deviation of the Cyclical Component



Note: Red dotted lines indicate the 68% confidence band.

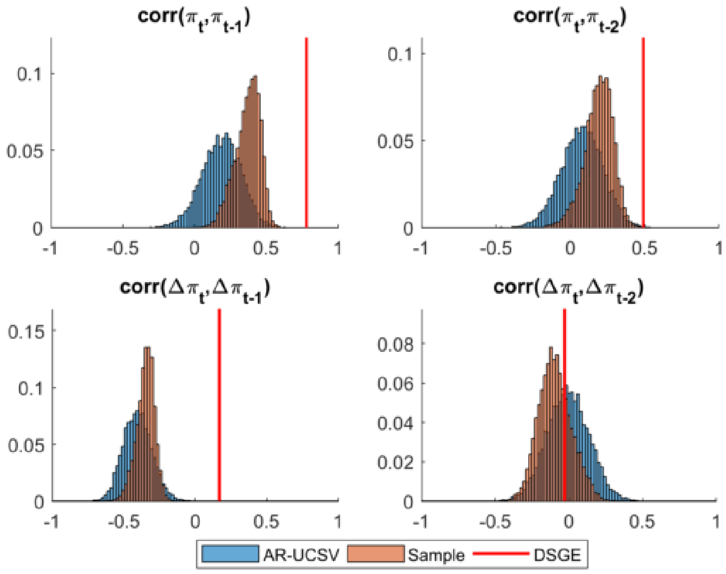
Figure 5 reports the estimated trend τ_t, the persistence parameter ρ_t, and the standard deviation of the cyclical component together with the corresponding 68% confidence bands. It turns out that the inflation trend decreases over time. The fall can be explained by the changes in the inflation target (see Figure 4). Nevertheless, while the fall in the trend over the period examined exhibits some degree of statistical significance, the evidence for anti-inflationary bias (i.e., whether the trend resides often close to the lower bound of the target or the tolerance band for the target) and the effect of unconventional measures (the rebound of the trend in 2014) are not statistically significant.

The persistence of the inflation process represented by the AR parameter of the cyclical component ρ_t suggests that around one third of expected inflation is given by the observed current inflation and two thirds are related to the slow-moving part of inflation (the trend). The weight of observed inflation increases over time, but the increase is not statistically significant.

The standard deviation of the cyclical component decreased between 1998 and 2000, which could be viewed as success of the introduction of inflation targeting as a disinflation strategy. The second peak in the variance can be observed around the Global Financial and Economic Crises around 2009, when the Czech economy was subjected to extreme foreign demand shocks, and later also by oil price shocks, and the external environment switched to a low-inflation one.

The posterior mean of the standard deviation of the trend equals $\sigma_\tau = 0.02$. When compared with the estimated standard error of the cyclical component (Figure 5), it follows that the majority of the variation in the inflation process comes from the noise in the cyclical component. Finally, the posterior mean of the standard deviation in the process that drives the change in the AR parameter equals $\sigma_\rho = 0.001$, suggesting very slow evolution in the persistence of the cyclical component of inflation.

Figure 6: Simulated Autocorrelations from the AR-UCSV (Blue Bars) and DSGE (Red Line) Models and Sample Autocorrelations (Brown Bars)



