

A Panel Smooth Transition Model for the Exchange Rate Pass-Through: New Evidence from the New EU Member States

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Abstract

This paper studies the existence of nonlinear behaviour in exchange rate pass-through (ERPT) for a set of new EU Member States (NMS). To account for nonlinearities, we implement a Panel Smooth Transition Regression (PSTR) approach, where the extent of pass-through is allowed to respond nonlinearly to the inflation environment. The main advantage of PSTR framework is the use of a grid search to select endogenously the threshold level(s) allowing to identify different inflation regimes. Using quarterly data for ten NMS over 1996-2013, our results suggest a significant regime-dependence of the pass-through to inflation levels. We find that in the long-run the sensitivity of import prices to exchange rate is about 0.78% when CPI inflation is below a threshold of 4.56%, but beyond this threshold level, the degree of pass-through becomes higher and reaches a full ERPT. Furthermore, our findings corroborate those obtained for the NMS in the sense that falling inflation rates are closely linked to the decline in pass-through estimates since the late 1990s, although ERPT in the NMS remains on average larger than for industrial countries. These results contribute to lively debate on the path towards the Euro and how ECB could reduce the inflation differentials with the enlargement of the monetary union.

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1 Introduction

In January 1, 2015, Lithuania is 19th country to adopt the European single currency and its national central bank becomes a member of the Eurosystem. Lithuania is one of the central and eastern European economies to get EU membership on May 2004 with Czech Republic, Poland, Hungary, Slovenia, Slovakia, Latvia, Estonia, Malta and Cyprus. Now is joining the Eurozone after Slovenia, Slovakia, Latvia, Estonia, Malta and Cyprus. For the new EU member states (NMS), forgoing their local currencies to join a monetary union would pose a significant challenge, since a country adopting the euro cedes its monetary policy to the European Central Bank (ECB), and no longer has the option of using monetary policy to respond to local conditions. The impact of the monetary policy decisions on the single currency may induce different effects on expenditure switching and price level movements regarding the extent of the transmission of exchange rate changes to domestic prices. Thereby, a common exchange rate movement, in the absence of a national monetary policy, may have differential impact on different the NMS of of the European Union, leading notably to possible divergence in inflation rates.

Besides, these countries have made substantial progress in transforming their economies since the demise of the Soviet-type communist regimes at the start of the 1990s. They have developed structural reforms and implemented macroeconomic stabilization programs of great diversity in monetary policy frameworks and exchange rate regimes, which could affect the extent to which exchange rates changes affect domestic prices. For instance, the NMS have maintained different exchange systems over time. Different regimes was adopted included currency boards, fixed pegs to a basket, crawling pegs, managed float and free float. As discussed in literature on the NMS, the exchange rate pass-through (ERPT) may differ according to the nature of the exchange rate regime in place (see e.g. Bitans, 2004; Coricelli et al., 2006; María-Dolores, 2010; Beirne and Bijsterbosch, 2011, among others). Also, there was a dramatic changes in the inflation levels across Central and Eastern European Economies. Most of the central and eastern European countries shifted from a high inflation environment, at the beginning of the 90s, to a relatively moderate inflation rate by the late 90s. Thus, measures of the degree of pass-through may differ across studies due to the factors that have changed during the transition process.

As a matter of fact, previous studies followed diverse approaches in order to take into this changing macroeconomic environment in the NMS of the EU. The major drawback of the cited literature is the failure to account for policy shifts by using the relevant empirical techniques. This could result in biased and underestimated ERPT estimates. This issue is of notable importance due

to the great diversity in exchange rate regimes and in inflation environment the central and eastern European countries. For instance, several studies assume linearity in the transmission of exchange rate changes rather than testing it. Although they have put forth the role of macroeconomic factors, they fail to recognize the nonlinear behaviour of ERPT mechanism.

For more accuracy, in our study we propose to implement panel smooth transition regression (PSTR) model in order to capture the nonlinear behavior of the pass-through. We investigate how the inflation environment impact nonlinearly the ERPT. The main advantage of PSTR framework is the use of a grid search to select endogenously threshold level(s) allowing to identify different inflation regimes. We focus on ten new EU member states (Bulgaria, Czech Republic, Estonia, Hungary, Lithuania, Latvia, Poland, Romania, Slovenia and Slovakia) where a marked diminishing tendency in the inflation rates was started at the end of the 1990s. The model is estimated using quarterly data spanning the period 1996:1 to 2013:4.

