The impact of labor cost growth on inflation in selected CESEE countries

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We analyze the relationship between labor cost and inflation in selected economies in Central, Eastern and Southeastern Europe (CESEE) by using a medium-scale time-varying parameter vector autoregressive model. The proposed framework makes it possible to control for potential movements in the underlying transmission mechanisms, stochastic volatility and flexible model selection. We use our model to simulate the effect of an unexpected labor cost shock and assess the dynamic reaction of inflation over the estimation period. Our findings indicate that a 1 percentage point increase in unit labor cost translates into higher inflation rates in most countries considered. However, the magnitude of the inflation reaction is very heterogeneous across countries and over time: the lowest response was observed for Bulgaria between 2008 and 2012 and the highest median response for Hungary between 2005 and 2007 (more than 0.4 percentage points). Moreover, we find that the wage-inflation pass-through weakened after the global financial crisis for most countries under consideration.

JEL classification: C11, C15, C32, E24, E31
Keywords: inflation, pass-through, labor cost, Bayesian methods, time-varying parameters

A commonly held view in economics is that prices reflect firms' costs of production plus a markup. As labor costs represent a sizable share of total production costs,[2] they should have a direct impact on the price level of an economy based on this notion. Accordingly, a move in labor costs should translate into changes of the inflation rate.

There are several reasons why this relationship does not necessarily hold in the short run. Not all components of the consumption basket are equally sensitive to changes in labor cost: Think of energy prices, which are mostly determined in global markets, food prices, which fluctuate because of weather conditions, housing prices, which depend to a considerable extent on the supply of housing, or the area of regulated prices, which is governed by political decisions. Furthermore, labor is not the only production factor. Changes in the costs of other tangible and intangible production inputs may interfere with wage developments. Finally, firms may decide to compensate for an increase in labor cost by lowering their profit margin to retain market share and/or to avoid costs associated with changing prices. In the longer run, however, a persistent increase in wages should at some point lead to higher inflation.

Even though this proposition has been reviewed extensively in the literature, the empirical results remain inconclusive. Peneva and Rudd (2017), for example, find that changes in labor cost have had only little material effect on price inflation in the U.S.A. in recent years. Depending on the specific measure of compensation used in the estimations, they find that the pass-through of labor cost to prices has either fallen over the past decades or that independent changes in labor cost have had essentially no material effect on inflation in recent years.

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[2] *The share of wage costs in the selected CESEE countries ranges from 73% to 85% (lowest in the Czech Republic and Slovakia and highest in Bulgaria and Croatia). Figures based on 2017 data (Eurostat).*

Bobeica, Ciccarelli and Vansteenkiste (2019), to the contrary, find a strong link between labor cost and price inflation in the four major economies of the euro area and across three main sectors. The relationship, however, is time varying and depends on the state of the economy (especially with regard to the level of inflation) and shocks hitting the economy (see also Forbes et al., 2018).[3]

Empirically, it even remains unclear whether shifts in labor cost precede or follow prices.[4] Knotek and Zaman (2014) report moderate (and over time, declining) correlations between inflation and different wage measures at all leads and lags. In the same vein, Church and Akin (2017) find evidence for both hypotheses in a model using U.S. consumer and producer prices and employment costs. A model calibrated only for core price indexes excluding food and energy suggests that shocks to consumer prices have a significant long-run effect on employment costs but not vice versa. This implies that even the direction of transmission between wages and prices remains ambiguous.

This study sheds new light on the relationship between labor cost and prices by focusing on the experience of selected Central, Eastern and Southeastern European EU Member States (i.e. all CESEE EU Member States except the Baltic countries). These eight countries were chosen because of an especially interesting combination of wage and price developments in recent years: strong wage growth observed against the background of tightening labor markets since 2016 coupled with sustainably low inflation (or even deflation). Average nominal hourly wage growth in the CESEE region accelerated to around 12% in the second half of 2018, far outstripping productivity developments. At the same time, HICP and core inflation remained rather contained at around 2% and 1.5%, respectively. We estimate the effect of an increase in unit labor cost (ULC, i.e. compensation adjusted for productivity) on inflation in the eight selected CESEE countries to assess whether the pass-through has changed over time in the region.

Our research question is addressed through a novel macroeconometric model that allows for drifting parameters and error variances. To circumvent issues associated with overparameterization, we propose using recent shrinkage techniques that push irrelevant predictors toward zero. Moreover, the question of whether coefficients should be time varying or constant is handled through mixture innovation components on the state innovation variances. Since an exact estimation of such a model is unfeasible, we use a straightforward approximation as proposed in Huber et al. (2019) that relies on approximating the mixture indicators during Markov chain Monte Carlo (MCMC) sampling. The resulting MCMC draws are then used to compute the dynamic responses of inflation to ULC shocks.[5]

[3] *In a state-dependent VAR setting, the authors find that the pass-through is stronger at high levels of inflation, while the wage-price link is weak in times of low inflation (or deflation). They also show results for conditional and unconditional forecasting performance finding that, in the four major euro area countries and for most sectors, labor cost has more forecasting power for price inflation than the other way around.*

[4] *From a theoretical perspective, labor cost and inflation are expected to be closely interrelated only in the long run, while, in the short run, firms might be willing to sell at any given price set by the market. Price rigidities, such as menu costs, might also impede sudden price adjustments. Moreover, in New Keynesian models, wages are often determined according to inflation expectations; therefore, depending on whether price or wage rigidities are prevailing, we could expect prices either to follow or lead wages.*

[5] *Time-varying parameter models have been successfully used for forecasting GDP growth in CESEE economies (see, for instance, Feldkircher and Hauzenberger, 2019).*

Our findings point toward a positive, but relatively weak, relationship between ULC growth and inflation for most CESEE countries considered. Furthermore, the responses of price growth to ULC shocks tend to vary strongly across countries and over time. The strongest effect of a 1 percentage point increase of ULC growth on inflation is observed between one quarter and one year after the shock. The median impact of the ULC shock reaches a maximum of more than 0.4 percentage points in Hungary between 2005 and 2007. At the other end of the spectrum, a ULC shock of 1 percentage point translates into a deceleration of inflation by 0.05 percentage points in Bulgaria in the period between 2008 and 2012 (however, those estimates are insignificant). Moreover, we find that most countries experienced a weakening of the pass-through after the global financial crisis. This result corroborates what we see in the data, namely that the effect of the recent strong growth in labor cost on inflation has been rather moderate so far.

This paper is structured as follows: Section 1 describes the data and shows some descriptive statistics. Section 2 introduces the econometric framework, briefly discusses the prior setup and outlines the estimation strategy. Section 3 presents the results including the dynamic responses of inflation to ULC shocks. The final section summarizes the results, elaborates on some further research questions and concludes the paper.

1 Data description

In our analysis, we concentrate on the link between ULC[6] and HICP inflation in eight Central, Eastern and Southeastern European EU Member States: Bulgaria, Croatia, the Czech Republic, Hungary, Poland, Romania, Slovakia and Slovenia. We collected quarterly data over the period from Q1 1995 to Q4 2018. It needs to be noted, however, that in several cases time series are notably shorter (some start only around the year 2000). The dataset includes the following series: the log difference of ULC, the log difference of HICP inflation, real GDP, the nominal effective exchange rate, the log difference of a commodity price index (HWWI index including food, raw materials and energy), oil prices and one-month money market rates. Further details on data sources and data series included in the estimations are provided in the annex.

It should be noted that a considerable number of authors claim that the inclusion of global variables is becoming increasingly important for explaining local consumer price inflation (see, e.g., Borio and Filardo, 2007; Kabukçuoğlu and Martínez-García, 2018; Kamber and Wong, 2018). The inclusion of financial variables and nominal exchange rate and global price indices (such as oil and energy prices) accommodates the exposure of CESEE countries to external shocks in our model.

Chart 1 shows annual changes in ULC and the HICP. In the period under review, price developments were characterized by a broad-based trend of disinflation approximately up until 2005, reflecting economic stabilization after the early years of transition, increased competition (especially at the international level), a shift of monetary policy away from exchange rate stabilization toward inflation targeting

