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Asymmetries in effects of domestic inflation drivers in the Baltic States: a Phillips curve-based nonlinear ARDL approach

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ABSTRACT
This study investigates asymmetry in the impact of domestic inflation drivers in the Baltic States, focusing on the output gap and unemployment gap. We aim to reveal how positive and negative changes in these economic activity indicators affect the inflation rate by employing a nonlinear autoregressive distributed lag approach (NARDL) and the Phillips curve. Empirical results demonstrate the long-run asymmetry as inflation in Estonia and Lithuania responds more significantly to positive changes in the output gap, whereas negative changes in the unemployment gap exhibit a stronger long-run impact on inflation in all three countries. These findings mainly suggest some extent of downward price rigidity in the Baltic economies, indicating a nonlinear Phillips curve and relatively large costs of disinflation policy directed to aggregate demand reduction. Further analysis reveals that increasing downward nominal wage flexibility could reduce these asymmetries in Estonia and, to a lesser extent, in Latvia and Lithuania.

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Inflation; output gap; unemployment gap; nonlinear ARDL model; the Baltic States

1. Introduction

Inflation rate dynamics in an economy is determined by numerous factors. Some of them are external, such as global economic disorders, whereas others relate to domestic inflation drivers. The impact of external factors on the inflation rate in recent years has often been the subject of research since the process of globalization brings a greater connection between national economies (e.g. Auer et al., 2017; Ciccarelli & Mojon, 2010; Jordan, 2016; Nagy & Tengely, 2018; Živkov et al., 2019). In these studies, oil prices, changes in the exchange rate or international economic disorders were used as explanatory variables. Domestic inflation drivers are generally linked to the pressures of aggregate demand and the dynamics of aggregate supply in the observed country. Although in contemporary conditions almost all economies are subject to external influences that reflect on the inflation rate dynamics, the growing body of literature confirms that domestic inflation drivers have a more significant impact on the inflation rate than the external ones, even in explaining the 'missing deflation puzzle' which has been observed in

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most economies during and after the Great Recession (for instance, Bobeica & Jarociński, 2019; Globan et al., 2016; Hałka & Kotłowski, 2017; Lanne & Luoto, 2013). The Phillips curve-based analysis in the European Central Bank's Economic Bulletin of 2017 (European Central Bank, 2017a) showed that global inflation drivers had a statistically significant impact on inflation in the Eurozone only between 2008 and 2009, whereas in the period 2012–2015 domestic factors were dominant. Also, Abdih et al. (2018) in a recent study demonstrated that 'domestic factors dominate global factors in explaining recent inflation dynamics' in the euro area. They also concluded that 'the domestic Phillips curve can explain the "deflation puzzle", with inflation persistence being the key factor behind recent low inflation' (Abdih et al., 2018, p. 6).

Among domestic inflation drivers, the output gap, as the ratio between actual and potential gross domestic product, is frequently used as a measure of aggregate demand pressures on inflation. Large body of literature confirmed that dynamics of the output gap could explain the movement of the inflation rate relatively good (Başer Andiç et al., 2015; Bjørnland et al., 2008; Claus, 2000; Lawless & Whelan, 2011; Mehra, 2004; Neiss & Nelson, 2005; Saman & Pauna, 2013). An alternative measure commonly used in empirical research relates to the unemployment gap, as a deviation of unemployment from its natural rate, i.e. NAIRU rate (Non-Accelerating-Inflation Rate of Unemployment). The positive unemployment gap, for instance, is a result of excess labour supply on the labour market, making downward pressure on wages and inflation rate.

Domestic inflation drivers on the supply side are generally related to the changes in both productivity and the production inputs prices, leading to cost-push inflation. A usual approach to explaining inflation dynamics assumes that prices are determined as a mark-up on a firm's production costs (Galí, 2000). Accordingly, changes in the marginal cost ultimately causing changes in product prices. Numerous studies confirm this statement, most often using unit labour costs as an approximation for real marginal costs (for example, Alexová, 2012; Furuoka, 2016; Galí, 2000; Tatierská, 2010). Yet, there are also studies that dispute the importance of the labour share and unit labour costs for explaining the dynamics of inflation, arguing these are not a credible measure of the real marginal costs (Karabarbounis & Neiman, 2013; King & Watson, 2012; Peneva & Rudd, 2017). Put it simply, labour's share in income is generally countercyclical, whereas there is empirical evidence that marginal cost is procyclical (Ball & Mazumder, 2011).

The impact of domestic inflation drivers in empirical studies was the most commonly investigated assuming the symmetry (linearity) of their relationship with inflation, i.e. positive and negative changes in their values (serving as regressors) have symmetrical effects on the inflation rate (as dependent variable) (e.g. Ball & Mazumder, 2019; Dotsey et al., 2018; Hałka & Kotłowski, 2017; Jašová et al., 2018; Tiwari et al., 2014; Zhang & Murasawa, 2011). However, the revealing of asymmetry in this relationship can provide useful information for economic policymaking, since policy measures with different timing and intensity should be applied in the case of positive and negative changes in these inflation drivers.

The aim of this study is, thus, to investigate the presence of asymmetries in the effects of the output gap and unemployment gap on the inflation rate analysing the data for the Baltic States (Estonia, Latvia, and Lithuania). These countries have been chosen for analysis due to several facts. First of all, in recent years, the inflation rate in all Baltic countries, especially in Latvia and Lithuania, was often above the average of the Eurozone (European

Bank for Reconstruction and Development, 2018), raising the question about the relative impact of domestic (country-specific) inflation drivers in these countries, apart from the impact of common factors that all other EMU members are exposed. Since these being among the fastest-growing economies in the Eurozone, the inflation increases might be explained, to a certain extent, by the real output exceeding the potential one. Indeed, there is evidence that the Baltic economies are operating above the potential in recent years (Swedbank, 2018). Besides that, there was a significant growth of unit labour costs in the Baltic economies over the past few years (European Central Bank, 2017b), as a result of a deceleration in productivity growth and an acceleration in the real compensation per worker growth. This opens a question about the relative impact of nominal and real wage dynamics on inflation in the Baltic countries. Since being the economies with relatively institutionally-flexible labour market (low union density rate, low employment protection legislation strictness, flexible labour policy), as documented by, for instance, Paas et al. (2002) and Eamets and Masso (2005), it is interesting to investigate whether the wage adjustment process follows that flexibility. The presence of some kind of wage rigidity can explain the price inflation dynamics and the asymmetry in the economic activity-inflation nexus. Finally, this study could fill the gap in the literature dealing with the issue of the domestic drivers' impact on the inflation rate dynamics in the Baltic States, which was mainly based on the assumption that the impact is symmetric (for instance, Dabušinskas & Kulikov, 2007; Masso & Staehr, 2005; Meļihovs & Zasova, 2009; Reigl, 2017; Vanags & Hansen, 2008; Virbickas, 2012).

In order to analyse the above-mentioned, we applied a nonlinear autoregressive distributed lag approach (NARDL), recently proposed by Shin et al. (2014). As suggested by these authors, this approach has important advantages over the existing estimation techniques, such as Error Correction Model (ECM), the Markov-switching ECM (Psaradakis et al., 2004) and the Smooth transition ECM (Kapetanios et al., 2006), since it can model jointly the cointegration dynamics and long- and short-run asymmetries in the context of an unrestricted error correction model. In this paper, the NARDL model specification is derived from the New Keynesian Phillips curve relationship, which includes the expected inflation rate and the output gap or, alternatively, the unemployment gap as determinants of the actual inflation rate. The potential presence of asymmetries in the relation between these variables indicates nonlinearity of the Phillips curve, which can be concave, convex or bent over, as suggested by some of the recent studies (e.g. Bildirici & Özaksoy, 2016; Bildirici & Sonustun Özaksoy, 2018; Kobbi & Gabsi, 2017). The nonlinearity of the Phillips curve has important implications for economic policy modelling and achieving a goal of price stability.