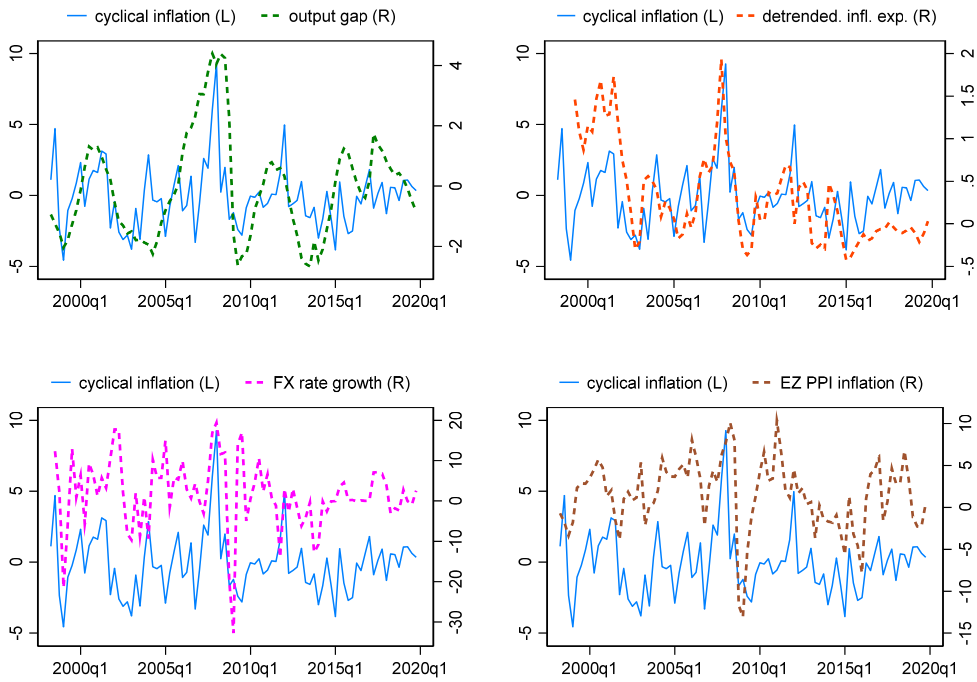
Note: The empirical distributions of the sample autocorrelation are estimated employing a block bootstrap of window size equal to 9.

Looking back at our motivation for the UCSV type of model for inflation discussed in Section 2, Figure 6 shows the simulated autocorrelation of inflation and differenced inflation from the estimated AR-UCSV model and the moments found in the data and simulated in the DSGE model. The AR-UCSV model seems to capture the basic correlations of the inflation process more accurately than the DSGE model. This observation provides some justification for modeling inflation by simple UCSV-type models.

4. Regression Analysis of Inflation Components

After decomposing the process for inflation into a trend and a cycle component, we are interested in how each component is related to other macroeconomic variables that can potentially affect prices in the economy. Figure 7 displays the cyclical component of inflation together with the output gap, inflation expectations, currency appreciation, and Eurozone inflation. The overall picture indicates a solid degree of comovement between the series and thus the possibility that the behavior of inflation is driven by the standard fundamentals. In the rest of this section, we will investigate these relationships more formally by estimating various regressions with a single additional macroeconomic variable at a time. Our results should therefore be interpreted as statistical associations rather than causal effects. Nevertheless, we believe such reduced-form evidence can be informative about macroeconomic patterns that take into account the existence of the inflation trend.

Figure 7: Cyclical Component of Inflation



The variables we consider include measures of domestic and foreign economic activity such as the output gap and cyclical unemployment, expectations about future inflation, changes in the exchange rate, growth in prices of oil, wages, and foreign producer prices, and financial conditions such as quantities of credit and spreads on interest rates. The full list of the variables we use is provided in Table 1. The data are quarterly and run from 1998 (depending on data availability) to 2019Q4. Our output gap variable is obtained as the HP-filtered cycle of real GDP, while the unemployment gap is the difference between the actual rate and the OECD measure of the NAIRU. We use three series of inflation expectations, always regarding inflation over the next four quarters: those of financial analysts, those of firm managers (both sourced from regular surveys conducted by the CNB), and those of the central bank (published regularly in the CNB's Inflation Reports). We also include growth in the EUR/CZK exchange rate, which is relevant in the open-economy setting. Credit conditions are captured by the spread between average rates on bank loans to firms and the interbank rate, as well as by the growth rates of credit to the private nonfinancial sector and of the broad money aggregate. The growth rates of Czech nominal wages,[7] the Brent crude oil price, and the Eurozone PPI capture additional price pressures and global inflation trends (cf. Ciccarelli and Mojon, 2010). Finally, we also construct a measure of the global output gap as a weighted average of OECD countries' output gaps (again estimated by the HP filter) weighted by the volume of Czech exports to each country (similar to Borio and Filardo, 2007). Table B1 in Appendix B displays the number of observations and summary statistics for inflation, macroeconomic variables, and for the trend component estimated in the previous section.