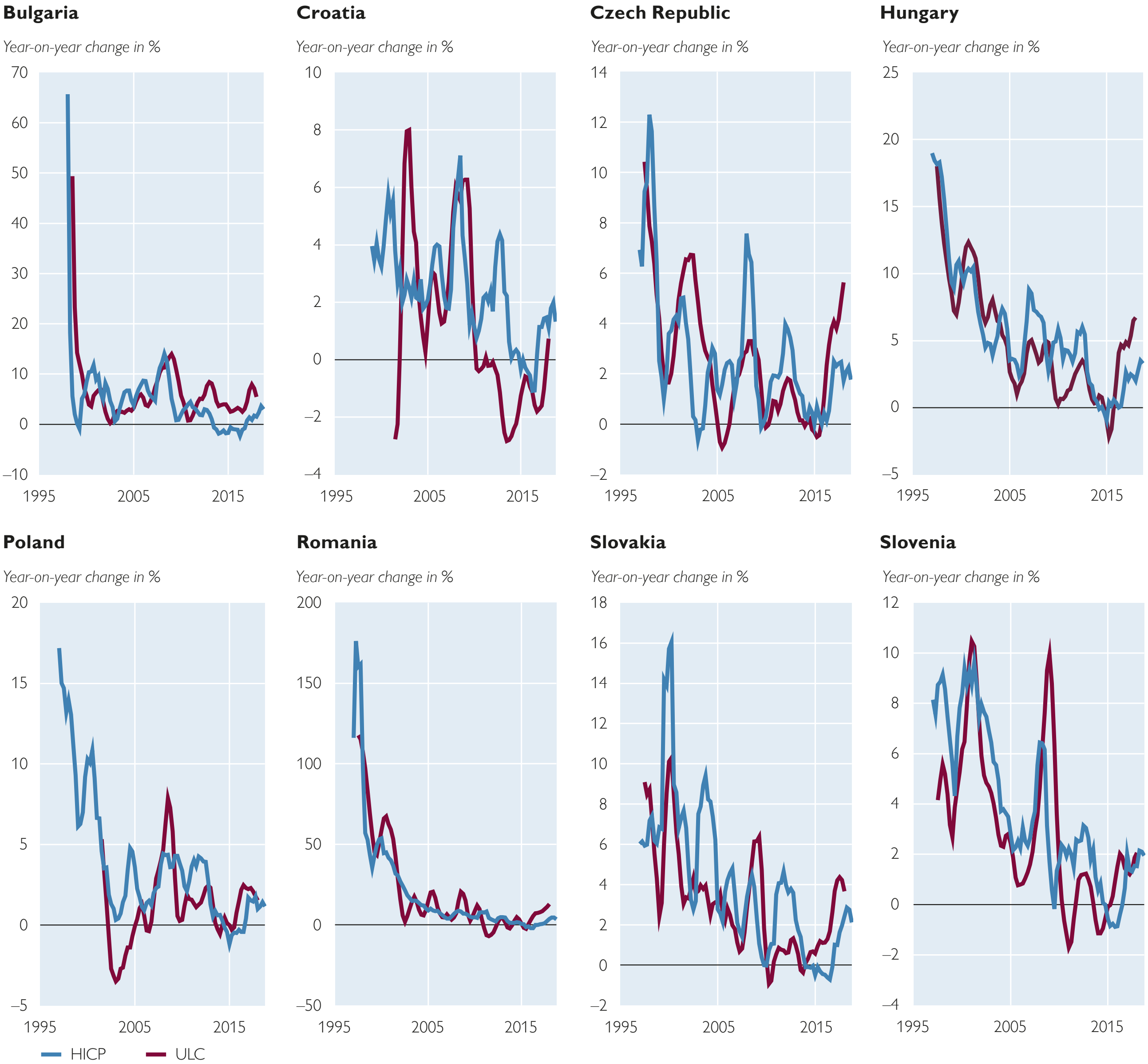
[6] *With unit labor cost (ULC) we refer to the ratio of labor cost to labor productivity on the level of the whole economy. Nominal ULC is calculated as follows: the ratio of compensation of employees to hours worked divided by the ratio of GDP to hours worked. In our analysis, we use ULC instead of compensation of employees since, from a theoretical perspective, only wage increases in excess of productivity growth should put upward pressure on prices. As a robustness check, we also ran the estimations for a wage (compensation of employees) shock. These results are available upon request.*

in many countries and – later on – a stronger reform momentum in the run-up to EU accession. ULC figures by and large mirrored this downward trend. Compensation of employees, however, tended to grow faster than inflation, implying increasing real wages. The latter tendency was a by-product of the greater economic catching-up process. In some countries, however, strong real wage advances can also be partly related to a cyclical overshooting.

In the boom years around the 2004 EU enlargement round, prices and ULC again trended up notably, reflecting buoyant (partly credit-fueled) domestic demand and record-high GDP growth as well as tightening labor markets amid

Chart 1

HICP and ULC over time

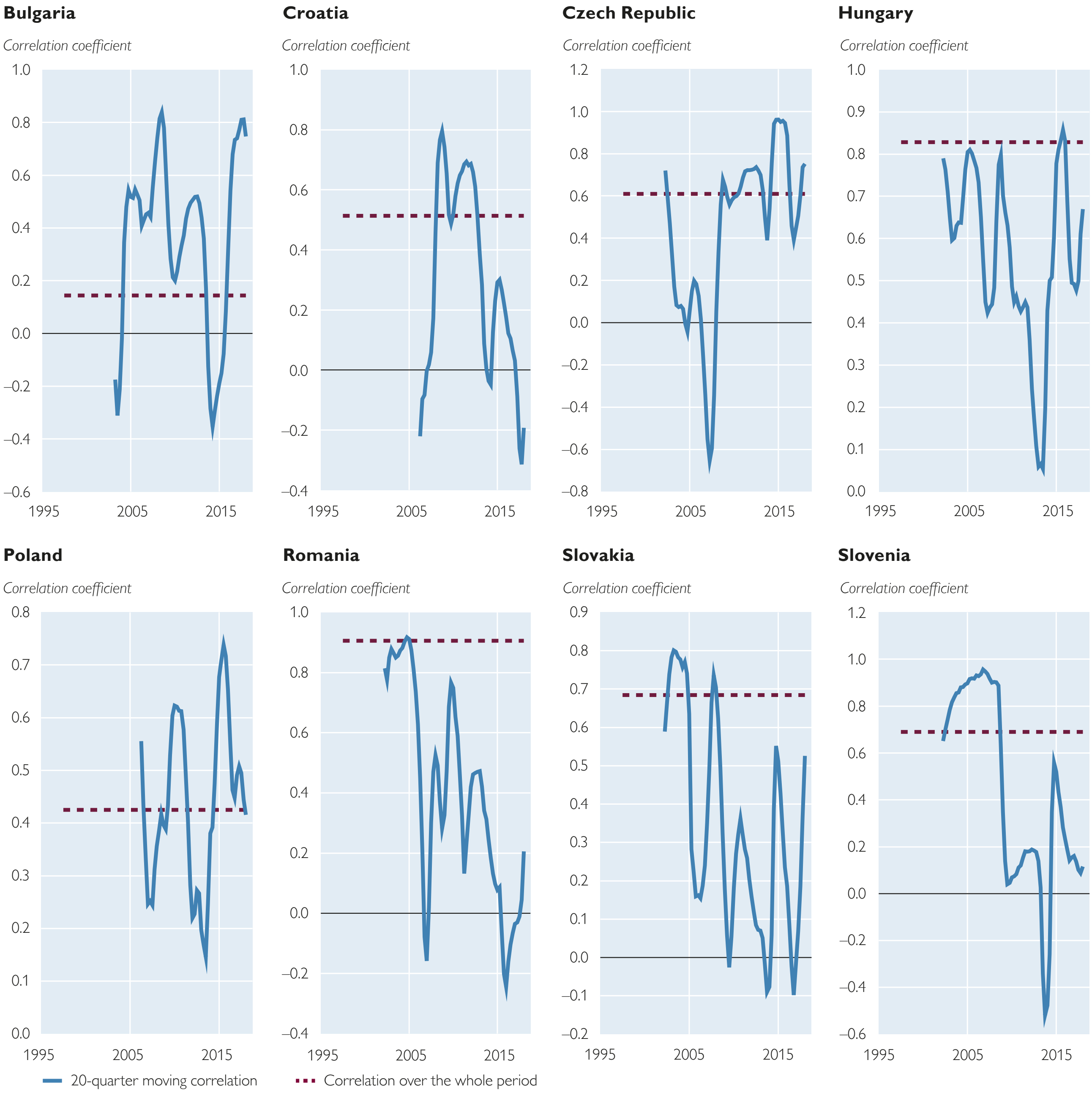


Source: Eurostat.

continuing emigration. The crisis of 2008 and the subsequent years put an end to this phase and sent prices on a downward trend. This trend – temporarily interrupted between 2011 and 2013, when oil prices climbed to above USD 100 per barrel – culminated in a period of deflation around 2015 and 2016. ULC dynamics were heterogeneous. After a notable decline in the aftermath of the crisis, ULC growth again accelerated somewhat until late 2012. At that time, the sovereign debt crisis had sent the euro area into recession for some quarters already, which

Chart 2

20-quarter moving correlation versus correlation over the whole sampling period



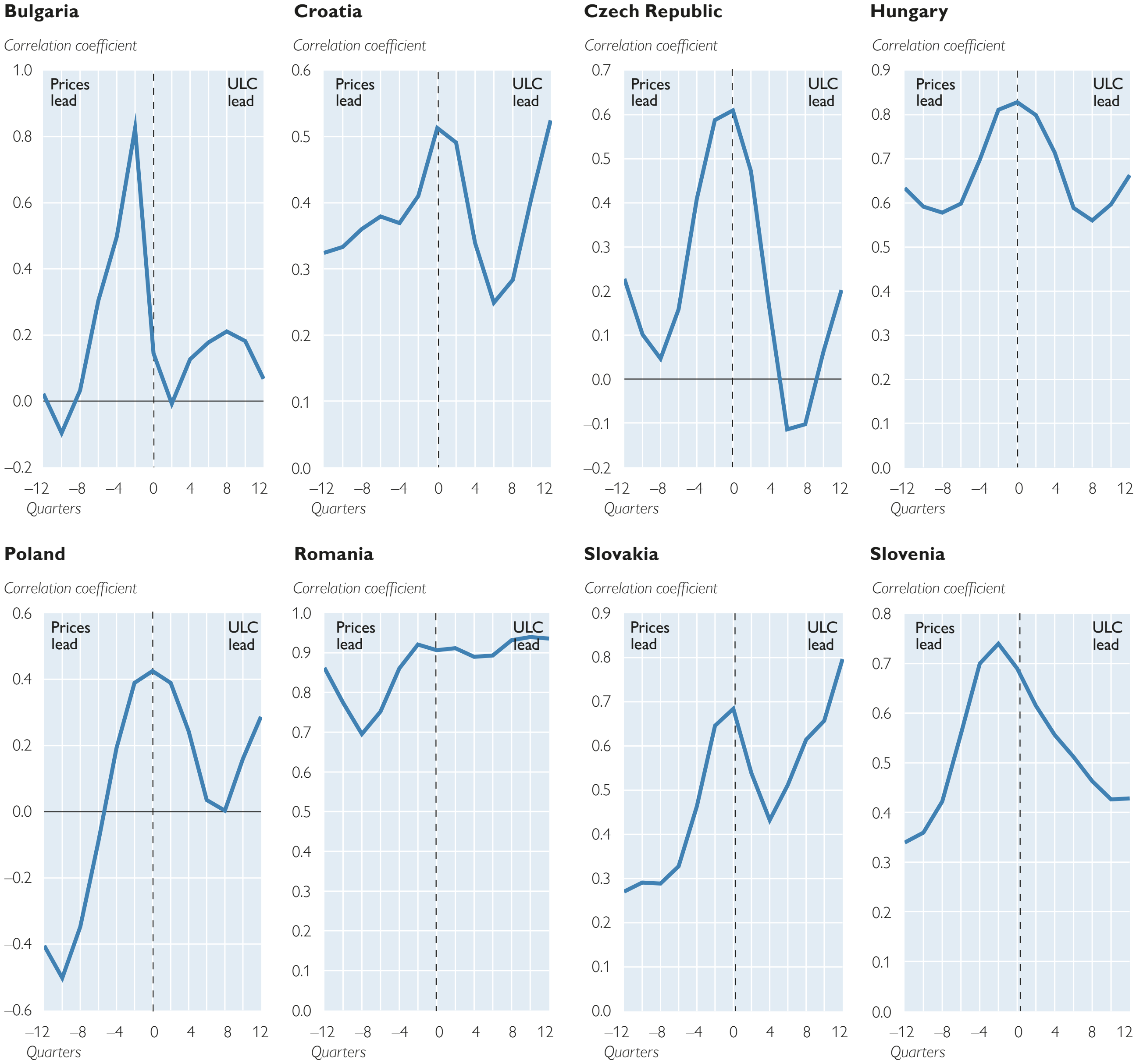
Source: Authors' calculations.

negatively impacted on general economic and wage developments in CESEE. In recent years, ULC dynamics again gained speed and ULC growth reached the highest levels in a decade in many CESEE countries. At the same time, inflation rates remained broadly contained despite an economy in full swing.