This paper contributes to the existing empirical literature at least in two aspects. First, the analysis in the paper is focused on the nonlinear nature of changes in domestic inflation drivers in the Baltic States. Accordingly, it reveals whether positive changes in their value have a different impact on the inflation rate than the negative ones, giving an insight into the additional characteristics of the inflation process in these countries. For instance, if the inflation rate responds more significantly to positive than the negative changes in the output gap, it can be a sign of a certain degree of downward price rigidity, which could increase the costs of disinflationary economic policy. Second, we employ nonlinear ARDL modelling in the estimation of the wage Phillips curve in these economies in order to reveal the role of nominal and real wage rigidities as the potential causes for observed asymmetry in the relationship between inflation and economic activity. The analysis shows that in the relatively flexible Baltic labour markets there is downward wage stickiness which might prevent the complete adjustment of prices to economic disturbances. Although this is not the first paper which deals with the problem of macro-level wage rigidity in the Baltic States (see, for example, Babetskii, 2007; Radziwiłł & Walewski, 2003; Von Hagen & Traistary, 2005), this study adds further to this issue by using relatively novel econometric methodology and the more recent dataset.

The rest of the paper is structured as follows. The second section gives an overview of the relevant empirical literature. The third section explains the econometric methodology and the fourth section contains a data set. The fifth section gives a presentation of empirical results and discussion, whereas the sixth analyses possible causes of asymmetry. The last section concludes.

2. Review of the empirical literature

Empirical verification of the importance of domestic inflation drivers is a part of the growing body of literature published in recent years. In these studies, some variants of the Phillips curve were most commonly used. For instance, Zhang and Murasawa (2011) extend the standard New Keynesian Phillips curve by introduction of the multivariate model-based output gap. They show that this new measure of the output gap is a valid driving force for inflation in China, which is not the case with traditional output gap measures. Tiwari et al. (2014) investigate the relation between inflation and the output gap in France, by application of a wavelet transform approach. Their findings indicate that the output gap leads inflation in the short and medium run and that is able to predict the inflation dynamics. Chin (2019) analyses the New Keynesian Phillips curve in the United States with parameters of a trade-off between inflation and real economic activity which are allowed to vary across time. Applying the General Method of Moments (GMM), he demonstrates that the estimated parameter of output-inflation trade-off is time-varying and is larger in periods with high inflation. The extent of price rigidity is one of the most important factors for the explanation of the parameter time-variation. Jašová et al. (2018) investigate the effects of domestic and global output gaps on inflation, measured by the Consumer Price Index, for a panel of 26 advanced and 22 emerging economies. They show that both kinds of output gaps are significant drivers of inflation both in pre-crisis and post-crisis periods.

The impact of the real marginal costs (most commonly approximated by real unit labour costs) on the inflation rate, based on the assumption that prices are determined as a mark-up on firms' costs, is often analysed in empirical research. Notwithstanding, their significance for an explanation of the inflation dynamics is quite controversial. For example, Tatierská (2010) shows the importance of the unit labour costs as determinants of price level dynamics in eight of the eleven euro area countries. Alexová (2012) find that inflation in the half of ten observed new EU members from Central and Eastern Europe was caused by cost-push factors and in the rest by demand-side factors. On the other hand, King and Watson (2012) show that real unit labour costs have no potential to explain the inflation dynamics in the United States accurately and that rather real factors have an impact on the labour's share in a manner largely unrelated to inflation. Peneva and Rudd (2017) come to similar conclusions for the same country. Karabarbounis and

Neiman (2013) stress the declining tendency of the labour share and unit labour costs in a number of countries since the 1980s, concluding that these could not be the appropriate proxy for the real marginal costs. Razgūnė and Lazutka (2015) confirm that tendency in the case of the Baltic States.

In recent studies, the nonlinear character of the relationship between the inflation rate and various macroeconomic variables is emphasized. Thus, Bildirici and Özaksoy (2016) analyse nonlinearities in the Post Keynesian Phillips curve in Canada, in order to investigate the degree of labour market flexibility. They reveal a bidirectional causality relationship between inflation, unemployment, and economic growth, indicating that the labour market is flexible. They also concluded that there are asymmetries in the long-run relationship between those variables. Lepetit (2018) investigate the role of labour market asymmetries in setting the optimal monetary policy, by analysing unemployment fluctuations in a New Keynesian model with search and matching frictions. He conclude that these asymmetries are crucial in generating a significant trade-off between inflation and unemployment and that monetary policy should respond to both inflation and unemployment. Using the nonlinear ARDL approach combined with causality methods, Bildirici and Sonustun Özaksoy (2018) analyse the relationship between inflation and unemployment in Japan, Turkey, the USA, and France. They try to investigate whether the structure of the Phillips curve in these countries is backward bending, which might indicate that the relationship between inflation and unemployment is positive. They conclude that there is a rather negative long-run relationship between the variables, which is also asymmetric. Kobbi and Gabsi (2017) check the nonlinearity of the hybrid New Keynesian Phillips curve in Tunisia, by application of the Logistic Smooth Transition Regression (LSTR) model. They find evidence that inflation responds significantly to the output gap only in the case when the inflation rate exceeds a certain threshold. The price rigidity dominates when the inflation rate is relatively low.

Although the specificity of the inflation drivers in former transitional economies has been frequently investigated (e.g. Basarac et al., 2011; Bouda, 2013; Danišková & Fidrmuc, 2011; Vasilev, 2015; Vašíček, 2011), there are relatively few macroeconomic studies concerning inflation rate dynamics in the Baltic countries. Thus, Masso and Staehr (2005) use single-equation and panel GMM estimation for the Phillips curve in the Baltic States in order to investigate the importance of international prices adjustment, exchange rate, labour market tendencies, and the output gap. They conclude, *inter alia*, that excess capacity in the labour market have no effect on inflation, whereas changes in the output gap can explain some extent of inflation dynamics. Meļihovs and Zasova (2007) find that the output gap, foreign price shocks, and expected future inflation rate have a significant impact on the core inflation rate in Latvia. Dabušinskas and Kulikov (2007) estimate the New Keynesian Phillips curve for Estonia, Latvia, and Lithuania using the GMM estimation. One of their main conclusions is that the inflation process depends mainly on inflation expectations and lagged inflation, whereas the role of real factors of inflation dynamics, such as marginal costs, is rather limited. These authors also find out some extent of price rigidity in observed countries, which have an influence on the inflation rate dynamics. Similar findings are documented in Virbickas (2012) for Lithuania. Furuoka (2016) analyse the validity of the New Keynesian Phillips curve in the Baltic States. According to that study, the inflation dynamics is mainly determined by forward-looking inflation expectations, whereas marginal costs do not expose a significant impact on the inflation rate. Reigl (2017) show how the headline and core inflation rate in Estonia can be forecasted by application of the factor models. One of his key conclusions is that the quality of the forecasting process depends on the relationship between a number of factors and the size of the dataset.

3. Econometric model

3.1. Specification of the Phillips curve relationship

The Phillips curve generally captures the relationship between the inflation rate and some measure of real economic activity. In the New Keynesian Phillips curve, that measure is presented by dynamics of marginal cost and is usually represented by the equation (Galí, 2000):

$$\pi_t = \beta E_t(\pi_{t+1}) + \lambda mc_t, \tag{1}$$

where π_t and $E_t(\pi_{t+1})$ represent current and expected future inflation rate, respectively, mc_t denotes the marginal costs, whereas λ indicates the frequency of price changes. Due to the lack of data about real marginal costs dynamics, most empirical studies employ proxies like labour share (or unit labour costs), output gap, and unemployment gap. Although advocated by numerous researchers (e.g. Chin, 2019; Furuoka, 2016; Galí et al., 2001; Tatierská, 2010), the application of labour share and unit labour costs as a proxy for marginal costs became increasingly problematic (Coibon & Gorodnichenko, 2015; Karabarbounis & Neiman, 2013; King & Watson, 2012; Lindé, 2005; Rudd & Whelan, 2005). Accordingly, we employ the Phillips curve specification with an output gap and, alternatively, the unemployment gap, as measures of economic activity.

Starting for Equation (1) we assume that the following relation is valid, as suggested in Galí (2000):

$$mc = \kappa(y_t - y_t^*), \tag{2}$$

from which it follows:

$$\pi_t = \beta E_t(\pi_{t+1}) + \lambda\kappa(y_t - y_t^*) \tag{3}$$

where y_t and y_t^* refer to actual and potential output, respectively. Bearing in mind the appropriate data about the expected inflation rate are not available for the Baltic States, we use the past inflation rate as a regressor, which is advocated by numerous empirical studies (e.g. Ball & Mazumder, 2011; Coibon & Gorodnichenko, 2015; Rudd & Whelan, 2005). This approach also coincides with the studies which have cast doubt on the importance of rationality and 'forward-looking' behaviour in the process of forming inflationary expectations (e.g. Batini et al., 2005; Fuhrer, 1997; Guay & Pelgrin, 2004). Hence, Equation (3) gets the following form:

$$\pi_t = \pi_{t-1} + \gamma x_t + \varepsilon_t, \tag{4}$$

where π_{t-1} denotes the inflation rate in the previous period, γ is the coefficient that measures the impact of the change in the output gap on the inflation rate, x stands for the output gap ($x_t \equiv y_t - y_t^*$) whereas ε_t denotes an error term.