Table 1: List of Variables

group	variable	description	source
inflation	INFL	CPI inflation (%, qoq, ann., s.a.)	CZSO, own calc.
economic activity	OUTGAP	output gap, HP filt. (%)	CZSO, own calc.
	UGAP	unempl. gap wrt. OECD NAIRU (%)	CZSO, OECD
	GLOBGAP	global output gap (%)	OECD, IMF, own calc.
exchange rate	GCZKEUR	growth in CZK/EUR rate (%, qoq, ann., pos.=apprec.)	CNB
expectations	IEFA	infl. exp., fin. analysts (%, yoy)	CNB
	IEFM	infl. exp., firm managers (%, yoy)	CNB
	IECNB	infl. exp., CNB (%, yoy)	CNB
credit	GM2	growth of M2 (%, qoq, ann.)	CNB
	SPREAD	spread btw. rate on loans to firms and PRIBOR (%)	CNB
	GCRED	growth of credit to priv. nonfin. sector (%, qoq, ann.)	BIS
prices	GOIL	growth of Brent oil price (%, qoq, ann.)	World Bank
	GWAGE	growth of average nominal wage (%, qoq, ann., s.a.)	CZSO, own calc.
	GEZPPI	growth of Eurozone PPI (%, qoq, ann.)	Eurostat

Note: CZSO denotes the Czech Statistical Office.

Our analysis is inspired by Forbes et al. (2019). The first set of regressions relates observed inflation to its trend component and an additional macroeconomic variable:

$$INFL_t = \alpha + \beta \cdot TREND_t + \gamma \cdot XVAR_t + \varepsilon_t. \qquad (1)$$

Finding that *XVAR* helps to explain inflation even conditional on the trend would suggest it is related to the cyclical component of inflation. The regression coefficients and their standard errors are shown in Table 2. In addition, the last column displays the partial R-squared, defined as the share of the residual variation in inflation not explained by a model with only a trend that is explained by adding *XVAR*.

We find that, in agreement with the logic of the open-economy Phillips curve, the local and global output gap, inflation expectations, the exchange rate, and foreign prices are statistically significant at the 1% level of significance and exhibit a nontrivial partial R^2. The estimates seem to have the correct signs and are also quantitatively consequential. To get some idea about the effect of a usual change of the explanatory variable on inflation implied by the estimated regression, we normalize the coefficients by the volatility of each variable and divide by four and find that a one standard deviation movement in each of these variables is associated with a movement in (non-

[7] Since the average wage displays clear seasonality, we calculate and use a seasonally adjusted series using the X-13ARIMA-SEATS method.

annualized) qoq inflation ranging from 0.15 pp (the exchange rate) to about 0.50 pp (expectations of analysts and managers), with output gaps, foreign prices, and CNB expectations being in the middle with values of 0.20–0.25 pp.[8] The remaining variables are not significant at the 5% level.

Table 2: Regression of Inflation on the Trend and an Additional Variable

	const	trend	xvar	partial R^2
OUTGAP	-1.72	1.69***	0.58***	0.21
	(1.08)	(0.40)	(0.12)	
UGAP	-2.07	1.85***	-0.48*	0.05
	(1.44)	(0.62)	(0.26)	
GLOBGAP	-1.05	1.39***	0.86***	0.23
	(0.96)	(0.38)	(0.11)	
GCZKEUR	-1.15	1.38***	0.07***	0.08
	(0.92)	(0.38)	(0.03)	
IEFA	2.26	-2.53	2.24***	0.20
	(2.19)	(1.70)	(0.78)	
IEFM	2.04	-2.25***	2.06***	0.37
	(1.46)	(0.86)	(0.34)	
IECNB	-2.53	1.01	1.05*	0.10
	(3.52)	(2.03)	(0.58)	
GM2	-0.88	0.98	0.10*	0.05
	(3.11)	(1.43)	(0.05)	
SPREAD	1.65	1.69	-2.13*	0.06
	(4.66)	(1.48)	(1.26)	
GCRED	-1.89	1.71***	0.02	0.00
	(1.19)	(0.47)	(0.04)	
GOIL	-1.41	1.52***	0.01	0.03
	(0.95)	(0.38)	(0.00)	
GWAGE	-1.27	1.15	0.15	0.05
	(1.95)	(1.02)	(0.15)	
GEZPPI	-1.40**	1.41***	0.21***	0.15
	(0.65)	(0.30)	(0.04)	

Note: HAC-robust standard errors in parentheses (* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$). Each row displays the results of a regression as in equation (1) with a different additional variable. The partial R^2 column shows what proportion of the variance unaccounted for by the trend only is explained by adding XVAR.