Chart 2 shows the correlation between inflation and ULC growth over the whole available time period as well as for a moving window of 20 quarters. Over the whole period, correlations range from 0.14 in Bulgaria to 0.91 in Romania, with the correlation coefficient for the region as a whole amounting to an average of 0.58.

Chart 3

Correlation coefficients of HICP inflation and ULC growth at different leads and lags



Source: Authors' calculations.

For different time frames, the correlations are rather heterogeneous across countries. In a greater regional perspective, however, co-movements are highest in phases of strong economic growth. For the region on average, for example, the correlation coefficient was highest in the five years up to 2008 (0.66), i.e. the boom years after EU accession up to the crisis. Before and after that, correlation was notably weaker. Recently, the average correlation coefficient increased from a low of 0.02 in 2014 to 0.61 in 2018.

Chart 3 shows cross-correlations, which enables a simple examination of the lead-lag structure of the correlations. If increases in compensation systematically come ahead of price inflation in the data, then the strongest cross-correlation would be expected between labor cost inflation in quarter t and price inflation in a later quarter t+x (i.e. higher correlation would appear at point x on the right-hand side of the plots).

This exercise draws a rather heterogeneous picture of the lead-lag structure connecting ULC growth and price inflation. In general, however, the single-strongest correlations usually appear in a specification with no lag or a specification with only a moderate lag of two quarters. Furthermore, we find rather strong correlations in models where ULC growth leads price growth by 12 quarters (e.g. Croatia, Poland, Romania and Slovakia).

2 Econometric framework

To investigate the relationship between ULC growth and inflation in the selected CESEE countries, we consider both the short- and long-run relationships among the two quantities of interest. First, we analyze how the persistent components of ULC and trend inflation co-move over time. Then, we estimate the short-run relationship by considering impulse responses to an increase in ULC growth on inflation over different horizons. For this purpose, we use a time-varying parameter vector autoregressive model with mixture innovations and stochastic volatility (TVP-VAR-SV), which is described in this section. The less technical reader can skip this section and move to section 3.

The dynamic econometric framework adopted is summarized in section 2.1. In section 2.2, we discuss the prior setup used and briefly consider the MCMC algorithm employed.

2.1 TVP-VAR-SV model with mixture innovations

The goal of the present paper is to investigate the dynamic relationship between labor cost growth and inflation in selected CESEE countries. When addressing this research question, we encounter several difficulties that we aim to solve using a novel econometric model. First, for some of the economies we consider in this paper, the length of the time series is rather short (i.e. starting around the early 2000s). This calls for shrinkage techniques to obtain reliable estimates and alleviate overfitting issues. Second, most CESEE countries have undergone structural changes and may have experienced shifts in the wage-inflation pass-through, implying that the parameters of the underlying regression model are possibly time varying. Finally, the estimation period we consider is characterized by major economic shocks that are assumed to affect not only the transmission channels but, more importantly, the magnitudes of the structural shocks.

A model that can handle all three issues raised above is a time-varying parameter VAR model with stochastic volatility (TVP-VAR-SV) proposed in Primiceri (2005) and Cogley and Sargent (2005). This model assumes that the relationship between an M-dimensional vector of macroeconomic quantities y_t evolves according to:

$$y_t = (I_M \otimes x_t')\beta_t + \varepsilon_t, \tag{1}$$

where $x_t=(y_{t-1}',\dots,y_{t-p}',1)$ is a $K(=pM+1)$-dimensional vector that includes the lags of y_t, and a constant β_t is a MK-dimensional vector of time-varying regression coefficients, while $\varepsilon_t \sim N(0,\Sigma_t)$ is a Gaussian white noise shock vector with time-varying variance-covariance matrix Σ_t. In what follows, we decompose Σ_t as follows:

$$\Sigma_t = Q_t H_t Q_t', \tag{2}$$

where Q_t denotes a lower unitriangular matrix (i.e. lower triangular with a unit diagonal) of dimension MxM and $H_t=diag\{exp(h_{1t}),\dots,\exp(h_{Mt})\}$ is a diagonal matrix, while $h_{jt}(j=1,\dots,M)$ denotes the time-varying variances. For simplicity, we store all free elements in Q_t in a v-dimensional vector q_t (with $v = M\frac{M-1}{2}$).

For convenience, we stack the VAR coefficients and the covariance parameters in a $(K+v)$-dimensional vector $\alpha_t=(\beta_t',q_t')'$. Consistent with the literature, α_t evolves according to a random walk process,

$$\alpha_t = \alpha_{t-1} + \eta_t. \tag{3}$$

We let η_t denote a Gaussian error term with $\eta_t \sim N\{0,diag(\theta_{1t},\dots,\theta_{K+v,t})\}$. The error variances in equation (3) are given by:

$$\theta_{jt} = \theta_{j1}d_{jt} + \theta_{j0}(1-d_{jt}), \tag{4}$$

where d_{jt} is a binary indicator that follows a Bernoulli distribution with $Prob(d_{jt}=1)=p_j$ and $\theta_{j1} \gg \theta_{j0}$ are scaling parameters with $\theta_{j0}\approx 0$. This specification implies that if $d_{jt}=1$, the corresponding regression coefficient α_{jt} varies over time, whereas in the case that $d_{jt}=0$, the change in α_{jt} is essentially zero. In what follows, we do not estimate θ_{j0} but set it equal to $\left(\frac{0.1}{4}\right)^2\hat{\theta}_j$, with $\hat{\theta}_j$ denoting the ordinary least squares variance obtained by estimating a linear regression model.

This specification turns out to be a highly flexible variant of a mixture innovation model originally proposed in Gerlach et al. (2000) and Koop et al. (2009). Our model allows for flexible testing whether coefficients should be time varying, constant or a mixture of both (i.e. dynamic over certain periods in time). Notice that if $d_{jt}=0$ for all j and t, we obtain a constant parameter VAR with stochastic volatility. The indicators make it possible to obtain a parsimonious model specification. This feature is crucial for our present application, since the number of quantities in y_t is moderate and the length of the time series rather short.

Our model is completed by the assumption that the logarithm of the error variances follows a random walk with constant error variances. This captures the notion that the log volatilities feature a rather smooth evolution through time.

2.2 Bayesian prior setup and estimation

In this section, we briefly sketch our estimation strategy. Since the model outlined in section 2.1 is heavily parameterized and the likelihood function of the model is difficult to optimize, we follow a Bayesian approach. This implies that we have to specify prior distributions on all key parameters.

Starting with the initial state α_0, we use a Normal-Gamma prior in the spirit of Griffin and Brown (2010) that flexibly pushes elements in α_0 to zero. For a given element in α_0, this prior is given by:

$$\alpha_{0j}|\tau_j \sim N\left(0, \frac{\tau_j}{2\lambda}\right),\ \tau_j \sim G(c_0, c_0), \lambda \sim G(e_0, d_0), \tag{5}$$

where τ_j denotes a local scaling parameter that features a Gamma prior with hyperparameter c_0, while λ is a global shrinkage parameter that also follows a Gamma distribution with parameters e_0 and d_0 a priori. A large λ forces all elements in α_0 to zero, while the presence of the local shrinkage parameters τ_j allows for non-zero $\alpha_{0j}'s$. We follow Huber and Feldkircher (2019) and set e_0=d_0=0.01 and c_0=0.1. Finally, on θ_{j1} and the error variances of the stochastic volatility processes, we use inverted Gamma priors set to be weakly informative.

Since estimating a mixture innovation model with $K+v$ Bernoulli indicators is computationally challenging, we adopt the approximation proposed in Huber et al. (2019) and adopt the corresponding MCMC algorithm. This implies that the full history of d_t is approximated through

$$\hat{d}_{jt}^{(i)} = \begin{cases} 1\ if\ \left|\Delta\alpha_{jt}^{(i-1)}\right| > c_j \\ 0\ if\ \left|\Delta\alpha_{jt}^{(i-1)}\right| \leq c_j \end{cases} \tag{6}$$

during MCMC sampling. Here, we let c_j denote a threshold that features a uniformly distributed prior and the superscript (i) indicates the i^{th} MCMC draw. This approximation captures the notion that if the absolute change in the dynamic regression coefficients exceeds a threshold c_j, we allow for this change by using a high variance of the innovation. By contrast, if it falls below the threshold we use a process innovation variance close to zero and effectively rule out movements in the parameters going from time t–1 to t.

Our MCMC algorithm cycles between the full conditionals that are mostly available in closed form. One exception is the full conditional posterior of c_j, where an inverse transform sampling step is performed. For all empirical results that follow, we use 30,000 MCMC iterations and discard the first 15,000. More details on the proposed model, estimation technique and convergence characteristics can be found in Huber et al. (2019).