In this paper, we also introduce the unemployment gap into the Phillips curve relation, as a measure of capacity utilization in the labour market, which is other common approach in economic literature (Ball & Mazumder, 2011; Ball & Mazumder, 2019; Blanchard, 2016;

Dotsey et al., 2018; Matheson & Stavrev, 2013; Rondina, 2018; Svensson, 2015). The Phillips curve is formulated as follows:

$$\pi_t = \pi_{t-1} + \nu\psi_t + \varepsilon_t, \tag{5}$$

where ψ stands for unemployment gap ($\psi_t \equiv u_t - u_t^*$), ν is the coefficient that measures the impact of the change in the unemployment gap on the inflation rate, whereas u_t and u_t^* denote actual unemployment rate and NAIRU rate, respectively. The impact of the output gap and unemployment gap on inflation in Equations (4) and (5) is expected to be opposite due to the negative relationship between output and unemployment captured by Okun's law.

3.2. Linear and nonlinear ARDL model specifications

Starting from Equation (4) we formulate the following linear (symmetric) ARDL (p, q) model in the error correction form, as suggested by Pesaran et al. (2001):

$$\Delta\pi_t = \alpha_1 + \beta_1\pi_{t-1} + \gamma x_{t-1} + \sum_{i=1}^{p}\kappa_i\Delta\pi_{t-i} + \sum_{i=0}^{q}\mu_i\Delta x_{t-i} + \varepsilon_t, \tag{6}$$

where α_1 represents the constant, β_1 and γ denote long-run coefficients, κ_i and μ_i are the short-run coefficients, and p and q denote lag length. Following Shin et al. (2014) we decompose the movements in the output gap (x_t) into its increasing and decreasing partial sum, i.e. $x_t = x_0 + x_t^+ + x_t^-$, where x_t^+ and x_t^- are partial sum processes of positive and negative changes in the output gap, respectively, which are generated as follows:

$$x_t^+ = \sum_{i=0}^{t}\Delta x_i^+ = \sum_{i=0}^{t} max(\Delta x_i, 0); \tag{7}$$

$$x_t^- = \sum_{i=0}^{t}\Delta x_i^- = \sum_{i=0}^{t} min(\Delta x_i, 0). \tag{8}$$

By replacing the positive and negative partial sum of x_t in the linear ARDL model (6) we get the partial asymmetry cointegration equation or NARDL model:

$$\Delta\pi_t = \alpha_1 + \beta_1\pi_{t-1} + \gamma^+x_{t-1}^+ + \gamma^-x_{t-1}^- + \sum_{i=1}^{p}\kappa_i\Delta\pi_{t-i} + \sum_{i=0}^{q}(\mu_i^+\Delta x_{t-i}^+ + \mu_i^-\Delta x_{t-i}^-) + \varepsilon_t. \tag{9}$$

Likewise, we reformulate the Equation (5) in order to get the linear ARDL (m, n) model with the unemployment gap (ψ_t) as an explanatory variable:

$$\Delta\pi_t = \alpha_2 + \beta_2\pi_{t-1} + \nu\psi_{t-1} + \sum_{j=1}^{m}\vartheta_j\Delta\pi_{t-j} + \sum_{j=0}^{n}\tau_j\Delta\psi_{t-j} + \varepsilon_t, \tag{10}$$

where α_2 denotes the constant, β_2 and ν are the long-run coefficients, ϑ_j and τ_j are the short-run coefficients whereas m and n represent the lag length. The partial sum of positive (ψ_t^+) and negative (ψ_t^-) changes in the unemployment gap, respectively, are generated as follows:

$$\psi_t^+ = \sum_{j=0}^{t} \Delta\psi_j^+ = \sum_{j=0}^{t} max(\Delta\psi_j, 0) \tag{11}$$

$$\psi_t^- = \sum_{j=0}^{t} \Delta\psi_i^- = \sum_{j=0}^{t} min(\Delta\psi_j, 0). \tag{12}$$

Finally, by introducing the positive and negative partial sum of ψ_t in the Equation (10) we get the NARDL model representing the effects of the changes in the unemployment gap on the inflation rate:

$$\Delta\pi_t = \alpha_2 + \beta_2\pi_{t-1} + \nu^+\psi_{t-1}^+ + \nu^-\psi_{t-1}^- + \sum_{j=1}^{m} \vartheta_j\Delta\pi_{t-j} + \sum_{j=0}^{n} (\tau_j^+\Delta\psi_{t-j}^+ + \tau_j^-\Delta\psi_{t-j}^-) + \varepsilon_t. \tag{13}$$

The presence of cointegration, as well as the long-run and short-run asymmetries between the variables, is tested using the standard Wald test, which is a common approach in empirical research (e.g. Bildirici & Özaksoy, 2016; Kobbi & Gabsi, 2017; Shin et al., 2014; Tang & Bethencourt, 2017). All null hypothesis formulations are given in section 5, along with estimation results.

4. Data

For this study, we use quarterly data about the real gross domestic product (in USD), the unemployment rate (the share of unemployed persons in the total labour force) and Harmonized Consumer Price Index (CPI), as a measure of the inflation rate, for Estonia, Latvia, and Lithuania. The data are collected from the OECD database. All data are seasonally adjusted and converted into logarithmic form. The value of the output gap was calculated as a log ratio of actual to potential real output, which is obtained using Hodrick-Prescott (HP) filter (Hodrick & Prescott, 1981), as suggested by numerous researchers (Jašová et al., 2018; Mehra, 2004; Neiss & Nelson, 2005; Tiwari et al., 2014). The smoothness parameter λ in the HP filter takes the value of 1600, which is a common approach when it is applied to quarterly data (Flashel et al., 2008). Analogously, the unemployment gap was calculated as a log ratio of the actual unemployment rate to the NAIRU rate obtained *via* the HP filter. The analysis covers a time span from the first quarter of 1998 to the fourth quarter of 2018 (84 observations). The earlier period is not included in the analysis to abstract the impact of intensive transitional reforms on inflationary processes in the Baltic States. At the same time, the observed period includes events that significantly influenced the inflation dynamics in these countries, such as the Russian financial crisis in 1998, the accession of Baltic countries to the European Union in 2004, and the Great Recession of 2008.

Figure 1 presents the empirical dynamics of CPI inflation, output gap and unemployment gap in the Baltic States. There is a relatively similar pattern in the movement of these variables among countries. Intuitively, one can observe the negative trade-off between the output gap and unemployment gap, as suggested by Okun's law. Accordingly, we expect these drivers to affect inflation with an opposite sign. Bearing in mind the evident impact of the Great Recession on the time series, we introduced a dummy variable (D) into the analysis in order to capture the effects of the crisis. Following Tang and