As a robustness check, Table B2 in Appendix B presents the partial R^2 also for a specification with an additional three lags of the *XVAR* variable. Although the fit is increased more for some variables than for others, the set of variables with the most explanatory power stays largely the same. Another possible issue is bias stemming from measurement error in the trend series, since we do not observe its true value and work with its posterior mean. Table B3 in Appendix B presents the distribution of the regression point estimates (of the *XVAR* coefficient) obtained from 30,000 MCMC draws of the trend series. The means of the point estimates are largely consistent with the results in Table 2, while their dispersion is small relative to the regression standard errors, indicating that the effects of measurement error in our setting are not serious.

[8] The positive coefficient on exchange rate growth could be viewed as counterintuitive; the positive sign implies the interpretation that an acceleration in appreciation coincides with the increase in inflation. The sign of the coefficient is probably a consequence of a decreasing pace of economic convergence, reflected in a fall in equilibrium appreciation together with a fall in trend inflation, which is not necessarily fully accounted for by the estimated trend.

The coefficient of the trend is close to (and not statistically significantly different from) unity for most variables, which is consistent with the functional form of the AR-UCSV model. The exception is the inflation expectations of analysts and managers, conditional on which inflation seems to be related negatively to the trend. One possible explanation for this pattern is the simultaneity associated with the forward-looking nature of expectations. If expectations themselves are positively related to current inflation and the trend, conditioning on expectations could induce a negative correlation between the two, similar to collider bias. Nevertheless, the relatively high share of the variance explained indicates that expectations are still related to inflation in a way that is not simply subsumed by the trend.

Interestingly, our results for the cyclical components are in contrast to those found for the UK in Forbes et al. (2019) and for the US in Cecchetti et al. (2017). Using a similar reduced-form approach, these papers report a weak relationship between inflation and traditional explanatory variables such as measures of slack or expectations, once movements in the trend are taken into account.

The above results indicate that the variables suggested by the open-economy Phillips curve model are related to inflation even when controlling for long-run drift in inflation. On the other hand, we may be interested in whether changes in the inflation trend itself can also be explained by other macroeconomic developments. We thus estimate another set of regressions relating the change in trend (expressed in basis points, as the trend changes only a little from quarter to quarter) to the current and lagged values of a particular variable:

$$100 \cdot \Delta TREND_t = \alpha + \sum_{i=0}^{3} \gamma_i \cdot XVAR_{t-i} + \varepsilon_t. \qquad (2)$$

In the above equation, most variables are used in levels, because they are already defined in terms of gaps or growth rates and thus are stationary. Inflation expectations are the exception, since they inherit the nonstationarity of trend inflation, so they are entered in differences. We also exclude measures of real activity, i.e., the output gap and the unemployment gap, since we do not expect changes in these variables to have a permanent impact on the inflation trend.

The results are presented in Table 3. While several variables are formally statistically significant (in the sense of a joint test of whether all the $XVAR$ coefficients are zero), if we focus on the sum of the coefficients as a measure of the cumulative impact on the trend component of inflation, the most important predictors are inflation expectations, especially those of financial analysts. A permanent jump of 1 pp in analysts' expectations (i.e., a one-time shock to the difference) would be associated with a cumulative increase in trend inflation of more than 6 basis points over four quarters. For comparison, during most of our sample, trend inflation was falling at a rate of merely 10 basis points per year. Of course, given the reduced-form nature of our estimates, it is difficult to say whether changes in expectations cause the fall in trend inflation, or whether market participants simply anticipate future changes in the trend.

Table 3: Regression of Change in Trend on an Additional Variable

	xsum	pval	R^2
GCZKEUR	-0.09	0.08	0.04
	(0.09)		
ΔIEFA	6.17***	0.04	0.13
	(2.11)		
ΔIEFM	4.07***	0.02	0.12
	(1.50)		
ΔIECNB	2.13*	0.21	0.04
	(1.15)		
GM2	-0.04	0.99	0.00
	(0.11)		
SPREAD	1.76	0.00	0.20
	(1.73)		
GCRED	-0.00	0.28	0.01
	(0.09)		
GOIL	0.01	0.63	0.01
	(0.01)		
GWAGE	-0.06	0.13	0.05
	(0.26)		
GEZPPI	0.05	0.32	0.02
	(0.12)		

Note: HAC-robust standard errors in parentheses (* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$). Each row displays the results of a regression as in equation (2) with a different additional variable. The "xsum" column presents the sum of the coefficients for 4 lags of XVAR. The "pval" column presents a joint test of whether the coefficients of all lags of XVAR are zero.