3 Empirical results

In this section, we start by describing the time-varying low-frequency relationship between ULC and inflation. This analysis makes it possible to investigate how the persistent components of ULC growth and trend inflation co-move over time. Then, we consider the short-run relationship between ULC growth and inflation by considering impulse responses to a 1 percentage point increase in ULC growth on inflation over different horizons.[7]

[7] *We also ran the estimations for a simple wage shock (compensation of employees) and the results did not change qualitatively.*

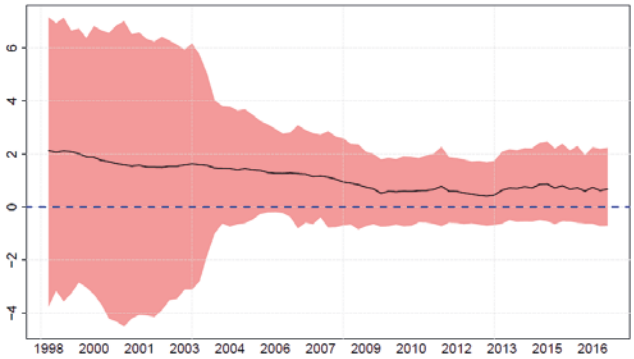
Before proceeding to the empirical analysis, it is worth emphasizing that the time variation in all quantities reported in this section arises from three different sources: changes in the reduced-form VAR coefficients, in the matrix of impact innovations and in the stochastic volatilities. The first two sources potentially impact key transmission mechanisms, while changes in error volatilities mainly reflect movements in uncertainty associated with one-step-ahead prediction errors.

Chart 4.1

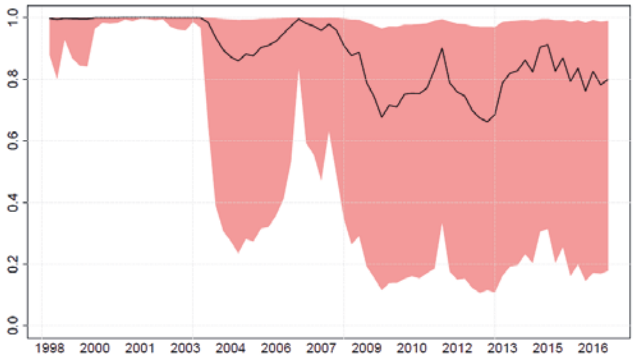
Low-frequency relationship between wages and inflation, and the corresponding measure of fit

Bulgaria

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation

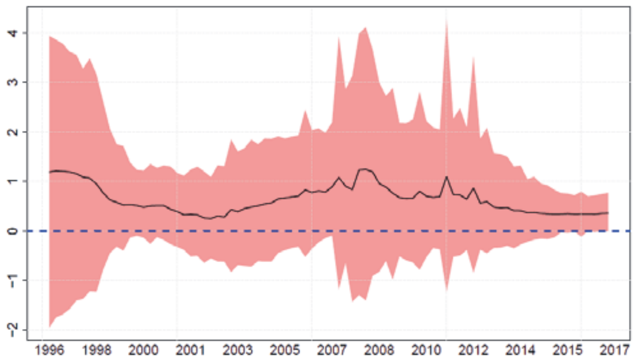


R-squared of wages over price inflation

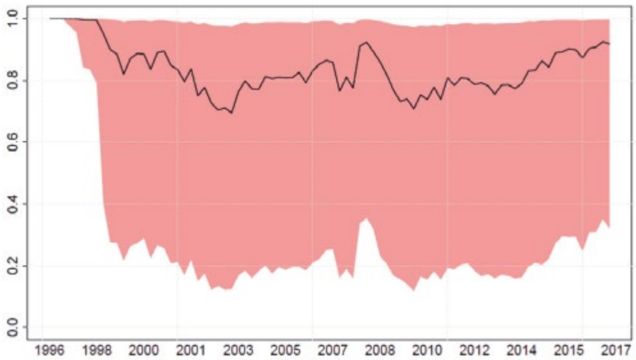


Czech Republic

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation

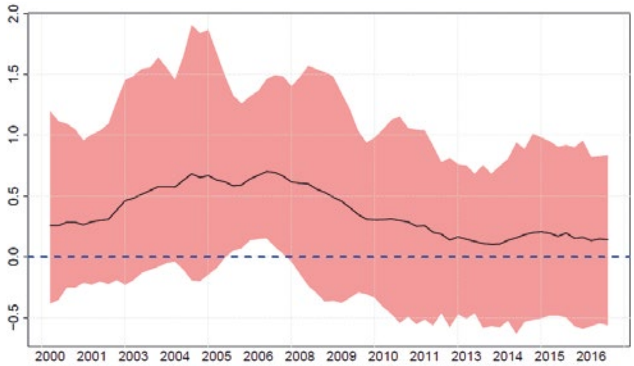


R-squared of wages over price inflation

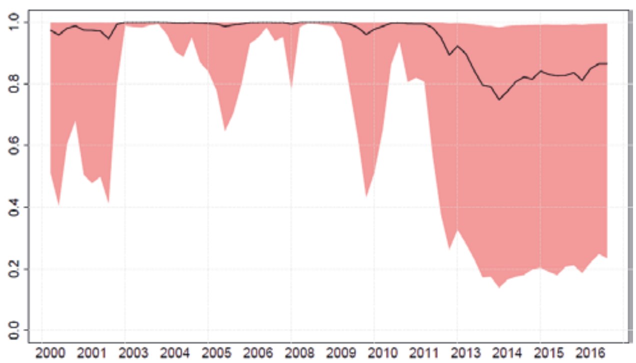


Croatia

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation



R-squared of wages over price inflation



Source: Authors' calculations.

Charts 4.1 to 4.3 show the low-frequency relationship between ULC growth and inflation (i.e. the long-term evolution of the regression coefficients of wages on inflation) on the left-hand side, together with the corresponding measure of fit, the squared coherency, on the right-hand side (for further details, see Sargent and Surico, 2011; Kliem et al., 2016).[8]

The plots on the left-hand side generally show a positive persistent relationship between ULC growth and inflation. However, the relationship is not equally strong in all countries. Over the whole observation period, we see the lowest coefficients for Poland and Croatia (below 0.5) and the highest coefficient for Bulgaria (around 2). Furthermore, coefficients change notably over time. In particular, the economic and financial crisis seems to have altered the relationship between ULC growth and inflation. In most countries, we see a weakening of the coefficients around 2008. In Croatia, for example, the coefficient declines from around 0.8 in 2007 to close to zero in 2014 and the following years. We also see an increase in the confidence bands around the coefficients in several countries, indicating a higher level of uncertainty in the estimations (e.g. for the Czech Republic, Croatia, Romania, Slovenia and Slovakia).

A changing relationship between ULC growth and inflation is also illustrated by the plots on the right-hand side of charts 4.1 to 4.3. These charts depict the coherency measure, a measure that is akin to the R-squared of a regression model, and show how much of the variation in price inflation is explained by ULC dynamics. We find that ULC was driving the variation in inflation in the years up to the crisis in most countries. Exceptions include Poland and the Czech Republic, where the R-squared is notably more volatile and the confidence bands are much larger than in the other countries of the CESEE region. At around 2008, ULC growth loses explanatory power for inflation variation in all countries. At the same time, confidence bands in the plots of the R-squared measure increase notably for most countries.

In what follows we describe the country-specific time-varying impulse responses by inflation to a 1 percentage point increase in ULC growth to assess whether the magnitude of the pass-through has changed over time. For the identification we employ Cholesky ordering with zero restrictions. Following Bobeica et al. (2019), we assume that inflation reacts with a lag to movements in ULC. In addition, we allow financial variables to react within the same quarter, following the assumption that financial markets react more quickly to shocks in the economy than consumer prices.[9]

Due to the time-varying nature of the model, we plot the distribution of the responses over the estimation sample and for four different time horizons: one quarter, one year, two and three years. The responses of inflation to a 1 percentage point increase in ULC growth are presented in charts 5.1 and 5.2. Additional responses showing the subsequent reaction of ULC to a ULC shock can be found in the annex.

[8] *We can rewrite our model in state-space form and compute the spectral density matrix at frequency zero in time t. These figures can be considered an approximation of the sum of distributed lag coefficients of a two-sided least-squares projection of order infinity of inflation on ULC growth. Another feasible approach would be to estimate the low-frequency component of inflation and ULC by relying on a filter and then compute the slope of a scatter plot between the two low-frequency components (see Whiteman, 1984; Kliem et al., 2016). This approach, however, would make it difficult to spot any time variation in the low-frequency relationship.*

[9] *Robustness checks were carried out with respect to different orderings of the endogenous variables to identify the shock. It should be noted that the one-year-ahead impulse responses to a 1 percentage point increase in ULC do not change when altering the ordering of the variables.*

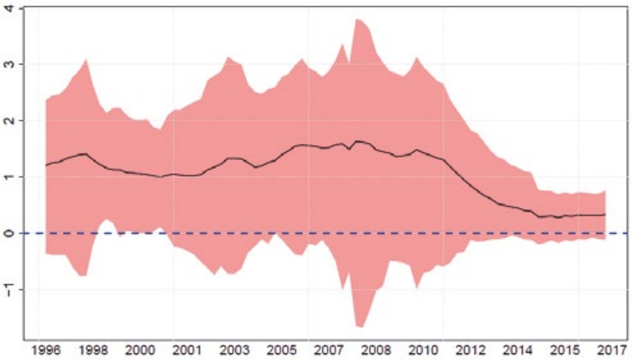
Looking at the reaction of inflation to a 1 percentage point increase in ULC growth (charts 5.1 and 5.2), we observe that the responses are highly time- and country-dependent. In Bulgaria and Slovenia, the responses remain insignificant over the whole sample and for all horizons, indicating a weak link between wages and price inflation. Moreover, in Bulgaria, the median response of inflation turns