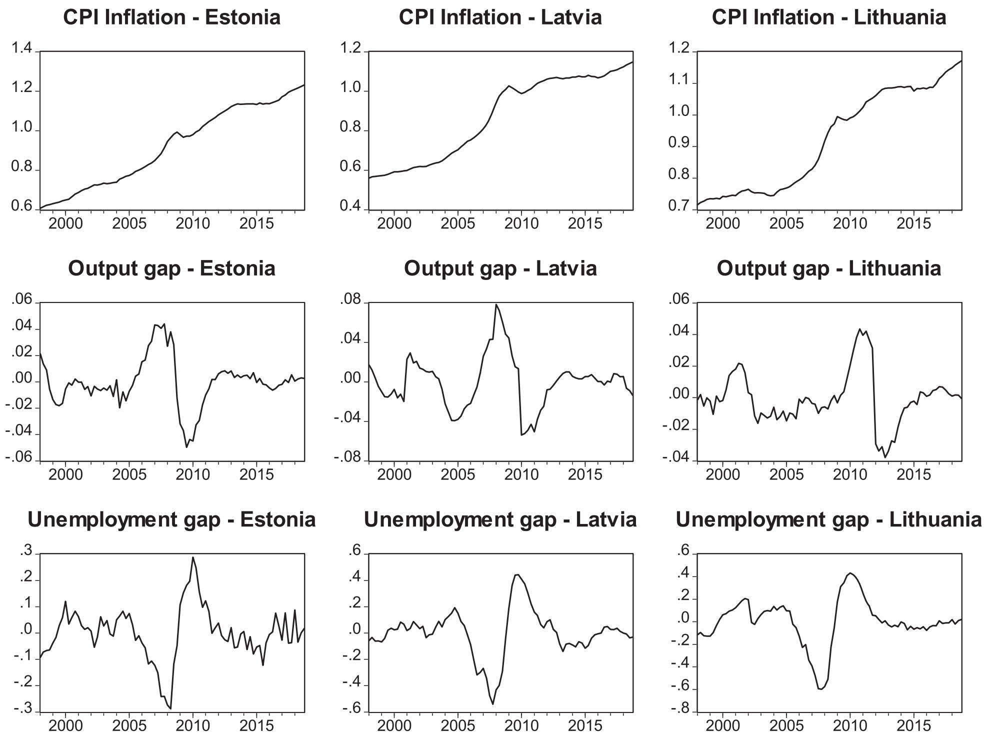

Figure 1. Empirical dynamics of CPI inflation, output gap and unemployment gap in the Baltic States.

Bethencourt (2017), we chose the first quarter of 2008 as a date when the crisis started. For every period equal to or greater than 2008Q1, *D* takes the value 1, otherwise, *D* equals 0. In the case when the coefficient of the dummy variable is statistically significant, it may be a sign that the structural change occurred in the observed country due to the Great Recession.

5. Empirical results and discussion

In the ARDL and NARDL models, the regressors should not be integrated of the same order, unlike the error correction models (Pesaran et al., 2001; Katrakilidis & Trachanas, 2012). All variables should be integrated of order I(0) or I(1) or mutually integrated, but none of them should be I(2) in order to calculate valid *F*-statistic (Shin et al., 2014). Hence, the first step in the empirical analysis is unit root testing. Table 1 reports the results of the Augmented Dickey-Fuller (ADF) test (Dickey & Fuller, 1979) and the KPSS test (Kwiatkowski et al., 1992). It is evident that these tests give mixed results, but generally it can be concluded that all variables are stationary in level and/or in the first difference.

Since the trend of all observed variables in the Baltic States, as can be seen from Figure 1, has been dramatically changed due to the Great Recession, it can be a sign that there is a structural break in time series. ADF and KPSS tests provide no information about structural breaks, so there is a possibility they offer biased results, as suggested by Baum (2004). Hence, we employed the Zivot-Andrews test (Zivot & Andrews, 1992), in order to get more robust results about the stationarity of the observed time series (Table 2). The

Table 1. The results of the ADF and KPSS test.

Variable	ADF Constant	ADF Constant & Trend	KPSS Constant	KPSS Constant & Trend
Estonia				
π	−0.380 (1)	−2.284 (1)	1.149[a] [7]	0.143[c] [6]
x	−4.262[a] (3)	−4.225[a] (3)	0.048 [6]	0.047 [6]
ψ	−3.924[a] (3)	−3.913[b] (3)	0.046 [6]	0.045 [6]
$\Delta\pi$	−4.395[a] (0)	−4.368[a] (0)	0.088 [5]	0.087 [5]
Δx	−4.008[a] (1)	−3.976[b] (1)	0.052 [5]	0.041 [5]
$\Delta\psi$	−9.573[a] (0)	−9.517[a] (0)	0.043 [5]	0.039 [5]
Latvia				
π	−0.819 (1)	−2.182 (1)	1.101[a] [7]	0.174[c] [7]
x	−4.597[a] (5)	−4.575[a] (5)	0.048 [4]	0.047 [4]
ψ	−3.492[b] (1)	−3.469[b] (1)	0.043 [6]	0.046 [6]
$\Delta\pi$	−2.917[b] (0)	−2.910 (0)	0.152 [6]	0.134[c] [6]
Δx	−3.111[b] (3)	−3.070 (3)	0.068 [13]	0.066 [13]
$\Delta\psi$	−4.394[a] (0)	−4.365[a] (0)	0.038 [5]	0.036 [5]
Lithuania				
π	0.105 (1)	−2.481 (2)	1.113[a] [7]	0.133[c] [7]
x	−4.102[a] (3)	−4.076[b] (3)	0.047 [6]	0.047 [6]
ψ	−4.099[a] (2)	−4.075[b] (2)	0.053 [6]	0.055 [6]
$\Delta\pi$	−4.403[a] (0)	−4.435[a] (0)	0.150 [6]	0.122[c] [6]
Δx	−7.892[a] (0)	−7.843[a] (0)	0.032 [4]	0.032 [4]
$\Delta\psi$	−4.269[a] (3)	−4.245[a] (3)	0.042 [6]	0.040 [6]

Note: The significance levels: [a] −0,01; [b] −0,05; [c] −0,1. ADF tests the null hypothesis of unit root in time series, whereas KPSS tests the null of stationarity. For ADF test, the number in parenthesis indicates the lag order selected based on the Akaike information criterion (Akaike, 1981). The numbers in brackets (for KPSS test) indicate the truncation for the Bartlett Kernel, as suggested by the Newey-West test (Newey & West, 1987).

results confirm that the condition for implementation of the ARDL approach is fulfilled. The identified time break dates are ranged between 2007Q2 and 2013Q2, which is due to the impact of Great Recession.

In order to investigate the presence of a long-run relationship between variables in models (6), (9), (10) and (13), we applied the bounds-testing procedure advanced by Pesaran et al. (2001), and Shin et al. (2014). The optimal lag structure was chosen based on the Akaike Information Criterion (Akaike, 1981) and the Schwarz Information Criterion (Schwarz, 1978). Following Katrakilidis and Trachanas (2012) and Tang and Bethencourt (2017), we also used the general-to-specific procedure as the final model specification technique, by successively trimming all insignificant lags (starting from 4 lags). Table 3 reports the results of the bounds test to cointegration. It can be seen that the cointegration was not revealed in any linear (symmetric) ARDL specification, suggesting that the relationship between domestic inflation drivers and the inflation rate is rather nonlinear. Indeed, when one observes the nonlinear ARDL specifications, the cointegration

Table 2. Zivot-Andrews breakpoint unit root test.

Variable	Estonia t-statistic	Estonia Break date	Latvia t-statistic	Latvia Break date	Lithuania t-statistic	Lithuania Break date
π	−4.157[a]	2009Q3	−5.341[a]	2007Q3	−4.472[a]	2007Q3
x	−6.689[a]	2008Q4	−4.546[a]	2010Q1	−5.544[a]	2012Q1
ψ	−5.402[a]	2008Q3	−5.541[a]	2008Q3	−4.965[a]	2008Q3
$\Delta\pi$	−5.268[b]	2013Q2	−5.029[a]	2009Q2	−4.323[a]	2009Q2
Δx	−4.417[b]	2007Q2	−3.539[b]	2008Q3	−8.325[b]	2011Q3
$\Delta\psi$	−4.346[b]	2008Q3	−5.013[b]	2008Q1	−5.261[b]	2007Q4

Note: The significance levels: [a] −0,01; [b] −0,05; [c] −0,1. Null hypothesis: The series has a unit root with a structural break in both the intercept and trend.

between output gap and inflation is present in all countries, whereas the cointegration between the unemployment gap and inflation exists in the case of Estonia and Lithuania. In addition, the robustness of obtained results is confirmed by comparison with critical values provided by Narayan (2005), which are more suitable for sample sizes between 30 and 80 observations. These critical values for the sample size of 80 observations and the case of unrestricted intercept and no trend with 5% level of significance are: I(0) = 5.060, I(1) = 5.930 (for symmetric ARDL) and I(0) = 3.940, I(1) = 5.043 (for asymmetric ARDL).

Bearing in mind that the presence of cointegration provides no information about the direction of causality between the variables, the application of causality tests is required. When the time series are integrated of different orders, as in our case, the standard *F*-statistic for testing the Granger causality (Granger, 1969) may be misleading since the test does not have a standard distribution (Toda & Yamamoto, 1995). Hence, we used the Toda-Yamamoto approach to Granger causality test (Toda & Yamamoto, 1995) in the form of a standard VAR model based on a modified Wald test. This approach tests for the causality of time series in levels, thus reducing the risk of wrong identification of time series' order of integration (Wolde-Rufael, 2005).