The rest of the paper is organized as follows. Section 2 gives some stylised facts regarding inflation and exchange rate developments in the NMS during the past decade. The econometric approach is provided in the Section 3. Section 4 describes the data and the empirical specification. Section 5 gives the main empirical results and Section 6 concludes.

2 Pass-Through in the new EU Member States

In the beginning of the 1990s with the demise of the Soviet-type communist regimes, the NMS have made substantial progress in transforming their economies. They have developed structural reforms and implemented macroeconomic stabilization programs of great diversity in monetary policy frameworks and exchange rate regimes, which could affect the extent to which exchange rates changes affect domestic prices. Today a large majority of these countries have been able to join the European Union, in May 2004 (Czech Republic, Poland, Hungary, Slovenia, Slovakia, Latvia, Estonia, Lithuania, Malta, and Cyprus) and January 2007 (Romania and Bulgaria), and some of them have recently adopted the euro as a currency such as Slovenia, Slovakia, Latvia, Estonia and Lithuania.¹

As reported in the pass-through literature, many economic policy issues, such as the persistence of inflation, and the effect of entering into a monetary union, could influence the determination of the ERPT rate to prices and its

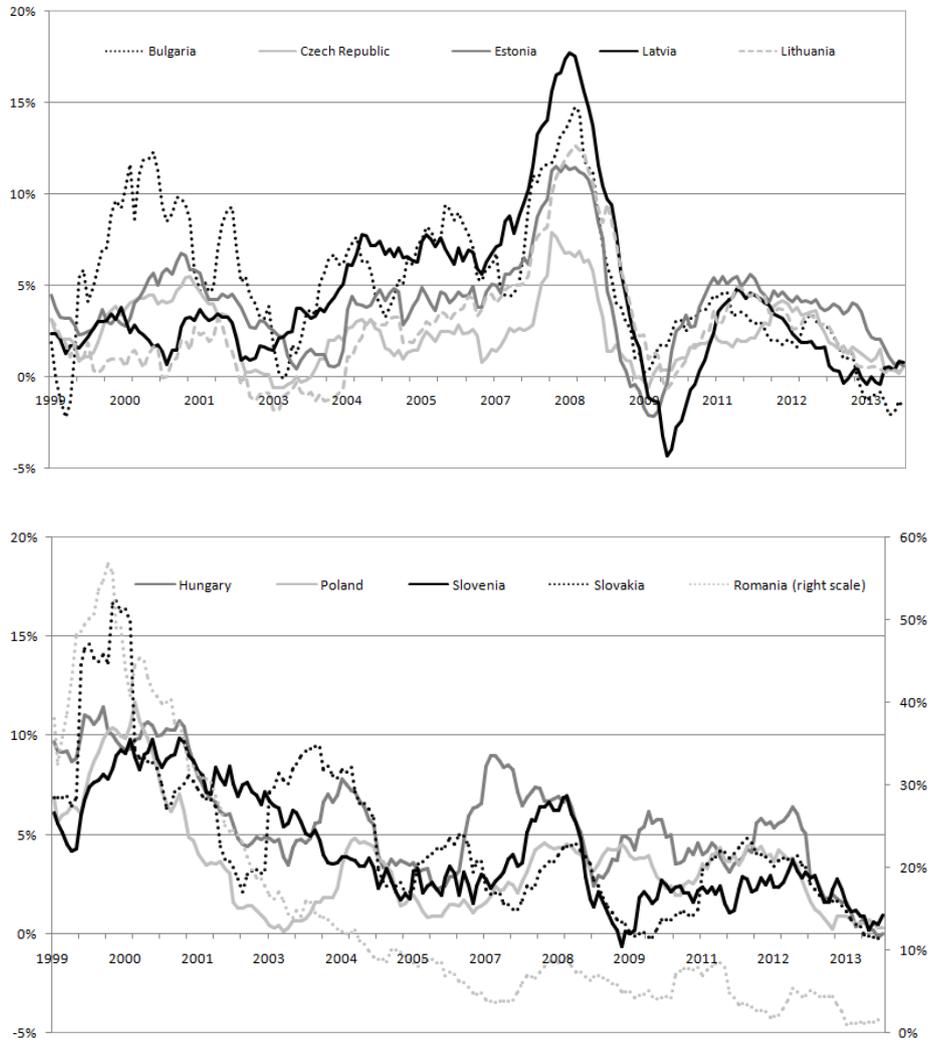
¹ The last one is Lithuania who has joined the monetary union by adopting the euro on January 1, 2015, and becomes the 19th Member State to adopt the single currency.