Chart 4.2

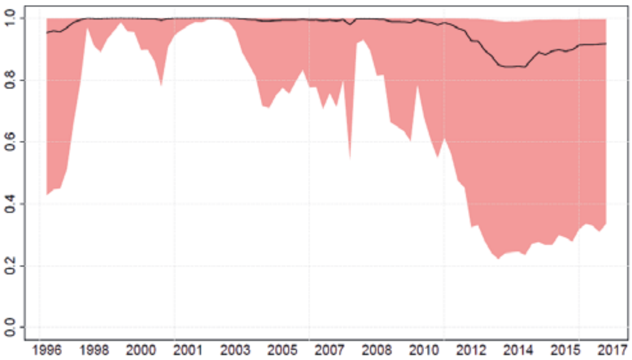
Low-frequency relationship between wages and inflation, and the corresponding measure of fit

Hungary

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation

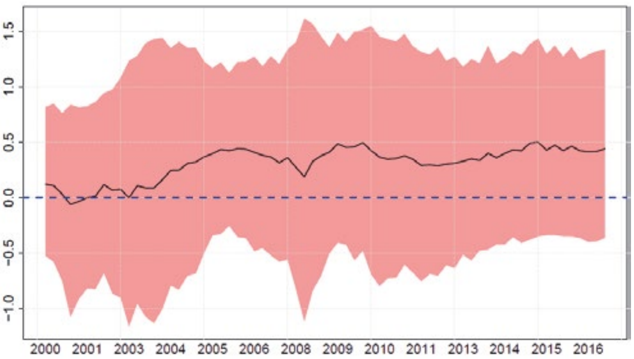


R-squared of wages over price inflation

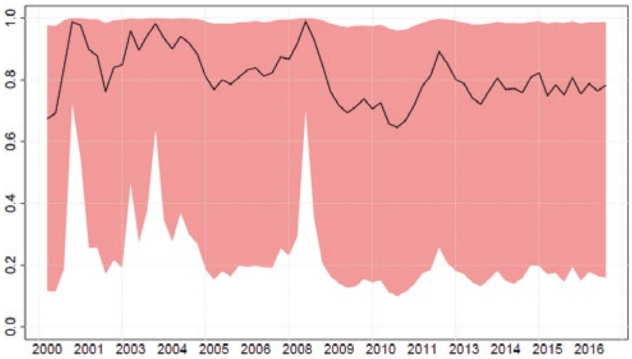


Poland

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation

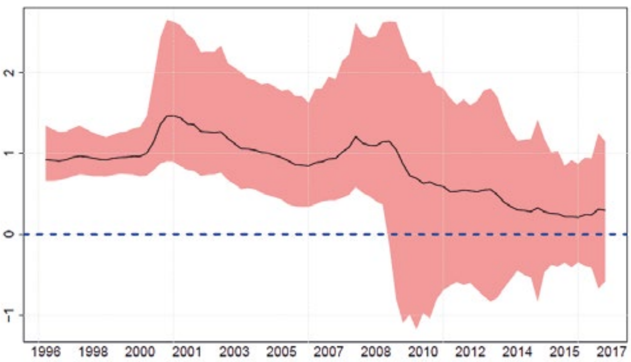


R-squared of wages over price inflation

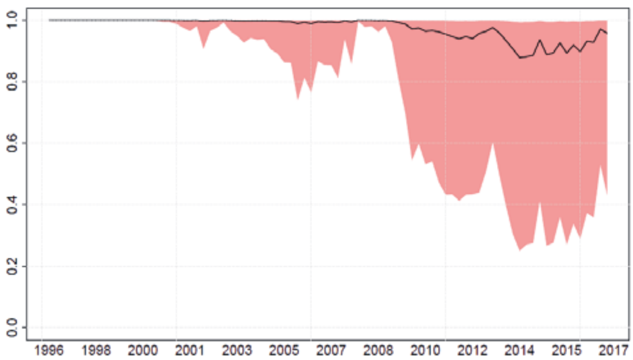


Romania

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation



R-squared of wages over price inflation



Source: Authors' calculations.

negative (approaching –0.05 percentage points) after the global financial crisis, indicating the most notable failure of the wage-inflation pass-through among the countries in the CESEE region. The peculiar responses of inflation to a ULC shock in Bulgaria could be due to the sustained wage growth that the country experienced, despite the long deflationary episode which lasted from 2013 to 2016. In Slovenia, the median responses remain very low over the whole sample but show slightly stronger and more persistent reactions in 2006 and in the last two to three years of the sample, due to the higher persistence of the labor cost shock (see chart A1.2 in the annex).

In the remaining countries, the link between labor cost and price inflation is more pronounced. In the Czech Republic, Hungary and Slovakia, the response of inflation to ULC shocks is positive and mostly significant already after the first quarter, while in Croatia, Poland and Romania, the reaction is slower and becomes significant only one year after the shock.

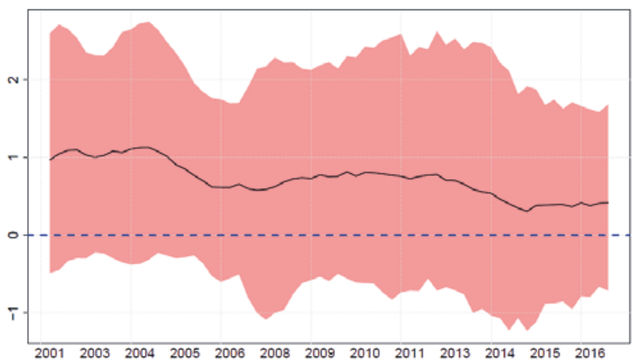
Median responses of inflation reach their maximum magnitude of about 0.4 percentage points in Hungary, followed by the Czech Republic, which shows a reaction of above 0.2 percentage points. The highest responses are mostly observed in the period between 2005 and 2008, but remain high in the Czech Republic up to

Chart 4.3

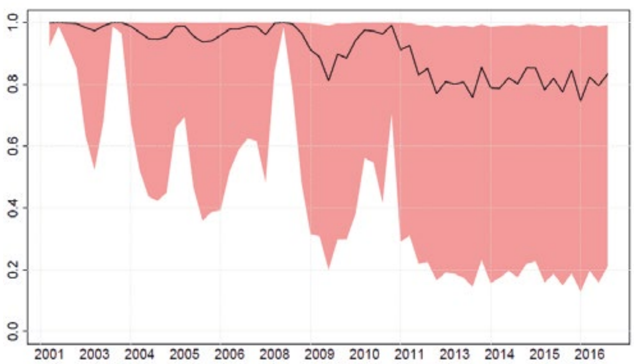
Low-frequency relationship between wages and inflation, and the corresponding measure of fit

Slovenia

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation

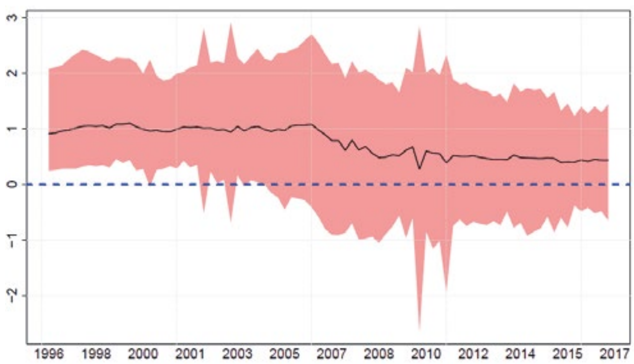


R-squared of wages over price inflation

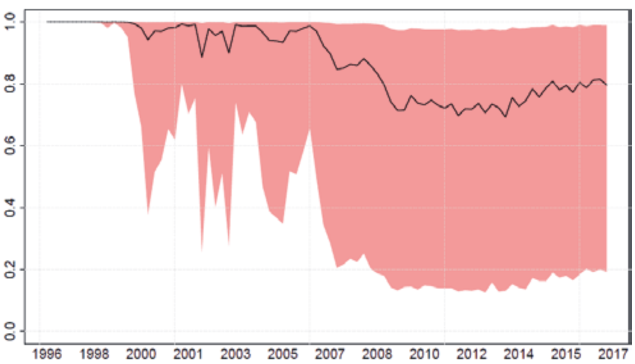


Slovakia

Median and 90% posterior percentiles of the time-varying coefficients of wages on inflation



R-squared of wages over price inflation



Source: Authors' calculations.

the end of the sample. Moreover, the shock displays relatively high persistence in the Czech Republic: Our simulations show strong and significant reactions of inflation (above 0.2 percentage points) at the two-year and three-year horizons in recent years. The Czech Republic is the only CESEE country with such long-lasting inflation reactions to labor cost shocks. This phenomenon is explained by the persistence of ULC shocks in recent years: As shown in chart A1.1 in the annex, ULC remains notably high even three years after the shock. In most of the other countries under observation, the shock has practically faded out after two years already.

Finally, in Poland, the effects of labor cost shocks appear most significant between 2005 and 2011, showing their peak one year after the shock and fading out quickly within the following year. In Romania, on the other hand, inflation reacts relatively strongly in the late 1990s and in the early 2000s, reaching a peak of 0.15 percentage points at the one-year horizon, but showing insignificant reactions otherwise.

As a general observation, prices tend to react less to ULC changes in the period after the crisis. This is true for most countries and most time horizons following a simulated shock (e.g. in Bulgaria, the Czech Republic, Croatia, Hungary and Poland). Furthermore, estimates tend to become less significant after 2008 in virtually all countries under observation. This provides further evidence for the weaker pass-through of wages to prices after the crisis, which we already observed in charts 4.1 to 4.3. Our results corroborate the findings in the literature claiming that the relationship between labor cost and price inflation varies over time and especially depends on the level of inflation: At high levels of inflation the pass-through is stronger, whereas during (and after) times of low inflation (or deflation) the wage-price link is weaker.

Taylor (2000) finds low inflation to be associated with lower expected inflation persistence and with a weaker wage-price and exchange rate pass-through. In fact, the degree to which firms match increases in marginal costs depends on how permanent these changes are (expected to be). When inflation is lower and more stable, changes in wages and prices are expected to be only temporary and firms will pass on less of these changes. This also implies that, under monetary policy aimed at price stabilization, the persistence of deviations of inflation from its trend is expected to be lower, and the pass-through will be weaker. It is worth noting that most of the selected countries show a strong decline in trend inflation in the late 1990s and early 2000s. Policy changes aimed at inflation stabilization in CESEE countries might partly justify the lower explanatory power of wage growth on inflation observed in most countries in the second part of the time series (see charts 4.1 to 4.3).