In order to apply the Toda-Yamamoto approach, we formulate the relation between time series for inflation, output gap and unemployment gap in the form of the VAR model. For instance, the causality between inflation and output gap (symmetric case) was investigated using the following system:

$$\pi_t = a_0 + \sum_{i=0}^{k} a_{1i}\pi_{t-i} + \sum_{j=k+1}^{dmax} a_{2j}\pi_{t-j} + \sum_{i=0}^{k} b_{1i}x_{t-i} + \sum_{j=k+1}^{dmax} b_{2j}x_{t-j} + e_{1t} \tag{14}$$

$$x_t = c_0 + \sum_{i=0}^{k} c_{1i}x_{t-i} + \sum_{j=k+1}^{dmax} c_{2j}x_{t-j} + \sum_{i=0}^{k} d_{1i}\pi_{t-i} + \sum_{j=k+1}^{dmax} d_{2j}\pi_{t-j} + e_{2t}, \tag{15}$$

where k denotes optimal lag length, whereas d_{max} refers to the maximal order of integration in observed time series. The optimal lag length was selected based on the Schwarz information criterion (Schwarz, 1978), which is the most accurate for quarterly

Table 3. Bounds test to cointegration in symmetric and asymmetric ARDL model.

	Estonia		Latvia		Lithuania	
	Symmetric ARDL (3,4)	Asymmetric ARDL	Symmetric ARDL (2,1)	Asymmetric ARDL	Symmetric ARDL (3,0)	Asymmetric ARDL
Explanatory variable: output gap						
F-statistic	1.323	6.815	4.255	5.810	0.313	6.371
Cointegration	No	Yes	No	Yes	No	Yes
	Symmetric ARDL (2,3)	Asymmetric ARDL	Symmetric ARDL (2,2)	Asymmetric ARDL	Symmetric ARDL (3,2)	Asymmetric ARDL
Explanatory variable: unemployment gap						
F-statistic	0.469	5.151	2.821	4.187	2.129	5.526
Cointegration	No	Yes	No	No	No	Yes

Note: Inflation rate is the dependent variable. The *F*-statistic was calculated using Wald test for the null hypothesis of no cointegration $\beta_1 = \gamma = 0$ (symmetric ARDL) and $\beta_1 = \gamma^+ = \gamma^- = 0$ (asymmetric ARDL) for a model with the output gap, and $\beta_2 = \nu = 0$ (symmetric ARDL) and $\beta_2 = \nu^+ = \nu^- = 0$ (asymmetric ARDL) for a model with unemployment gap. The critical values are obtained from Pesaran et al. (2001), Case III: Unrestricted intercept and no trend, level of significance 5%: I(0) = 4.94, I(1) = 5.73 (for symmetric ARDL), and I(0) = 3.79, I(1) = 4.85 (for asymmetric ARDL).

VAR models and sample sizes smaller than 120 observations, as suggested by Ivanov and Kilian (2005). The maximal order of integration of time series is 1.

Following Wolde-Rufael (2005) we conclude that in Equation (14) there is a Granger causality from x_t to π_t if $b_{1i} \neq 0\ \forall i$; likewise, in Equation (15) there is a Granger causality from π_t to x_t if $d_{1i} \neq 0\ \forall i$. By analogy with relations (14) and (15), we formulated and tested for causality the relation between unemployment gap and inflation. In addition, as suggested by Hatemi-J (2012), we tested for causality between positive and negative changes in output gap (unemployment gap) and inflation, which is important for further analysis of asymmetries. Table 4 reports the results of the Granger non-causality test for symmetric and asymmetric cases. It is evident that the causality between the output gap and inflation in Estonia is bidirectional for all pairs of variables, whereas the results are mixed for Latvia and Lithuania. By using an additional criterion based on the value of the χ^2 statistic, we can conclude that in all Baltic countries the changes in the output gap better explain inflation rate dynamics than *vice versa*, so taking the output gap as an explanatory variable is justified. The causality between the unemployment gap and inflation is unidirectional in most of the cases, with the unemployment gap as an explanatory variable.

The exact specification of the asymmetric ARDL model with the output gap is presented in Table 5. Since the presence of the long-run relationship between inflation and the output gap is already confirmed (Table 3), we first focus on the values and statistical significance of the long-run coefficients, L_x^+ and L_x^-. The significance is confirmed for coefficients of positive changes in the output gap (L_x^+) for Estonia and Lithuania, and negative ones (L_x^-) for Latvia. The estimated values of L_x^+ and L_x^- are 2.70 and 0.38 for Estonia, 0.25 and −2.41 for Latvia and 2.61 and 0.70 for Lithuania, respectively. We then may conclude that a 1% positive change in the output gap leads to a 2.7% increase in the inflation rate in Estonia and 2.61% in Lithuania. In the case of Latvia, the sign of estimated values is not in line with reported literature, since a 1% of negative change in output gap in Latvia results in a 2.41% rise in inflation. The estimated values for negative changes in Estonia and Lithuania, as well as the positive

Table 4. Results of symmetric and asymmetric Granger non-causality tests (Toda-Yamamoto procedure).

Symmetric case		Asymmetric (+) case		Asymmetric (-) case	
H_0	χ^2	H_0	χ^2	H_0	χ^2
Estonia					
$x0\!= \rightarrow 0/\pi$	22.254[a]	$x^+0\!= \rightarrow 0/\pi$	12.050[b]	$x^-0\!= \rightarrow 0/\pi$	23.421[a]
$\pi0\!= \rightarrow 0/x$	7.872[c]	$\pi0\!= \rightarrow 0/x^+$	9.628[c]	$\pi0\!= \rightarrow 0/x^-$	16.091[a]
$\psi0\!= \rightarrow 0/\pi$	6.972[a]	$\psi^+0\!= \rightarrow 0/\pi$	3.863	$\psi^-0\!= \rightarrow 0/\pi$	10.977[c]
$\pi0\!= \rightarrow 0/\psi$	0.002	$\pi0\!= \rightarrow 0/\psi^+$	8.115[b]	$\pi0\!= \rightarrow 0/\psi^-$	2.626
Latvia					
$x0\!= \rightarrow 0/\pi$	34.979[a]	$x^+0\!= \rightarrow 0/\pi$	13.910 [a]	$x^-0\!= \rightarrow 0/\pi$	30.659[a]
$\pi0\!= \rightarrow 0/x$	8.744[c]	$\pi0\!= \rightarrow 0/x^+$	4.802	$\pi0\!= \rightarrow 0/x^-$	11.052[b]
$\psi0\!= \rightarrow 0/\pi$	18.984[a]	$\psi^+0\!= \rightarrow 0/\pi$	13.201[a]	$\psi^-0\!= \rightarrow 0/\pi$	17.036[a]
$\pi0\!= \rightarrow 0/\psi$	0.079	$\pi0\!= \rightarrow 0/\psi^+$	25.270[a]	$\pi0\!= \rightarrow 0/\psi^-$	2.239
Lithuania					
$x0\!= \rightarrow 0/\pi$	28.491[a]	$x^+0\!= \rightarrow 0/\pi$	15.928[a]	$x^-0\!= \rightarrow 0/\pi$	17.862[a]
$\pi0\!= \rightarrow 0/x$	5.666	$\pi0\!= \rightarrow 0/x^+$	3.921	$\pi0\!= \rightarrow 0/x^-$	13.539[a]
$\psi0\!= \rightarrow 0/\pi$	10.967[a]	$\psi^+0\!= \rightarrow 0/\pi$	8.632[b]	$\psi^-0\!= \rightarrow 0/\pi$	6.787[c]
$\pi0\!= \rightarrow 0/\psi$	0.214	$\pi0\!= \rightarrow 0/\psi^+$	26.235[a]	$\pi0\!= \rightarrow 0/\psi^-$	0.498

Note: sign '$0\!= \rightarrow 0/$' means 'does not Granger cause'. The significance levels: [a] –0,01; [b] –0,05; [c] –0,1. x^+, x^-, ψ^+ and ψ^- denote partial sums of positive and negative changes in the output gap and unemployment gap, respectively.

changes in Latvia, are lower than the changes in the opposite direction, indicating the presence of asymmetries, but these results are not statistically significant. Indeed, the long-run asymmetry is confirmed by the Wald test (W_{LR}) in the case of all Baltic countries. On the other hand, there is not enough evidence of the short-run asymmetry, indicating that the impact of the output gap on inflation is rather symmetric in the short-run.