In Appendix C, we repeat the analysis using core inflation as the main dependent variable. The trend component of core inflation evolves similarly to the headline trend, although its overall level is somewhat lower. The results from the regressions are also broadly consistent with the headline estimates, but they do indicate a somewhat higher importance of domestic variables.

5. Policy Implications and Conclusion

In this paper, we explore a potential way of modeling the Czech inflation process using simple unobserved component stochastic volatility models. The first step of our analysis draws on the inflation series as the only observable series entering the analysis. Inflation is decomposed into a permanent component, which can be viewed as the inflation trend, and a transitory component that exhibits some moderate degree of persistence. The model does not assume any inflation trend – the trend is estimated. Moreover, no rational expectations are imposed.

The next step consists of a regression analysis of the two inflation components. The set of regressions allows for economic interpretation of the two inflation components. It turns out that the trend coincides with changes in inflation expectations, and the cyclical component to a great extent follows an open-economy Phillips curve.

The results can be contrasted with similar exercises conducted for the US and UK, which found no robust link of the real economy and inflation expectations to prices as suggested by the Phillips curve. The greater importance of these variables found in the case of the Czech Republic provides evidence for heterogeneity in inflation dynamics across economies, which should be taken into account by economists and policymakers.

Our analysis provides an alternative view of the inflation process and the nominal side of the economy in the Czech Republic to that stemming from a standard New Keynesian DSGE model. On the one hand, observed inflation, which is often below the inflation target, is consistent with partially adaptive expectations and with a trend residing below the target. This is the interpretation based on the AR-UCSV model. On the other hand, the same observed inflation can be interpreted by a DSGE model with an assumed inflation target and rational expectations. The interpretation stemming from this approach needs more (negative) structural shocks to explain observed below-target inflation. The two pictures of inflation dynamics based on the two models presumably provide different policy recommendations.

To conclude, papers applying UCSV-type models to US and UK inflation show that such simple frameworks can provide important insights into the dynamics of prices. This paper suggests that this approach is also useful in the context of Czech inflation and could serve as a complementary tool to help understand the inflation process and provide policy recommendations.

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Appendix A: Test of Stochastic Volatility in the Cyclical and Trend Components

To test for stochastic volatility in the inflation components, we employ the standard model of Stock and Watson (2007) in the noncentered parametrization, as, for example, in Chan (2018).[9] The reformulation enables us to test for the presence of stochastic volatility in the trend and cyclical/transitory components directly. The model is as follows:

$$
\begin{aligned}
\pi_t &= \tau_t + e^{\frac{1}{2}(h_0+\omega_h \tilde{h}_t)} \varepsilon_t^{\pi}, \\
\tau_t &= \tau_{t-1} + e^{\frac{1}{2}(g_0+\omega_g \tilde{g}_t)} \varepsilon_t^{\tau}, \\
\tilde{h}_t &= \tilde{h}_{t-1} + \varepsilon_t^{h}, \\
\tilde{g}_t &= \tilde{g}_{t-1} + \varepsilon_t^{g},
\end{aligned}
$$

where $\varepsilon_t^{\pi}, \varepsilon_t^{\tau}, \varepsilon_t^{h}$, and ε_t^{g} are independent $N(0,1)$. We largely follow the estimation setup in Chan (2018) and initialize the states with $\tau_1 \sim N\left(\tau_0, V_\tau exp\left(g_0 + \omega_g \tilde{g}_1\right)\right)$, $\tilde{h}_1 \sim N(0, V_h)$, $\tilde{g}_1 \sim N\left(0, V_g\right)$ for $V_h = V_g = 10$. In contrast to Chan (2018), we assume $\tau_0 = 6$ and $V_\tau = 1$, reflecting the fact of heightened inflation expectations at the beginning of the sample, which enters the estimation in the form of an informative prior. The priors on ω_h and ω_g are normal with zero mean and variances equal to 0.2 to get the prior means on ω_h^2 and ω_g^2 equal to 0.2 as in Stock and Watson (2007). Finally, we assume that $h_0 \sim N(0,10)$ and $g_0 \sim N(0,10)$. The inference is based on 100,000 draws with a burn-in period of 10,000.