evolution in different time horizons and sectors. In fact, there has been a lively debate on the path towards the adoption of the single currency for the NMS. On one hand, NMS countries have to meet an inflation criterion as set out in the Maastricht Treaty as well as their own inflation performances, which could be influenced by the degree of pass-through. Once they belong to the euro area, the effect of different rates of ERPT could contribute to national inflation differentials. On the other hand, the issue of the desirability and feasibility of maintaining flexible exchange rates and an independent monetary policy. As is well-known, the choice of a fixed or flexible exchange rate regime could influence the transmission of the currency movements to prices.

In this context, the pass-through of exchange rate changes to domestic prices is of considerable importance, as it determines the strength of the expenditure-switching effect and, in turn, the effectiveness of independent monetary policy in response to adverse macroeconomic shocks using the nominal exchange rate as the instrument. If the expenditure-switching effect is strong and the country has yet to achieve a satisfactory level of both nominal and real economic convergence with respect to the rest of the potential monetary union members, it may be advisable to postpone the date of entry into the Euro area because the adjustment with an independent monetary policy may be faster and smoother. Such independent monetary policy would also ensure better control of inflation and higher degree of output stability. If, on the other hand, the ability of exchange rate fluctuations to cushion idiosyncratic shocks is very low, or even counterproductive, it would be preferable to choose the strategy of fast Euro adoption, as this would not result in additional welfare cost, while the benefits of monetary union membership could be reaped sooner.

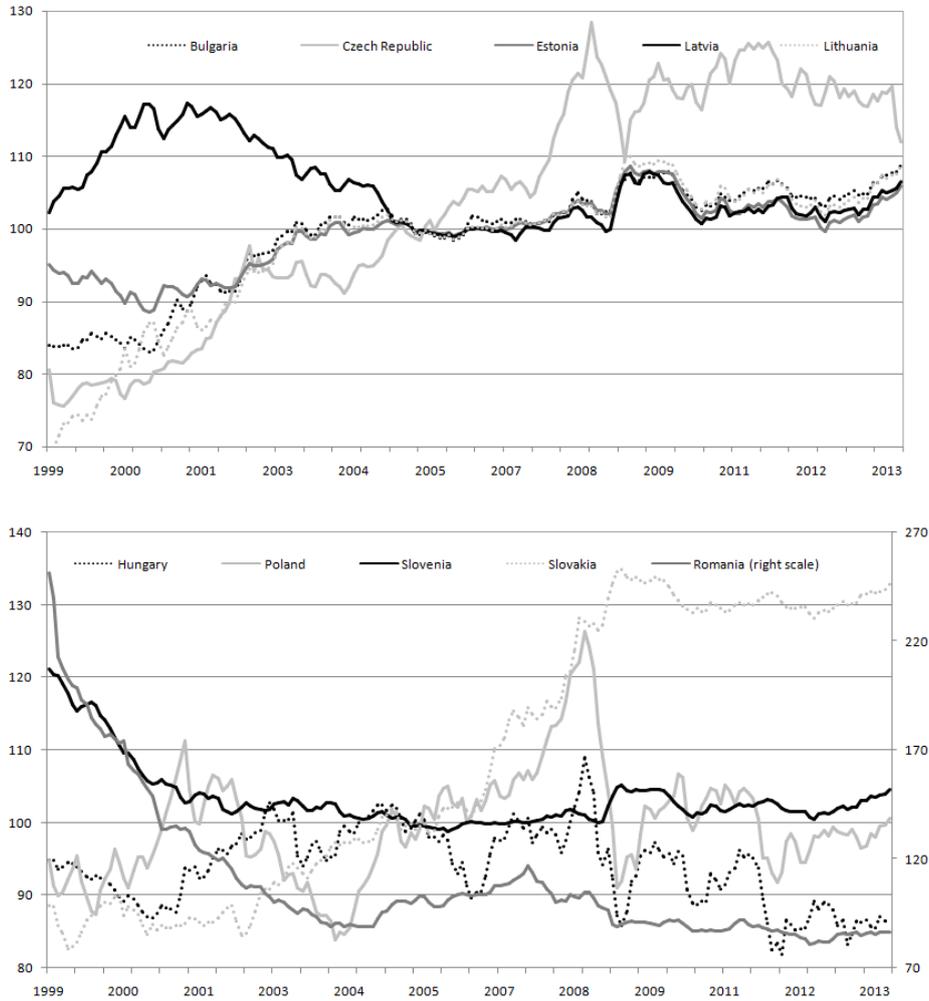
From a policy point of view, studying the extent of pass-through for the NMS is very appealing especially regarding the dramatic changes in the inflation environment. Most of the central and eastern European countries commuted from a high inflation environment, at the beginning of the 90s, to a relatively moderate inflation rate by the late 90s (see Figure 1). Between 2003 and 2005, however, inflation started to rise again in most countries in the region, particularly in Bulgaria, Estonia, Latvia and Lithuania, reaching a peak in 2008. In Romania and Slovakia, inflation continued to register a global descending trend compared to 2003. In all, when comparing the inflation regime before and after joining the EU, we point out a very clear diminishing tendency in the inflation rate for the large majority of NMS compared to the end of the 90s (see Figure 3).