In addition, it must be stressed that the estimated coefficients for past inflation (π_{t-1}) are statistically significant in all the three Baltic States, indicating that the price-setting process is characterized by the backward-looking behaviour in forming inflation expectations. In other words, our specification of the Phillips curve represents well the relation between expected and actual inflation rate in the Baltic countries. The results of residual diagnostic tests, as well as the tests for dynamic stability and functional form, indicate that the model is well specified. In addition, according to the coefficients of a dummy variable, the Great Recession did not lead to a structural change in these economies, except in Latvia. This is certainly surprising having in mind that the Baltic States were among European economies the most severely hit by the crisis.

Table 5. Estimation results for NARDL model with output gap (Equation (9)).

Variable	Estonia	Latvia	Lithuania
α_1	0.069 (0.001)	−0.034 (0.044)	0.046 (0.011)
π_{t-1}	−0.101 (0.002)	0.075 (0.022)	−0.066 (0.014)
x^+_{t-1}	0.272 (0.000)	−0.019 (0.807)	0.172 (0.014)
x^-_{t-1}	0.038 (0.431)	0.181 (0.007)	0.046 (0.356)
$\Delta\pi_{t-1}$	0.429 (0.000)	0.638 (0.000)	0.457 (0.000)
$\Delta\pi_{t-2}$	0.375 (0.002)	-	0.324 (0.015)
$\Delta\pi_{t-3}$	-	−0.203 (0.112)	-
$\Delta\pi_{t-4}$	−0.196 (0.047)	−0.092 (0.463)	0.173 (0.203)
Δx^+_{t-1}	−0.321 (0.055)	0.145 (0.467)	-
Δx^+_{t-2}	-	-	0.252 (0.267)
Δx^+_{t-3}	-	0.405 (0.038)	−0.221 (0.329)
Δx^+_{t-4}	−0.197 (0.239)	0.338 (0.107)	−0.394 (0.097)
Δx^-_{t-1}	0.403 (0.000)	0.278 (0.003)	-
Δx^-_{t-2}	0.081 (0.452)	-	-
Δx^-_{t-3}	−0.334 (0.003)	-	-
Δx^-_{t-4}	-	−0.061 (0.511)	-
D	0.003 (0.441)	−0.011 (0.012)	0.005 (0.264)
L^+_x	2.693 (0.000)	0.253 (0.792)	2.606 (0.016)
L^-_x	0.376 (0.467)	−2.413 (0.014)	0.697 (0.419)
R^2	0.679	0.818	0.501
F-stat.	11.677 (0.000)	24.781 (0.000)	6.589 (0.000)
JB test	0.541 (0.763)	0.137 (0.934)	2.432 (0.296)
BG LM test	0.357 (0.702)	0.456 (0.636)	0.166 (0.848)
ARCH test	0.031 (0.861)	1.125 (0.292)	2.219 (0.076)
Cusum test	Stable	Stable	Stable
Cusum Squared test	Stable	Stable	Stable
Ramsey RESET test	1.676 (0.200)	2.241 (0.139)	2.788 (0.069)
W_{LR}	122.077 (0.000)	56.829 (0.000)	17.503 (0.000)
W_{SR}	3.431 (0.069)	3.606 (0.062)	0.774 (0.382)

Note: value in parenthesis represents the corresponding *p*-value. *D* refers to dummy variable capturing effects of the Great Recession of 2008. L^+_x and L^-_x denote estimated long-run coefficients of positive and negative changes in the output gap, respectively, calculated as $L^+_x = -\hat{\gamma}^+/\hat{\beta}_1$ and $L^-_x = -\hat{\gamma}^-/\hat{\beta}_1$ (Equation (9)). JB, BG LM and ARCH denote Jarque-Bera test for normality, Breusch Godfrey test for higher-order autocorrelation and test for autoregressive conditional heteroskedasticity, respectively. Cusum and Cusum Squared are tests of dynamic stability based on cumulative sums of residuals. Ramsey RESET tests the null hypothesis of no functional form misspecification. W_{LR} and W_{SR} denote Wald tests for a null hypothesis of long-run and short-run symmetry, defined by $-\hat{\gamma}^+/\hat{\beta}_1 = -\hat{\gamma}^-/\hat{\beta}_1$ and $\sum_{i=0}^{q}\hat{\mu}^+_i = \sum_{i=0}^{q}\hat{\mu}^-_i$, respectively.

Table 6 reports the estimation results for the NARDL model with the unemployment gap as an explanatory variable. The cointegration is confirmed in the case of Estonia and Lithuania (Table 3), and the long-run coefficients of positive and negative changes in the unemployment gap (L_ψ^+ and L_ψ^-) in Estonia are −0.09 and −0.30, and in Lithuania 0.04 and −0.15, respectively. The estimated long-run coefficients of positive changes are not statistically significant. So, we may conclude that a 1% decrease in the unemployment gap leads to a 0.3% increase in inflation in Estonia and a 0.15% increase in inflation in Lithuania. Also, the presence of long-run asymmetry is confirmed in the case of all three countries, whereas the evidence in favour of the short-run asymmetry is found for Latvia and Lithuania. The coefficients of the dummy variable are not statistically significant. As in the previous case, the results of diagnostic tests, stability, and functional form confirm this model is well-suited.

Overall, our findings indicate that the changes in the output gap have a larger long-run impact on the inflation rate than the unemployment gap. In both cases, that impact is asymmetric. Statistical significance of the long-run parameters of positive changes in the output gap and negative changes in the unemployment gap in Estonia and Lithuania can be explained by the trade-off between unemployment and output captured by Okun's law. In addition, that could be evidence in favour of a certain extent of downward price rigidity since prices react more when the real output tends to rise relative to the potential

Table 6. Estimation results for NARDL model with unemployment gap (Equation (13)).

Variable	Estonia	Latvia	Lithuania
α_2	0.041 (0.065)	0.032 (0.156)	0.081 (0.000)
π_{t-1}	−0.061 (0.093)	−0.056 (0.171)	−0.111 (0.001)
ψ_{t-1}^+	−0.005 (0.401)	−0.002 (0.798)	0.004 (0.413)
ψ_{t-1}^-	−0.018 (0.024)	−0.018 (0.043)	−0.016 (0.001)
$\Delta\pi_{t-1}$	0.401 (0.000)	0.648 (0.000)	0.539 (0.000)
$\Delta\pi_{t-2}$	0.311 (0.012)	0.244 (0.078)	-
$\Delta\pi_{t-4}$	-	-	0.257 (0.018)
$\Delta\psi_t^+$	−0.061 (0.005)	-	-
$\Delta\psi_{t-1}^+$	-	-	−0.058 (0.003)
$\Delta\psi_{t-2}^+$	−0.032 (0.085)	−0.037 (0.024)	0.054 (0.008)
$\Delta\psi_{t-3}^+$	-	-	−0.061 (0.002)
$\Delta\psi_{t-4}^+$	0.058 (0.005)	0.061 (0.002)	-
$\Delta\psi_t^-$	0.007 (0.703)	-	0.018 (0.297)
$\Delta\psi_{t-1}^-$	-	−0.052 (0.012)	-
$\Delta\psi_{t-3}^-$	-	−0.023 (0.381)	-
D	−0.006 (0.219)	−0.000 (0.992)	0.008 (0.058)
L_ψ^+	−0.089 (0.382)	−0.036 (0.812)	0.036 (0.349)
L_ψ^-	−0.295 (0.003)	−0.315 (0.013)	−0.145 (0.000)
R^2	0.521	0.788	0.619
F-stat.	6.606 (0.000)	24.841 (0.000)	11.037 (0.000)
JB test	0.983 (0.612)	2.729 (0.255)	1.095 (0.579)
BG LM test	0.549 (0.581)	0.285 (0.753)	1.254 (0.292)
ARCH test	0.879 (0.351)	0.092 (0.763)	1.658 (0.202)
Cusum test	Stable	Stable	Stable
Cusum Squared test	Stable	Stable	Stable
Ramsey RESET test	1.601 (0.210)	1.869 (0.176)	2.359 (0.063)
W_{LR}	11.018 (0.002)	18.893 (0.000)	52.898 (0.000)
W_{SR}	1.238 (0.269)	5.533 (0.022)	7.482 (0.008)