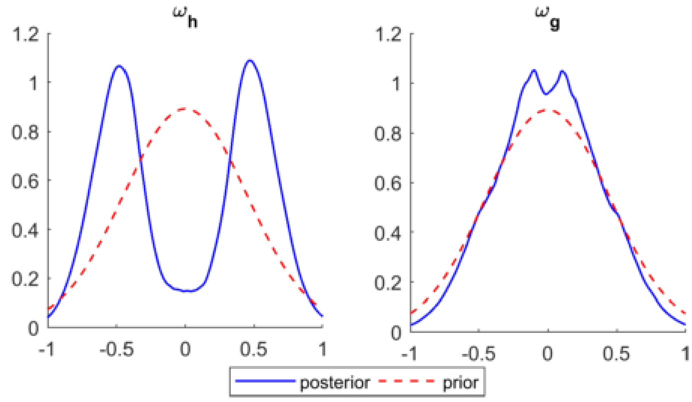
The MCMC algorithm is based on the fact that taking the log of the squared measurement equation (with the LHS variable being equal to $\pi_t - \tau_t$) implies a linear non-Gaussian state space form, with innovations in the measurement equation being distributed according to $\log \chi^2(1)$. Next, mixture of normals approximation is used to obtain an approximate linear and Gaussian state space form. The procedure is described in Kim et al. (1998).

The posterior mean of ω_h^2 is 0.28 and that of ω_g^2 is 0.14, with respective standard deviations of 0.20 and 0.19, suggesting high uncertainty related to the estimated parameters.[10] Note that the squared parameters are estimated (as noted by Chan, 2018, the signs of the two parameters are not identified) and it is difficult to infer on the probability that $\omega_h = 0$ and $\omega_g = 0$, i.e., that the volatility is constant. To get some idea about the variation in the volatility, the following figure shows the posterior densities for the two parameters. It suggests that there is some evidence of stochastic volatility of the transitory component parameter ω_h (little mass near 0). Regarding the stochastic volatility of the permanent component, it turns out that there is little information available in data to infer on this parameter (the prior and posterior densities of ω_g are almost identical).

[9] We use MATLAB code from Chan (2018) for this exercise.

[10] As a robustness check, we estimated the model with the prior on ω_h and ω_g distributed normally with mean zero and variance 0.1 (implying means of ω_h^2 and ω_g^2 equal to 0.1). The reasoning for this prior is that the posterior means of ω_h^2 and ω_g^2 for the G7 countries are closer to 0.1 than to 0.2, as documented in Chan (2018). The posterior mean of ω_h is 0.21 and that of ω_g is 0.10, with respective standard deviations of 0.16 and 0.14.

Figure A1: The Prior and Posterior Densities of $\omega_h$and ω_g



The formal statistical test of the presence of stochastic volatility in the two components is based on the Bayes factors of the unrestricted model and its restricted versions (i.e., the versions with $\omega_h = 0$, $\omega_g = 0$, or both). The Bayes factor is computed using the Savage-Dickey density ratio, i.e., the ratio of the prior to the posterior of the full model evaluated at the point constituting the restriction ($\omega_h = 0$, $\omega_g = 0$, or both). The details of the computation can be found in Chan (2018), Appendix A.

The log Bayes factors suggest a preference for the model with stochastic volatility in both components over the versions with stochastic volatility in the trend component only and without stochastic volatility (the log of the Bayes factor for the first comparison is 2.2 and that for the second is 24.6). At the same time, there is very weak evidence that stochastic volatility in the trend component is not favored by the data (the log of the Bayes factor of the unrestricted model to the model with constant volatility in the trend component is -0.1).

Appendix B: Additional Results of the Regression Analysis

Table B1: Summary Statistics

	count	mean	sd	min	max
INFL	87	2.32	2.31	-1.97	11.68
TREND	87	2.42	0.43	1.86	3.38
OUTGAP	87	-0.08	1.72	-2.67	4.39
UGAP	87	0.18	1.08	-2.00	1.73
GLOBGAP	87	0.03	1.25	-3.36	3.56
GCZKEUR	86	1.74	8.67	-32.74	19.41
IEFA	83	2.69	0.87	1.43	4.80
IEFM	83	2.71	0.99	1.30	5.00
IECNB	62	2.35	0.82	1.07	5.00
GM2	71	7.17	5.13	-9.12	19.63
SPREAD	64	1.49	0.31	0.89	2.35
GCRED	87	4.91	8.65	-36.31	30.59
GOIL	87	11.39	58.06	-206.61	148.00
GWAGE	79	5.15	3.67	-13.51	19.14
GEZPPI	87	1.50	4.19	-13.07	10.38
ΔTREND	86	-0.01	0.03	-0.07	0.06
ΔIEFA	82	-0.03	0.31	-0.93	0.73
ΔIEFM	82	-0.03	0.34	-1.20	1.30
ΔIECNB	61	-0.01	0.57	-1.63	1.58