Figure 1: Inflation rates evolution in the NMS over 1999-2013



Notes: Inflation levels in Romania are scaled right due to their higher variability in comparison to the rest of NMS. Source: Eurostat.

Figure 2: Nominal effective exchange rates development in the NMS during 1999-2013



Notes: There was a large decline in Nominal effective exchange rates for Romania over 1999-2013, that is why they are scaled right. Source: Eurostat.

Regarding exchange rate developments, the NMS have maintained different exchange systems over time. Different regimes were adopted included currency boards, fixed pegs to a basket, crawling pegs, managed float and free float. For instance, before joining the Eurozone, Estonia and Lithuania with a currency board as a unilateral commitment and Latvia with a ± 1 percent fluctuation band as a unilateral commitment. The different patterns of inflation dynamics in the NMS seem to be associated with different exchange rate regimes. As a matter of fact, developments in the nominal (effective) exchange rate, as

displayed in Figure 2, can have a short-term character but can equally be structural when associated with the real appreciation trend experienced by the catching-up economies. The gradual decline of inflation during the early stages of transition has been accompanied by a sizable appreciation of the real exchange rate. Where exchange rates were fixed (Bulgaria, Estonia, Latvia and Lithuania), the appreciation came as a result of inflation, while where a more flexible exchange rate regime was in place (Czech Republic, Hungary, Poland, Romania and Slovakia), the appreciation was due to a combination of nominal appreciation and inflation. A component of this trend appreciation can be considered an equilibrium phenomenon, in line with the Balassa-Samuelson effect that affects inflation and the real exchange rate in a catching-up phase.

In all, the degree of pass-through in central and eastern European economies may differ according to the nature of the exchange rate regime in place, and would have crucial policy implications for the new EU member states. Where the exchange regime is fixed, as in in monetary union, the real appreciation passes exclusively through an inflation differential. On the contrary, in a flexible exchange regime with inflation targeting, we might observe a tendency of appreciation in the nominal exchange rate, as the inflation target makes the existence of a sensible inflation differential less probable.

3 Econometric Approach

In order to investigate for the regime-dependence of ERPT to inflation environment, a nonlinear panel smooth transition model is estimated for the new EU Member States. Following González et al. (2005) and Fok et al. (2005), we consider the PSTR model with two extreme regimes and a single transition function:

$$y_t = \mu_i + \beta'_0 x_{it} + \beta'_1 x_{it} g(q_{it}; \gamma, c) + \varepsilon_{it}, \quad (1)$$

for $i = 1, \dots, N$, and $t = 1, \dots, T$, where N and T denote the cross-section and time dimensions of the panel, respectively. y_{it} is a scalar, μ_i is an unobservable time-invariant regressor, x_{it} is a k -dimensional vector of time-varying exogenous variables, q_{it} is an observable transition variable and ε_t are the residuals. Transition function $g(q_{it}; \gamma, c)$ is a continuous function of the observable variable q_{it} and bounded between 0 and 1, where c denotes the location parameter or the threshold between the two extreme regimes and γ determines the smoothness of the transition. Following the work of Teräsvirta