The weaker pass-through following low levels of inflation might also be explained with downward wage rigidities and different degrees of economic uncertainty which might drive firms' decisions to buffer increases in labor cost with profit margins (Daly and Hobijn, 2014, and Bobeica et al., 2019). In fact, in times of high inflation, firms are more likely to pass through higher production costs to prices, especially if they expect rising interest rates to squeeze future profit margins.[10]

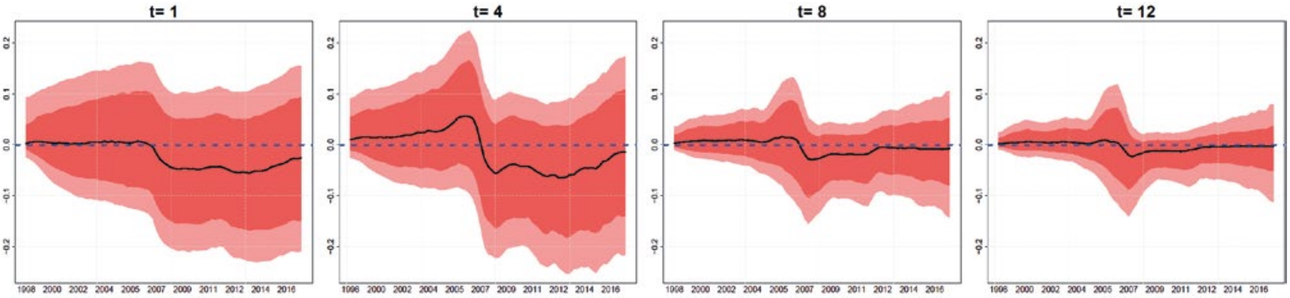
[10] *The opposite might hold if interest rates are expected to decrease when inflation is low (Bobeica et al., 2019).*

Chart 5.1

How does a 1 percentage point increase in ULC affect HICP inflation over time?

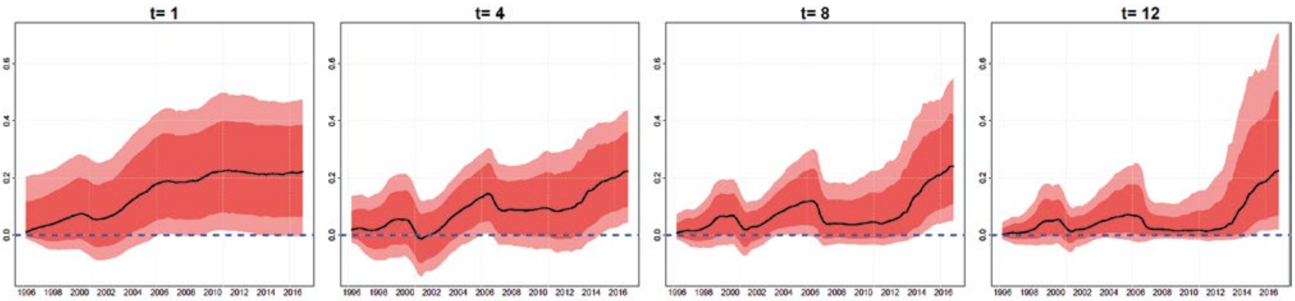
Bulgaria

Impulse responses after one quarter, one, two and three years



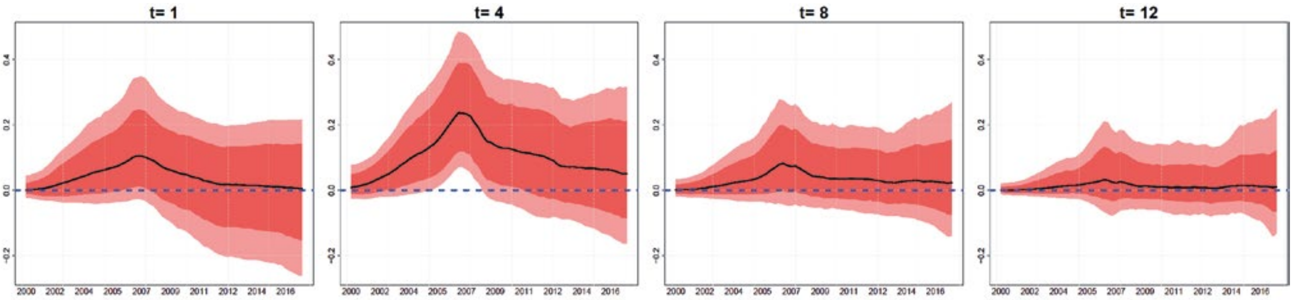
Czech Republic

Impulse responses after one quarter, one, two and three years



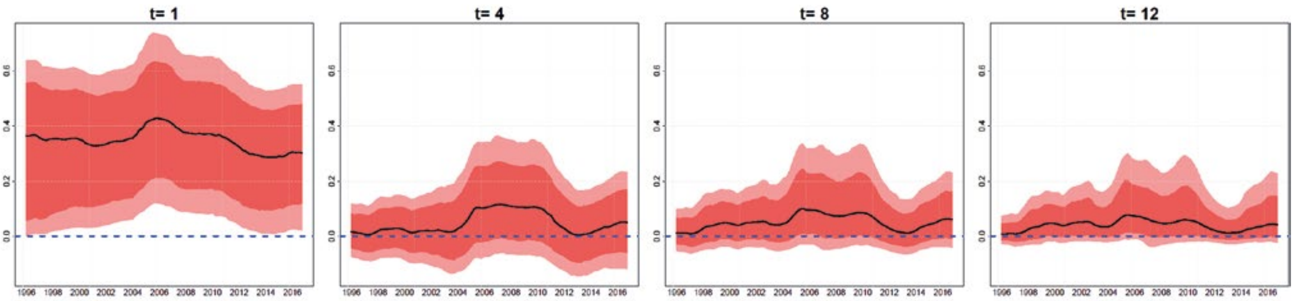
Croatia

Impulse responses after one quarter, one, two and three years



Hungary

Impulse responses after one quarter, one, two and three years



Source: Authors' calculations.

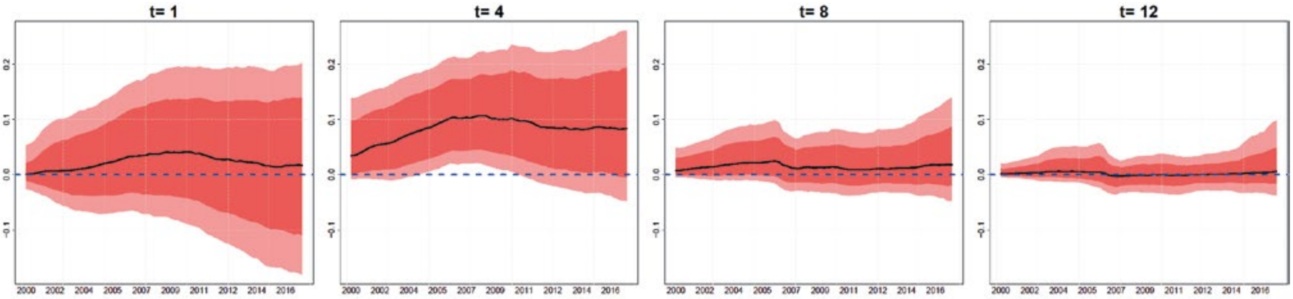
Note: Black lines represent the median response of inflation, whereas dark- and light-red shaded areas represent the 68% and 90% confidence bands.

Chart 5.2

How does a 1 percentage point increase in ULC affect HICP inflation over time?