Table B2: Regression of Inflation on the Trend and an Additional Variable, Partial R^2, Robustness to Lags

	no lags	with 3 lags
OUTGAP	0.205	0.214
UGAP	0.055	0.114
GLOBGAP	0.234	0.240
GCZKEUR	0.081	0.056
IEFA	0.200	0.252
IEFM	0.374	0.479
IECNB	0.099	0.242
GM2	0.050	0.057
SPREAD	0.063	0.105
GCRED	0.004	0.161
GOIL	0.026	0.110
GWAGE	0.055	0.081
GEZPPI	0.153	0.205

Table B3: Regression of Inflation on the Trend and an Additional Variable, Distribution of the Point Estimates for Different Draws of the Trend Component

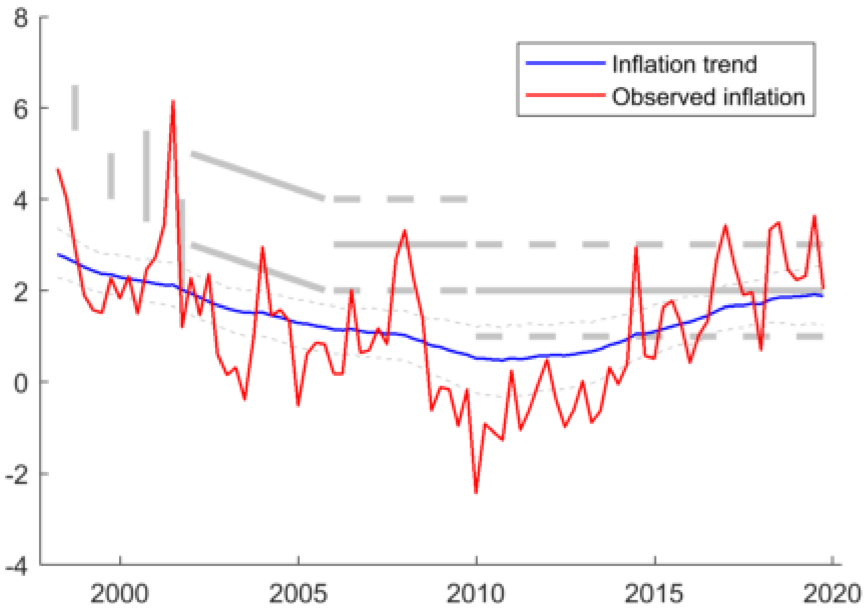
	mean	**sd**	**p5**	**p95**
OUTGAP	0.601	0.045	0.523	0.649
UGAP	-0.410	0.135	-0.713	-0.300
GLOBGAP	0.882	0.031	0.820	0.917
GCZKEUR	0.079	0.004	0.072	0.083
IEFA	1.590	0.303	1.178	2.256
IEFM	1.658	0.251	1.379	2.252
IECNB	1.218	0.116	0.957	1.418
GM2	0.105	0.010	0.088	0.123
SPREAD	-2.795	0.333	-3.128	-2.033
GCRED	0.013	0.008	-0.001	0.019
GOIL	0.007	0.001	0.006	0.008
GWAGE	0.143	0.013	0.119	0.158
GEZPPI	0.217	0.012	0.198	0.226

Appendix C: Core Inflation

The analysis in the main text concerns general inflation as represented by the consumer price index. The index contains some volatile items, and their presence can distort the analysis. Such volatile items are often outside the purview of monetary policy. They can, however, affect inflation expectations. To infer on the role of such items, a measure of core inflation is considered in this section. The core measure is adjusted inflation excluding fuels, which is an index of non-food items of the consumer basket excluding items with regulated prices, administrative interventions, and fuels. In what follows, we discuss the results based on the AR-UCSV model.