(1994) for the time series STAR models, González et al. (2005) consider the following logistic transition function:

$$g(q_{it}; \gamma, c) = \frac{1}{1 + \exp[-\gamma(q_{it} - c)]}, \quad \text{with } \gamma > 0 \quad (2)$$

The estimation of the PSTR in equation (1) consists of two steps. First, fixed effects are eliminated by removing individual-specific means. Then, in the second step, parameters γ and c are estimated by Non Linear Least Squares (NLS). An important issue to take into account is the selection of starting values of γ and c , since this notably determines the convergence procedure. To select good starting values, a two-dimensional grid search of 50 values of γ and 100 values of c is carried out. Given these grids, the vector with the minimum residual sum of squares is used to estimate the corresponding β_1 and β_2 .

Before estimating the PSTR it is important to test whether the regime-switching effect is statistically significant. González et al. (2005) outlined a procedure for testing linearity against a PSTR model. This is deemed important since the PSTR is not identified if the data-generating process (DGP) is linear, therefore a linearity test is viewed to be necessary to avoid the estimation of unidentified models. The null hypothesis is $H_0 : \beta_1 = 0$. The test is non-standard because under this null hypothesis, the PSTR model contains unidentified nuisance parameters. One solution adopted by González et al. (2005) is to replace $g(q_{it}; \gamma, c)$ by a first-order Taylor expansion around $\gamma = 0$ and test the linearity hypothesis as $H_0 : \gamma = 0$. After reparameterization, this leads to the following auxiliary regression:

$$y_t = \mu_i + \beta_0^* x_{it} + \beta_1^* x_{it} q_{it} + \varepsilon_{it}^* \quad (3)$$

where the parameters β_i^* are proportional to the slope parameter γ and $\varepsilon_{it}^* = \varepsilon_{it} + R_1 \beta_1^* x_{it}$, where R_1 is the remainder of the Taylor expansion. Therefore testing $H_0 : \gamma = 0$ in equation (1) is equivalent to testing $H_0^* : \beta_1^* = 0$ in equation (3). We can use the Wald, Fisher and Likelihood Ratio tests and their statistics respectively defined are as follows:

$$LM_w = \frac{TN(SSR_0 - SSR_1)}{SSR_0} \quad (4)$$

$$LM_F = \frac{TN(SSR_0 - SSR_1)/k}{SSR_0/(TN - N - k)} \quad (5)$$

$$LR = -2[\log(SSR_1) - \log(SSR_0)] \quad (6)$$

where k is the number of explanatory variables, SSR_0 is the panel sum of squared residuals under H_0 (linear panel model with individual effects) and SSR_1 is the panel sum of squared residuals under H_1 (i.e., the PSTR model with two regimes). Under the null hypothesis, the LM_w and LR statistics are distributed as a $\chi^2(k)$ and the LM_F statistic has an approximate $F(k, TN - N - k)$ distribution.

After estimating of at least one transition function, the next phase consists of testing the number of transition functions that must be included in the specification, *i.e.* the existence of remaining nonlinearity. Following a sequential procedure, as in González et al. (2005), we generalize the test to a r number of regimes to determine the number of transitions in the model.

$$y_t = \mu_i + \beta'_0 x_{it} + \sum_{j=1}^r \beta'_j x_{it} g_j(q_{it}^{(j)}; \gamma_j, c_j) + \varepsilon_{it}, \quad (7)$$

where the transition functions $g_j(q_{it}^{(j)}; \gamma_j, c_j)$, $j = 1, \dots, r$, are of the logistic type as in (2). For instance, suppose that we want to test whether there is one transition function, ($H_0 : r = 1$) versus there are at least two transition functions ($H_0 : r = 2$). Thus, we can write a model with $r = 2$ as:

$$y_t = \mu_i + \beta'_0 x_{it} + \beta'_1 x_{it} g_1(q_{it}^{(1)}; \gamma_1, c_1) + \beta'_2 x_{it} g_2(q_{it}^{(2)}; \gamma_2, c_2) + \varepsilon_{it}, \quad (8)$$