Note: The p-values are in parenthesis. L_ψ^+ and L_ψ^- denote estimated long-run coefficients of positive and negative changes in the unemployment gap, respectively, calculated as $L_\psi^+ = -\hat{\nu}^+/\hat{\beta}_2$ and $L_\psi^- = -\hat{\nu}^-/\hat{\beta}_2$ (Equation (13)). JB, BG LM and ARCH denote the Jarque-Bera test for normality, Breusch Godfrey test for higher-order autocorrelation and test for autoregressive conditional heteroskedasticity, respectively. Cusum and Cusum Squared are tests of dynamic stability based on cumulative sums of residuals. Ramsey RESET tests the null hypothesis of no functional form misspecification. W_{LR} and W_{SR} denote Wald tests for a null hypothesis of long-run and short-run symmetry, defined by $-\hat{\nu}^+/\hat{\beta}_2 = -\hat{\nu}^-/\hat{\beta}_2$ and $\sum_{j=0}^{q} \hat{\tau}_j^+ = \sum_{j=0}^{q} \hat{\tau}_j^-$, respectively.

one (actual unemployment falls below the NAIRU rate) than in the opposite case. These results coincide very well with the study of Dabušinskas and Kulikov (2007), who find evidence on price setting rigidity in the Baltic States and the significant impact of backward-looking behaviour in forming inflation expectations. In addition, some other empirical studies for European countries (including the Baltic economies), such as, for instance, Babetskii (2007) and Branten et al. (2018), also document the presence of price rigidity, mainly as a consequence of the nominal wage stickiness.

6. Possible causes of observed asymmetries

We now turn to further investigation of possible determinants of the asymmetric response of inflation to changes in economic activity in the Baltic States. Since there is some evidence in favour of downward price rigidity as a source of the asymmetry, an additional analysis of its possible causes could contribute to better understanding the inflation dynamics in these countries.

Considering the inability of policymakers in the Baltic States (as in the rest of the EMU member countries) to respond to macroeconomic shocks by using adjustment roles of the nominal exchange rate and independent monetary policy, the flexibility of the labour market become an important issue often stressed in the relevant literature (e.g. Anderton & Bonthuis, 2015; Babetskii, 2007; European Commission, 2003; Von Hagen & Traistary, 2005). Although labour market in the Baltic economies is commonly considered to exhibit institutional flexibility, due to low union density rate, low employment protection strictness and so on, as stressed by Eamets and Masso (2005), it must be noted that such flexibility can be accompanied by wage rigidity due to lower hiring and firing costs. Accordingly, our analysis of the asymmetry determinants is focused on the labour market and the role of nominal and real wage rigidity.

The presence of wage rigidity can prevent the complete adjustment of prices and have an impact on the inflation dynamics, as documented in a number of studies (e.g. Arpaia & Pichelmann, 2007; Babetskii, 2007; Branten et al., 2018; Daly et al., 2012; Daly & Hobijn, 2014; Iwasaki et al., 2018). This is due to the fact that labour compensation is a key determinant of firms' pricing behaviour. Indeed, as it is apparent from Figure 2, the nominal wage rate in all three Baltic economies tends to follow similar trends to the inflation rate, only with a higher magnitude. The lower panel of Figure 2 shows the relationship between real wages and productivity. The increases in real wages in excess of productivity gains put upward pressure on prices, which is evident in the pre-crisis period in all three Baltic economies, as well as in the period from 2012 onwards. During the Great Recession, real wages were significantly decreased, and their growth in subsequent years has been more moderate than productivity growth.

If some extent of downward wage rigidity is present in the Baltic economies that could explain, at least partially, the identified asymmetry in the effects of inflation drivers. Since an often-used measure of wage flexibility is the responsiveness of wages to the rate of unemployment, we use so-called wage Phillips curve, given as follows (Babetskii, 2007):

$$\Delta w_t = a_1 + a_2\psi_t + a_3\Delta\Upsilon_t + a_4\Delta\pi_t + \varepsilon_t, \tag{16}$$

where a_1 is a constant, Δw_t denotes nominal wage rate, ψ_t is the unemployment gap, $\Delta\Upsilon_t$ and $\Delta\pi_t$ refer to the rate of productivity growth and inflation rate, respectively, and ε_t

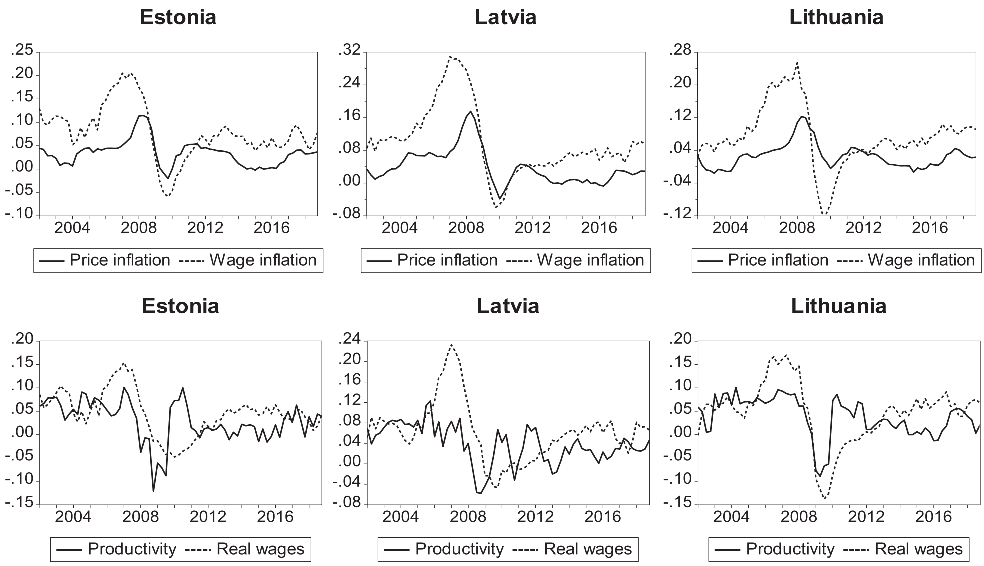

Figure 2. Empirical dynamics of the inflation rate, nominal and real wages and average productivity in the Baltic States.

Note: The data series refer to the percentage change of variables compared to the same quarter in the previous year.

represents an error term. Coefficient a_2 represents the responsiveness of the nominal wage rate to the unemployment gap and thus the wage flexibility. Following López-Villavicencio and Saglio (2012), we reformulate Equation (16) as a nonlinear ARDL model in the following form:

$$\Delta w_t = a_0 + a_1 w_{t-1} + a_2^+ \psi_{t-1}^+ + a_2^- \psi_{t-1}^- + a_3 \Upsilon_{t-1} + a_4 \pi_{t-1} + \sum_{j=1}^{m} \alpha_j \Delta w_{t-j} + \sum_{j=0}^{n} (\beta_j^+ \Delta \psi_{t-j}^+ + \beta_j^- \Delta \psi_{t-j}^- + \delta_j \Delta \Upsilon_{t-j} + \theta_j \Delta \pi_{t-j}) + \varepsilon_t. \tag{17}$$

where a_0 denotes constant, whereas a_1, a_2^+, a_2^-, a_3, a_4 and α_j, β_j^+, β_j^-, δ_j, θ_j are the long-run and short-run coefficients, respectively. The partial sum of positive (ψ_t^+) and negative (ψ_t^-) changes in the unemployment gap are generated using Equations (11) and (12), respectively. Equation (17) can be rewritten in terms of real wage (as CPI-deflated nominal wage) as follows:

$$\Delta(w_t/\pi_t) = a_0 + a_1 w_{t-1} + a_2^+ \psi_{t-1}^+ + a_2^- \psi_{t-1}^- + a_3 \Upsilon_{t-1} + \sum_{j=1}^{m} \alpha_j \Delta w_{t-j} + \sum_{j=0}^{n} (\beta_j^+ \Delta \psi_{t-j}^+ + \beta_j^- \Delta \psi_{t-j}^- + \delta_j \Delta \Upsilon_{t-j}) + \varepsilon_t. \tag{18}$$

Due to space limitations, we focus only on the long-run rigidities.[1] By calculation of the long-run coefficients for the positive and negative changes in the unemployment gap ($L_\psi^+ = -a_2^+/a_1$ and $L_\psi^- = -a_2^-/a_1$), we are able to capture asymmetries that may result from wage rigidities (López-Villavicencio & Saglio, 2012). Namely, if they are both significant and negatives, but L_{ψ}^{-} is larger than L_{ψ}^{+}, we conclude that there exists some downward wage rigidity.