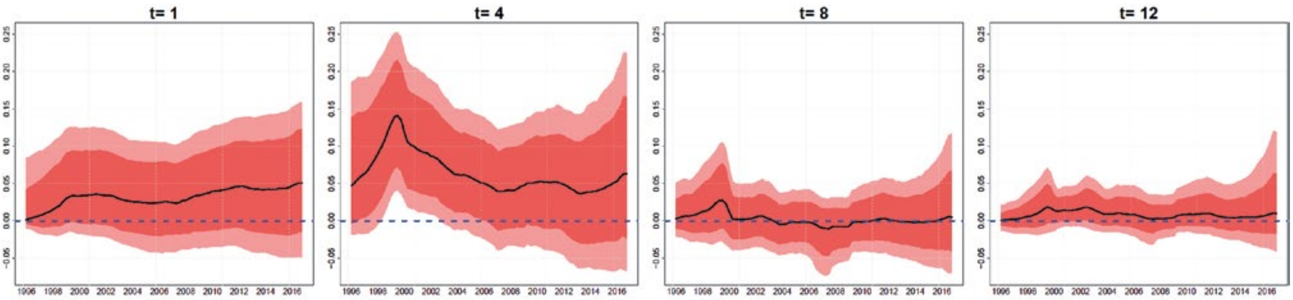
Poland

Impulse responses after one quarter, one, two and three years



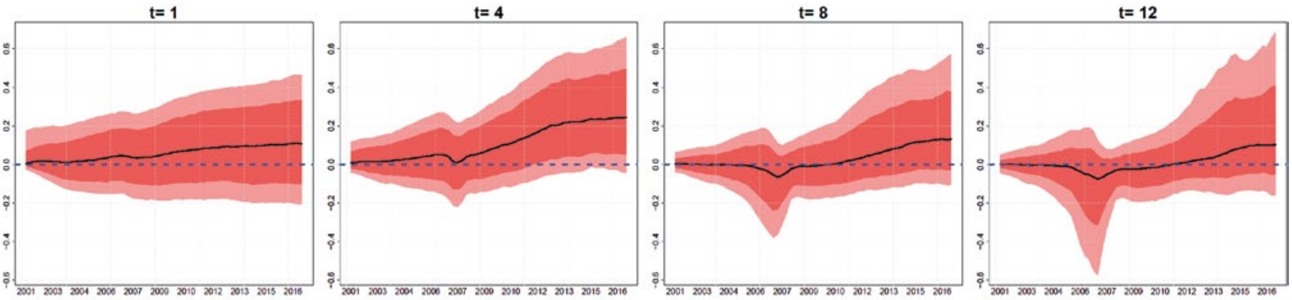
Romania

Impulse responses after one quarter, one, two and three years



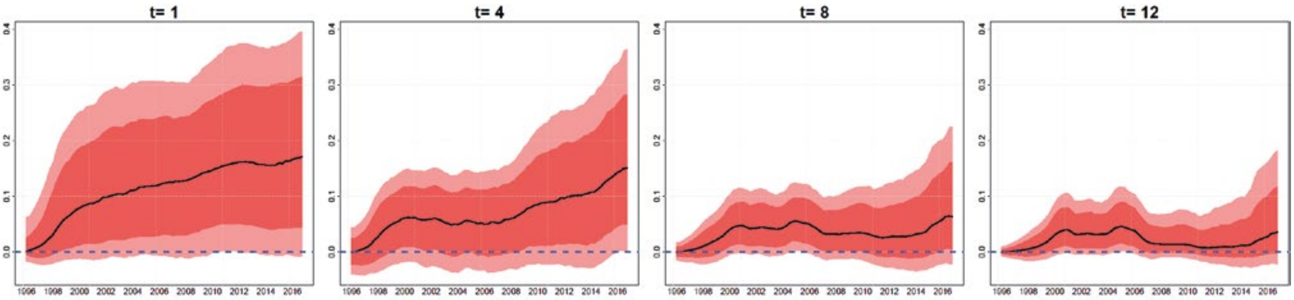
Slovenia

Impulse responses after one quarter, one, two and three years



Slovakia

Impulse responses after one quarter, one, two and three years



Source: Authors' calculations.

Note: Black lines represent the median response of inflation, whereas dark- and light-red shaded areas represent the 68% and 90% confidence bands.

Finally, Head et al. (2010) explain the sensitivity of pass-through to the level of inflation with different degrees of price dispersion in a high- and low-inflation environment and the varying impact of specific degrees of price dispersion on consumers' search intensity.[11]

Among other possible reasons for the weakening of the link between labor cost and inflation in the selected CESEE countries, the following factors should be taken into consideration: (1) Some of the newly created purchasing power from higher wages was used for imports in recent years, which lowered domestic price pressures. (2) Increased import competition did not allow companies to fully pass on higher wages to prices, which was mirrored in lower profit rates. (3) Nominal effective exchange rates appreciated from early 2012, which has led to associated quantity and price effects for imported goods. (4) Several countries reported higher saving rates, and (5) lower net inflows of labor income from abroad. Both of the latter two factors drained purchasing power from the market and thereby lowered inflationary pressure. Furthermore, at the onset of the crisis, ULC growth accelerated due to a contraction in output alongside a much lesser decline in employment and some downward rigidity of wages. Typically, firms treat such recession-caused ULC rises as temporary and refrain from raising prices (also to retain market share).

4 Conclusions

In recent years, many economies have experienced a weakening of the pass-through from labor cost to price inflation. This phenomenon seems to be related to times of recession and low inflation. It is also particularly common in emerging economies, where productivity growth is often followed by rapid wage increases, while price inflation is dampened by international competition.

In this paper we study the link between labor cost and price inflation in selected CESEE countries through a macroeconometric model that makes it possible to integrate drifting parameters and error variances. We use recent shrinkage techniques to circumvent issues associated with overparameterization and include mixture innovation components to make it possible to consider time-varying responses of inflation to ULC shocks.

Using country-specific quarterly data over the period Q1 1995 to Q4 2018, we find a positive relationship between ULC growth and price inflation in most CESEE countries under scrutiny. However, results appear to be very heterogeneous across countries and time.

Estimating the effect of a 1 percentage point increase in ULC growth on price inflation, we find that the shock does not lead to any significant increase in inflation in Bulgaria and Slovenia. In Bulgaria, the median response of inflation turns negative and reaches –0.05 percentage points after the crisis. Conversely, inflation reacts strongly to ULC growth in Hungary in the short term. The ULC shock increases price growth in the country by up to 0.4 percentage points after one quarter. In general, the effect of a ULC shock on inflation is highest between one quarter and one year after the shock, and becomes insignificant thereafter. The most important exception from

[11] *In a nutshell, the argument is the following: A low-inflation environment is characterized by a comparatively homogenous price dispersion. Any shock would notably disrupt price structures and consumers would start to compare prices more intensively and/or to look for substitute goods. This limits firms' price-setting power. In a high-inflation environment, to the contrary, price dispersion and consumers' search intensity is already high. The impact of a shock on the price-setting power of firms would therefore be lower.*

this pattern is the Czech Republic, where ULC shocks have strong and significant effects on inflation even after three years (at least in the period since 2014).

Overall, we also find that the link between ULC and price inflation weakened after the global financial crisis. As most countries entered a period of low inflation (or deflation), the above results corroborate findings in the literature claiming that the wage-price pass-through changes during times of low inflation (see, among others, Bobeica et al., 2019; Daly and Hobijn, 2014; Zanetti, 2007; Mehra, 2000). However, our results also show that the pass-through has again strengthened in some countries in the last few years. As mentioned above, this is especially true for the Czech Republic, but some improvement can also be observed for Hungary and Slovakia. This suggests that ULC growth might translate into larger price movements in the CESEE region in the near future. In fact, price readings of early 2019 indicate higher inflationary pressure amid unabated wage and ULC increases.

The change in the pass-through from ULC growth to inflation after the crisis is not fully explained by our paper and possibly related to the interaction of several factors. These include, among others: exchange rate and trade dynamics, international competition and changing profit shares and saving rates. Further research should be carried out in this direction to gain a more comprehensive understanding of inflationary processes in the CESEE region.

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Annex

Table A1

Data description

Indicator	Unit	Seasonal adjustment	Source	Transformation
All-items HICP	index, 2015=100	not adjusted	Eurostat	log, diff
GDP at constant prices	LCmn, CLV2010	not adjusted	Eurostat	log, diff, seasonally adjusted
Nominal unit labor cost per hour, whole economy	index, 2010=100	not adjusted	Eurostat	log, diff, seasonally adjusted
Nominal compensation per hour, whole economy	index, 2010=100	not adjusted	Eurostat	
Compensation of employees	LCmn	not adjusted	Eurostat	
Thousands of hours worked – total employees	000 hours	not adjusted	Eurostat	
GDP per hour worked	index, 2010=100	not adjusted	Eurostat	
GDP at constant prices	LCmn, CLV2010	not adjusted	Eurostat	
Thousands of hours worked – total employees	000 hours	not adjusted	Eurostat	
One-month money market rate, average	%	not adjusted	Bloomberg, Eurostat	
Nominal effective exchange rate	index, 2005=100	not adjusted	Eurostat	log, diff
HWWI index	index, 2015=100	not adjusted	HWWI	log, diff
Crude Oil, Brent, Spot	USD, average	not adjusted	Macrobond	log, diff

Source: Authors' compilation.

Table A2.1

Data description

	BG	HR	CZ	HU
	Period			
All-items HICP	Q1 97 to Q4 18	Q1 98 to Q4 18	Q1 96 to Q4 18	Q1 96 to Q4 18
GDP at constant prices	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 96 to Q3 18	Q1 95 to Q3 18
Nominal unit labor cost per hour, whole economy	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Nominal compensation per hour, whole economy	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Compensation of employees	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Thousands of hours worked – total employees	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
GDP per hour worked	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
GDP at constant prices	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Thousands of hours worked – total employees	Q1 95 to Q3 18	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
One-month money market rate, average	Q1 99 to Q2 18	Q3 97 to Q4 18	Q1 95 to Q4 18	Q4 96 to Q4 18
Nominal effective exchange rate	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
HWWI index	Q1 96 to Q3 18	Q1 96 to Q3 18	Q1 96 to Q4 18	Q1 96 to Q4 18
Crude Oil, Brent, Spot	Q1 95 to Q4 18	Q1 95 to Q4 18	Q1 95 to Q4 18	Q1 95 to Q4 18

Source: Authors' compilation.

Table A2.2

Data description

	PL	RO	SK	SI
	Period			
All-items HICP	Q1 96 to Q4 18	Q1 96 to Q4 18	Q1 96 to Q4 18	Q1 96 to Q4 18
GDP at constant prices	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Nominal unit labor cost per hour, whole economy	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Nominal compensation per hour, whole economy	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Compensation of employees	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Thousands of hours worked – total employees	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
GDP per hour worked	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
GDP at constant prices	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
Thousands of hours worked – total employees	Q1 00 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
One-month money market rate, average	Q1 95 to Q4 18	Q4 95 to Q4 18	Q2 95 to Q4 18	Q1 02 to Q4 18
Nominal effective exchange rate	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18	Q1 95 to Q3 18
HWWI index	Q1 96 to Q4 18	Q1 96 to Q4 18	Q1 96 to Q4 18	Q1 96 to Q4 18
Crude Oil, Brent, Spot	Q1 95 to Q4 18	Q1 95 to Q4 18	Q1 95 to Q4 18	Q1 95 to Q4 18

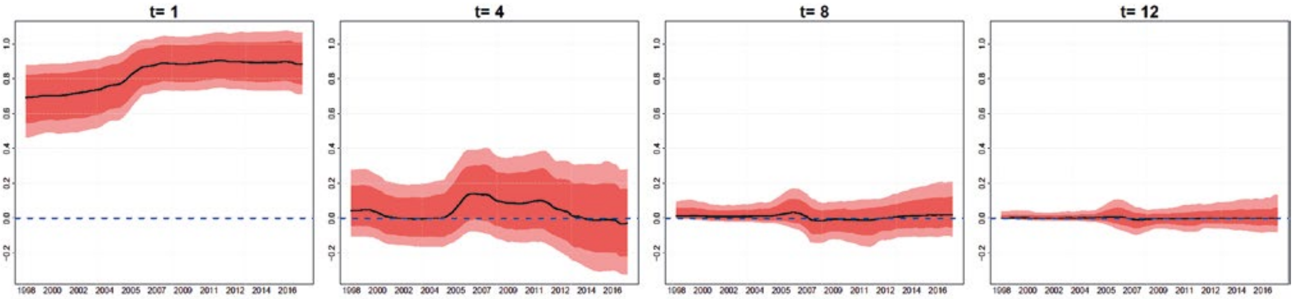
Source: Authors' compilation.