The rationality of the procedure is as before, that is, we replace the second transition function by its first-order Taylor expansion around $\gamma_2 = 0$ and then in testing linear constraints on the parameters. Thus, the first-order Taylor approximation of $g_2(q_{it}^{(2)}; \gamma_2, c_2)$ leads to the following auxiliary regression:

$$y_t = \mu_i + \beta'_0 x_{it} + \beta'_1 x_{it} g_1(q_{it}^{(1)}; \gamma_1, c_1) + \theta' x_{it} q_{it} + \varepsilon_{it}^* \quad (9)$$

Now, we test the null hypothesis that there is one transition function, versus the alternative that there are two. The null hypothesis of no remaining nonlinearity can thus be defined as $H_0 : \theta' = 0$. A similar logic is followed to compute the Wald, Fisher and Likelihood Ratio Tests as before.

4 Empirical specification and Data

The standard specification used throughout ERPT literature is based on the pricing behaviour of exporting firms. This suggests the following log-linear regression specification:

$$p_{it} = \alpha_i + \beta e_{it} + \gamma y_{it} + \delta w_{it}^* + \varepsilon_{it}, \quad (10)$$

where p_{it} is the local currency import prices, e_{it} is the nominal exchange rate, y_{it} is a primary control variable used to capture domestic demand conditions, and w_{it}^* is a measure of foreign producer cost which is a second control variable representing foreign exporters' costs. Our primary concern in this study is the pass-through elasticity which corresponds to the coefficient on the exchange rate β which is expected to be bounded between 0 and 1. A one-for-one pass through to changes in import prices, known as a complete ERPT, is given by $\beta = 1$. In this case, exporters let the domestic currency import prices affected by exchange rate move. While, when exporters adjust their markup, a partial or incomplete ERPT occurs and $\beta < 1$.

The aim of our empirical analysis is to give further evidence on the nonlinear behaviour of the ERPT with respect to inflation environment. As showed in Figure 3, there was a global descending trend in inflation rates for almost the NMS when comparing the inflation regime before and after joining the EU. Several studies underline the importance of inflation regime (see e.g. Taylor, 2000; Choudhri and Hakura, 2006; Bailliu and Fujii, 2004; Gagnon and Ihrig, 2004, among others). Countries with low relative inflation variability or stable monetary policies are more likely to have their currencies chosen for transaction invoicing, and hence more likely to have low pass-through to domestic prices. However, the major drawback is to assume linearity rather than testing it. For example, in their panel-data set of 11 industrialized countries, Bailliu and Fujii (2004) introduced interaction terms in their ERPT equation between the rate of change in the exchange rate and dummy variables that capture changes in the inflation environment. The authors use multiple break test developed by Bai and Perron (1998) to determine the timing of shifts in the inflation environments. For more accuracy, in our study we propose to implement a panel smooth transition framework where a grid search is used to select endogenously threshold level(s) allowing to identify different inflation regimes.

In order to capture the nonlinear behavior of the ERPT mechanism with respect to inflation environment, we extend the linear pass-through regression (10) by the following nonlinear specification:

$$p_{it} = \alpha_i + \beta e_{it} + \gamma y_{it} + \lambda w_{it}^* + \phi e_{it} g(q_{it}; \gamma, c) + \varepsilon_{it}, \quad (11)$$

The transition variables used as a measure of inflation environment is the CPI inflation rate $q_{it} = \pi_{it}$. According to equation (11), the ERPT elasticity is given by the following coefficient:

$$\text{ERPT} = \frac{\delta p_{it}}{\delta e_{it}} = \beta + \phi g(q_{it}; \gamma, c) \quad (12)$$