Similarly to the case with CPI inflation, the simpler UCSV model suggests the version of the model with a stochastic volatility component in the cyclical component only.

***Figure C1:** Quarter-to-Quarter Percentage Change in Seasonally Adjusted Inflation Rate Excluding Fuels in the Annualized Rate and the Estimated Trend in the AR-UCSV Model*



Note: 68% confidence interval indicated by thin dashed lines. Thick lines indicate the inflation target (the target range). Note that net inflation was targeted until 2001. Since 2002, headline inflation has been targeted.

Figure C1 shows that the inflation trend is located below the target or the tolerance band for the target over almost the whole period. The situation changed after the introduction of unconventional monetary policy measures.

Tables C1 and C2 are the core inflation counterparts of Tables 2 and 3 in the main text. Inspecting Table C1 suggests a more diminished role of inflation expectations and the exchange rate but a larger role of the unemployment gap in determining the cyclical component as compared to the results with headline inflation. Table C2 implies that changes in the core trend are, in addition to inflation expectations, related to the spread and wage growth. Since energy and food prices are to a large extent determined globally, the higher correlation of domestic variables with core inflation demonstrated by these results is not unexpected.

Table C1: Regression of Core Inflation on the Core Trend and an Additional Variable

	const	trend	xvar	partial R^2
OUTGAP	-0.93^{***}	1.71^{***}	0.23^{***}	0.16
	(0.18)	(0.15)	(0.04)	
UGAP	-0.64^{***}	1.53^{***}	-0.41^{***}	0.18
	(0.15)	(0.10)	(0.09)	
GLOBGAP	-0.85^{***}	1.63^{***}	0.33^{***}	0.16
	(0.22)	(0.20)	(0.05)	
GCZKEUR	-0.89^{***}	1.63^{***}	0.02^{*}	0.02
	(0.23)	(0.22)	(0.01)	
IEFA	-0.89^{*}	1.72^{***}	-0.02	0.00
	(0.51)	(0.31)	(0.25)	
IEFM	-1.25^{**}	1.60^{***}	0.16	0.02
	(0.52)	(0.29)	(0.23)	
IECNB	-1.85^{***}	2.22^{***}	0.24^{*}	0.05
	(0.33)	(0.10)	(0.13)	
GM2	-1.37^{***}	1.91^{***}	0.04^{*}	0.04
	(0.23)	(0.23)	(0.02)	
SPREAD	0.12	2.06^{***}	-0.83^{**}	0.08
	(0.76)	(0.19)	(0.38)	
GCRED	-0.90^{**}	1.69^{***}	-0.00	0.00
	(0.37)	(0.24)	(0.03)	
GOIL	-0.92^{***}	1.70^{***}	-0.00	0.00
	(0.29)	(0.22)	(0.00)	
GWAGE	-1.18^{***}	1.76^{***}	0.05	0.02
	(0.23)	(0.27)	(0.05)	
GEZPPI	-0.99^{***}	1.71^{***}	0.04	0.02
	(0.24)	(0.20)	(0.03)	

Note: HAC-robust standard errors in parentheses (* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$). Each row displays the results of a regression as in equation (1) with a different additional variable. The partial R^2 column shows what proportion of the variance unaccounted for by the trend only is explained by adding XVAR.

Table C2: Regression of Change in Core Trend on an Additional Variable

	xsum	pval	R^2
GCZKEUR	-0.47***	0.03	0.17
	(0.18)		
ΔIEFA	13.36***	0.00	0.15
	(3.55)		
ΔIEFM	6.57**	0.00	0.10
	(2.57)		
ΔIECNB	7.56**	0.00	0.13
	(3.48)		
GM2	-0.27	0.00	0.06
	(0.17)		
SPREAD	10.09**	0.03	0.33
	(4.48)		
GCRED	0.05	0.13	0.01
	(0.23)		
GOIL	-0.03	0.91	0.03
	(0.03)		
GWAGE	-0.97**	0.00	0.23
	(0.42)		
GEZPPI	-0.22	0.36	0.02
	(0.43)		

Note: HAC-robust standard errors in parentheses (* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$). Each row displays the results of a regression as in equation (2) with a different additional variable. The “xsum” column presents the sum of the coefficients for 4 lags of XVAR. The “pval” column presents a joint test of whether the coefficients of all lags of XVAR are zero.

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