The dataset consists of the quarterly time series about total wages and salaries (business economy), real labour productivity per employed person, unemployment rate and rate of inflation, collected from the Eurostat database. The analysed time span is from the first quarter of 2002 to the fourth quarter of 2018. As in the previous section, we use the dummy variable in order to capture the effects of the Great Recession.

The estimated long-run parameters are reported in Table 7. The values of the long-run coefficients for the unemployment gap suggest that there is some downward nominal wage rigidity in all three countries. A 1% increase in unemployment leads to a 0.28% decrease in the nominal wages in Estonia, 0.23% in Latvia and 0.21% in Lithuania. On the other hand, a 1% decrease in unemployment leads to a 0.33% increase in nominal wages in Estonia, 0.28% in Latvia and 0.23% in Lithuania. The inflation and productivity are positively associated with the nominal wage rate, but their coefficients are statistically significant only in Estonia.

Although the results confirm the presence of wage rigidity that could account for an identified asymmetry in the inflation response to economic activity, the Wald test for long-run asymmetry ($W_{LR(\psi)}$) indicates that only in Estonia nominal wages tend to react asymmetrically with respect to the changes in unemployment. The estimated long-run coefficients for the real wages in Estonia also indicate the presence of downward rigidity: a 1% decrease in unemployment leads to a 0.76% increase in the real wages, whereas a 1% increase in unemployment results in a real wages reduction for only 0.38%. This can be explained by the fact that the avoidance of cuts in nominal wages becomes an obstacle for real wage adjustments, especially during periods of low inflation.

Notwithstanding the evidence on downward wage rigidity, it must be stressed that the statistical significance and the negative sign of long-run coefficients for positive and negative changes in the unemployment gap indicate that nominal wages in the Baltic States are generally relatively flexible. In other words, nominal wages tend to rise when unemployment decreasing and to fall (a little bit slighter) when unemployment increasing. Overall, the values of these coefficients reveal that the unemployment elasticity of wages is the

Table 7. Long-run estimates for the wage Phillips curve parameters (Equation (17)).

Variable	Estonia	Latvia	Lithuania
Nominal wages			
a_1	−0.541 (0.000)	−0.218 (0.052)	−0.581 (0.000)
L_{ψ}^{+}	−0.277 (0.000)	−0.234 (0.009)	−0.211 (0.000)
L_{ψ}^{-}	−0.326 (0.000)	−0.287 (0.000)	−0.231 (0.000)
L_{π}	0.625 (0.002)	0.274 (0.611)	0.255 (0.304)
L_{Y}	0.325 (0.043)	0.472 (0.359)	0.312 (0.087)
$W_{LR(\psi)}$	6.067 (0.017)	0.969 (0.329)	0.791 (0.378)
Real wages			
a_1	−0.081 (0.010)	−0.035 (0.274)	−0.052 (0.129)
L_{ψ}^{+}	−0.377 (0.067)	−0.507 (0.384)	−1.268 (0.085)
L_{ψ}^{-}	−0.764 (0.011)	−0.882 (0.342)	−0.864 (0.066)
L_{Y}	0.621 (0.126)	3.073 (0.118)	3.329 (0.091)
$W_{LR}(\psi)$	9.905 (0.003)	0.685 (0.412)	1.096 (0.300)

Note: The *p*-values are in parenthesis. L_{ψ}^{+} and L_{ψ}^{-} denote estimated long-run coefficients of positive and negative changes in the unemployment gap, respectively, calculated as $L_{\psi}^{+} = -\hat{a}_2^{+}/\hat{a}_1$ and $L_{\psi}^{-} = -\hat{a}_2^{-}/\hat{a}_1$ (Equation (17)). L_{π} and L_{Y} are the estimated long-run coefficients for inflation and productivity, defined as $L_{\pi} = -\hat{a}_3/\hat{a}_1$ and $L_{Y} = -\hat{a}_4/\hat{a}_1$, respectively. $W_{LR(\psi)}$ denote Wald test for a null hypothesis of long-run symmetry of the relationship between unemployment gap and wage rate, defined by $= -\hat{a}_2^{+}/\hat{a}_1 = -\hat{a}_2^{-}/\hat{a}_1$.

highest in Estonia and the lowest in Lithuania. Our results are in line with the findings of Paas et al. (2002), who documented that the nominal wages are most rigid in Lithuania and most flexible in Estonia. The estimated values of the long-run coefficients also coincide with the research of Blanchflower (2001), who finds that the unemployment elasticity of wage in 15 transition countries (including the Baltic States) range from −0.02 to −0.46.

7. Conclusions

This study investigated the asymmetries in the effects of the output gap and unemployment gap on the inflation rate in the Baltic States, employing a nonlinear ARDL approach based on the Phillips curve. We found empirical evidence on long-run asymmetry in the impact of these inflation drivers, indicating that the Phillips curves in these countries are nonlinear. The estimated long-run coefficients indicate stronger and statistically significant effects of positive changes in the output gap on inflation in Estonia and Lithuania and negative changes in Latvia. The negative changes in the unemployment gap have a stronger and significant long-run impact on the inflation in all three countries, whereas the short-run asymmetry is revealed in the case of Latvia and Lithuania.

These findings indicate that there is some extent of downward price rigidity in the Baltic States. Further analysis was focused on the labour market flexibility as a possible source of identified asymmetries, having in mind the inability of policymakers in these economies to respond to macroeconomic shocks by using adjustment roles of the nominal exchange rate and independent monetary policy. The results documented the presence of downward nominal wage rigidities with respect to unemployment changes in all three countries, which could, at least partially, explain the observed asymmetry in the inflation response. However, the robustness check confirms the asymmetry of the wage-unemployment nexus only in Estonia.

According to the results, it appears that the disinflation policy directed to the aggregate demand reduction and inducing negative (positive) changes in the output gap (unemployment gap) is not likely to have a significant effect on inflation rate decelerating. Besides that, the disinflation policy could produce large costs due to downward price and wage rigidity. One can conclude that all Baltic economies, especially Estonia, could benefit from increasing downward wage flexibility. An additional problem is a lack of monetary sovereignty in the Baltic States as members of the Eurozone, which narrows the space for implementation of monetary policy measures. Hence, the fiscal policy remains a valuable tool, with different instruments of taxation and government spending that should be directed to the aggregate supply. These measures might lead to productivity growth acceleration, thus reducing the real marginal costs and relaxing the upward pressures on inflation. An important corollary of boosting productivity is the potential output growth, which is particularly needed for the Baltic States having in mind they operate above the potential, as reported in the official documents cited in this study.

Finally, there is a limitation of this research that should be noted, relating to the estimation of the output gap and unemployment gap. For instance, the accuracy of the output gap calculation depends on how precise one can estimate the potential output, which is further conditioned by the choice of filtering technique (Orphanides & van Norden, 2005). Besides that, the relationship between the output and unemployment gap and inflation is known only *a posteriori*, indicating that they cannot be so useful for inflation forecasting. However, it doesn't mean that knowing the characteristics of their relationship with inflation cannot provide useful information for economic policymaking. In addition, we focused on the domestic inflation drivers, which are traditionally incorporated in the Phillips curve, whereas the inflationary effect of the external economic environment appears only indirectly. It might make a model too country-specific but, at the same time, it is a way to capture important relations among relevant variables in order to propose adequate policy measures.

Some of the future research could take into account the impact of the domestic and external inflation drivers in other countries, such as the Central and Eastern European economies, as well as the most advanced economies of the West Europe and the United States. This would make a field for comparative analysis of inflation drivers' characteristics, which could provide additional information for a better understanding of the inflation process. That analysis should be dedicated to the investigation of the nonlinearities in macroeconomic relations connected with the inflation dynamics, as a promising field of research, what this study also demonstrated.

Note

1. The short-run estimates and the results of all pre- and post-estimation tests are available upon request.

Disclosure statement

No potential conflict of interest was reported by the author(s).

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