Chart A1.1

How does a 1 percentage point increase in ULC affect ULC over time?

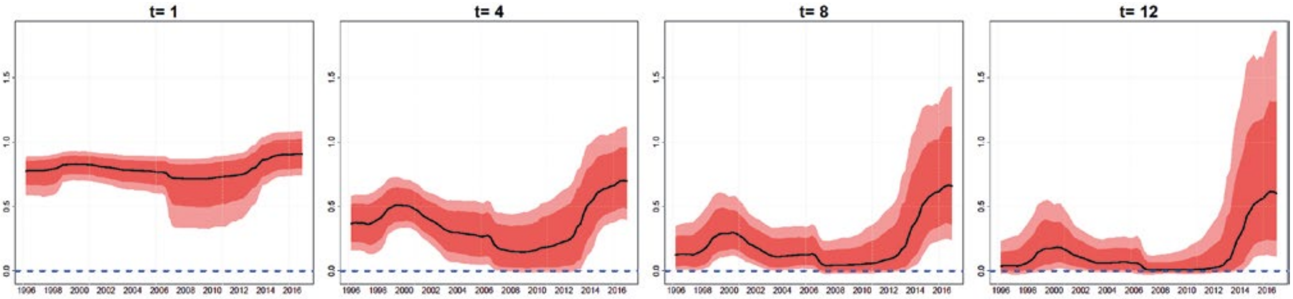
Bulgaria

Impulse responses after one quarter, one, two and three years



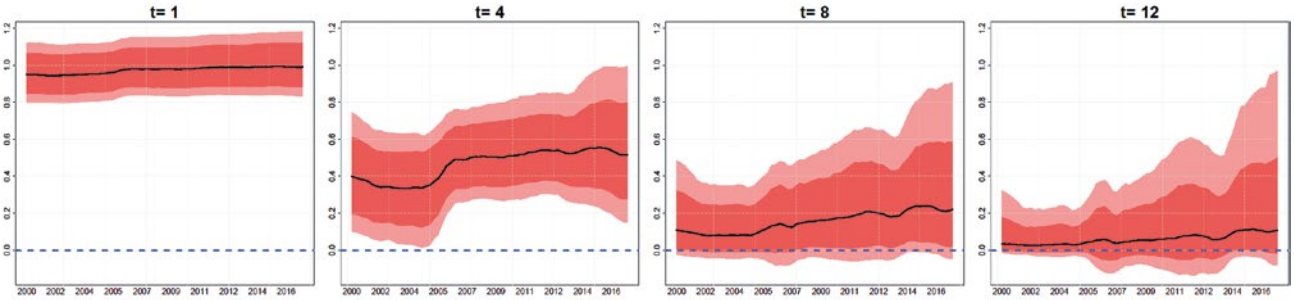
Czech Republic

Impulse responses after one quarter, one, two and three years



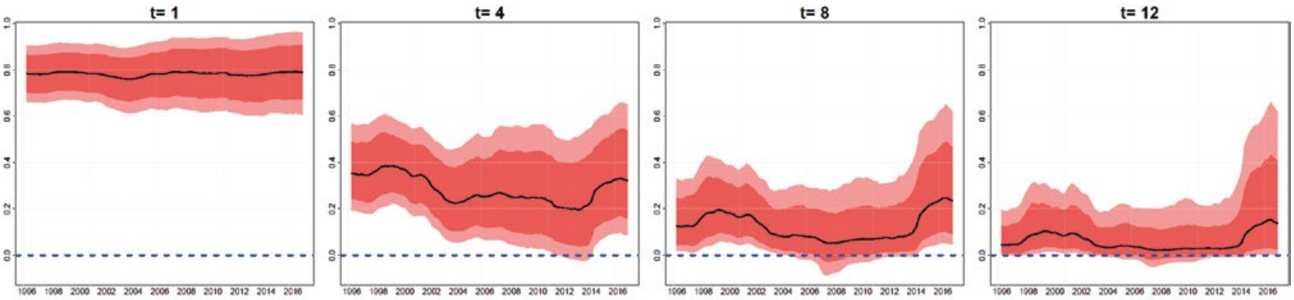
Croatia

Impulse responses after one quarter, one, two and three years



Hungary

Impulse responses after one quarter, one, two and three years



Source: Authors' calculations.

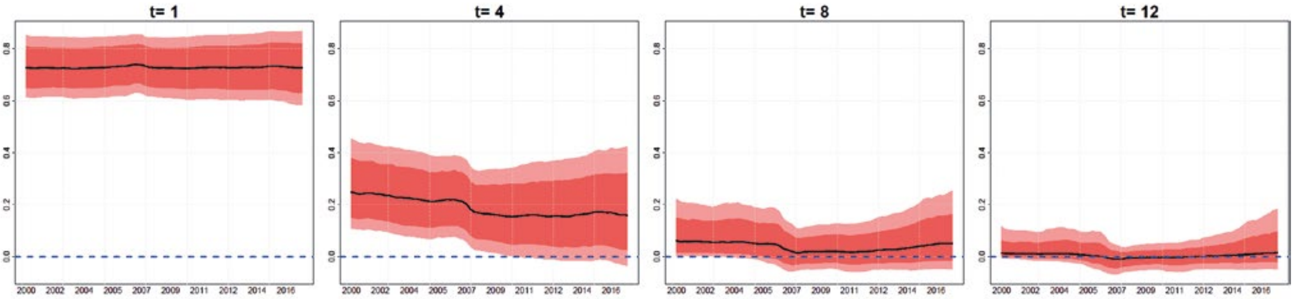
Note: Black lines represent the median response of inflation, whereas dark- and light-red shaded areas represent the 68% and 90% confidence bands.

Chart A1.2

How does a 1 percentage point increase in ULC affect ULC over time?

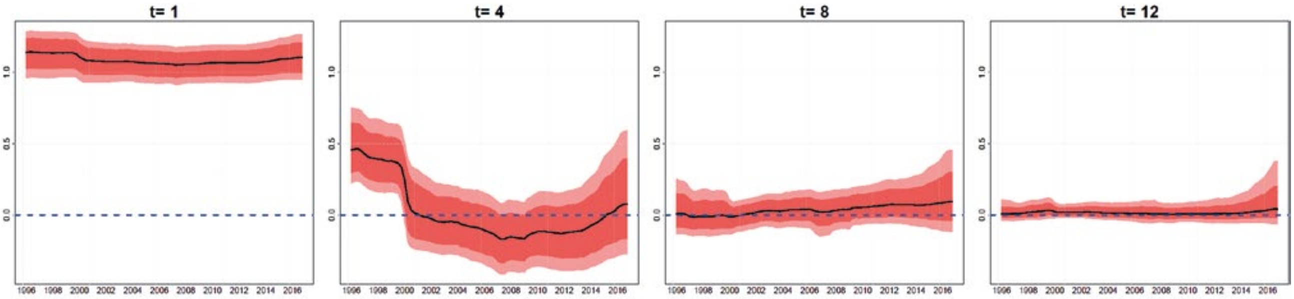
Poland

Impulse responses after one quarter, one, two and three years



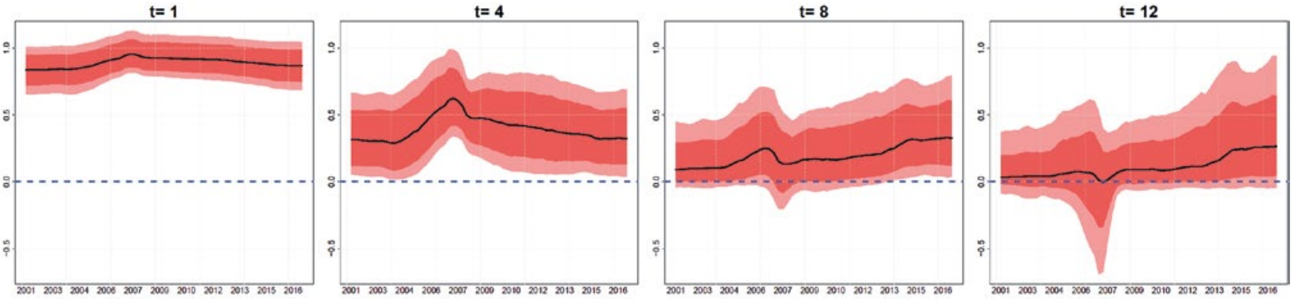
Romania

Impulse responses after one quarter, one, two and three years



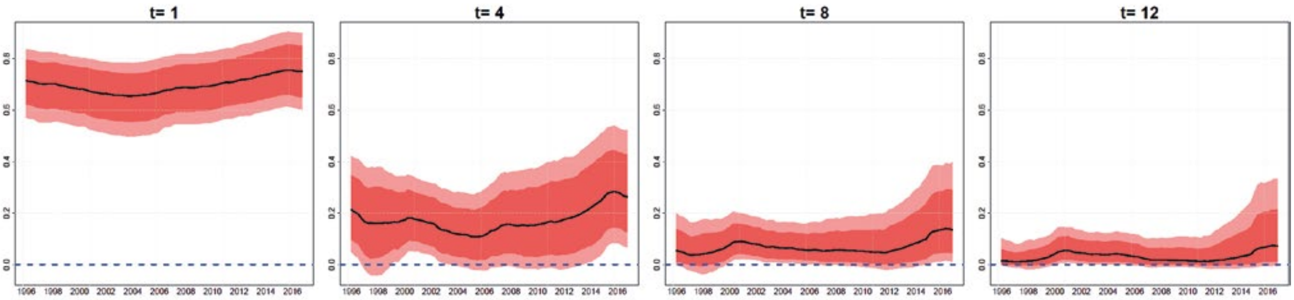
Slovenia

Impulse responses after one quarter, one, two and three years



Slovakia

Impulse responses after one quarter, one, two and three years



Source: Authors' calculations.

Note: Black lines represent the median response of inflation, whereas dark- and light-red shaded areas represent the 68% and 90% confidence bands.