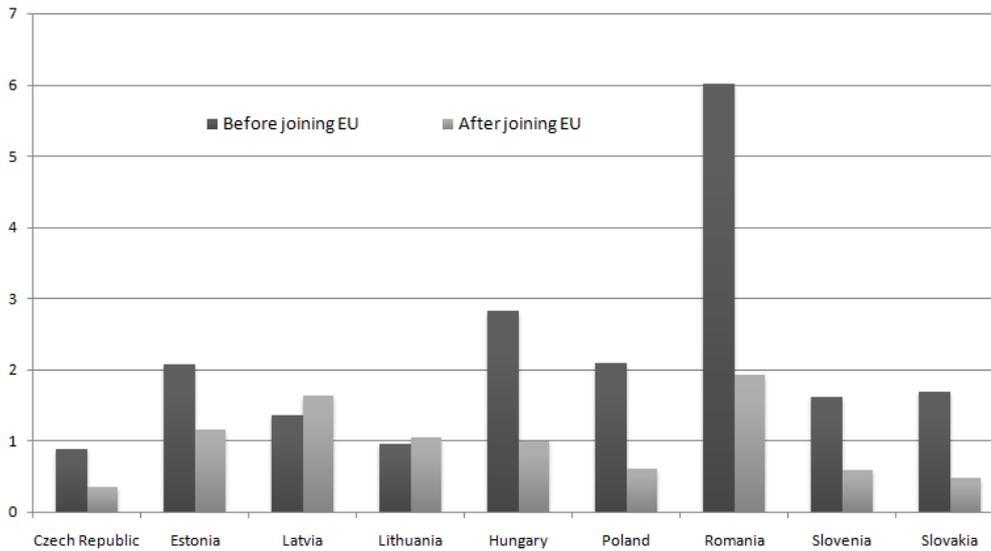
Due to the features of logistic STR models, long-run ERPT coefficient is expected to take different values depending on whether the transition variable, *i.e.* inflation level, is below or above the threshold. If the CPI inflation is below the threshold, *i.e.* $(q_{it} - c) \rightarrow -\infty$ which means $g(q_{it}; \gamma, c) = 0$, then the importing country experiences a low-inflation regime and pass-through elasticity is equal to: ERPT = β . If inflation rate is above the threshold value, *i.e.* $(q_{it} - c) \rightarrow +\infty$ which means $g(q_{it}; \gamma, c) = 1$, then the economy is in higher inflation episodes and pass-through coefficient becomes: ERPT = $\beta + \phi$.

Our PSTR pass-through equation (11) is estimated for ten new EU Member States (Bulgaria, Czech Republic, Estonia, Hungary, Lithuania, Latvia, Poland, Romania, Slovenia and Slovakia), using quarterly data spanning the period 1996:1 to 2013:4. All these data are extracted from the Eurostat database. For the domestic import prices p_{it} , we use import price index seasonally adjusted. as a proxy for the domestic demand y_{it} , we use the real GDP. For the exchange rate e_{it} , we use the nominal effective trade weighted series, with an increase means depreciation of the national currency, and a decrease means appreciation. To capture changes in foreign costs, we follow Bailliu and Fujii (2004) by constructing an exporter partners' cost proxy. In logarithms, this latter is measured as follow: $w_t^* \equiv q_t + ulc_t - e_t$, where q_t is the unit labor cost (ULC) based real effective exchange rate, ulc_t is the ULC in domestic country and e_t the nominal effective exchange rate. Finally, for our the transition variable, inflation rates series represents the quarterly change in consumer prices index (CPI).

To avoid the problem of spurious regression, as in Ben Cheikh and Cheik (2013), we employ a class of panel unit root and panel cointegration tests which allow for serial correlation between the cross-sections, *i.e.* the so-called second generation tests. We use the cross-sectionally augmented IPS panel unit root tests by Pesaran (2007) and the error-correction-based tests for panel cointegration by Westerlund (2007), which both account for possible cross-sectional dependencies. Panel unit root tests results are shown in Table 3 in Appendix A for all variables entering pass-through equation (10) in for both levels and first differences. In the level case, we are unable to reject the null hypothesis the null hypothesis of a unit root. For tests on the first differences, we can see that the null of non-stationarity is strongly rejected. Consequently, for variables belonging to the pass-through equation are $I(1)$. Next, we implement Westerlund's (2007) panel cointegration test where results are reported in Table 4 in Appendix A. According to the group-mean and panel test statistics, we can strongly reject the null of no cointegration. Thus, the presence of a long-run steady-state relationship between the variables

entering the ERPT equation is proved. Accordingly, we can obtain long-run ERPT coefficient from estimating equation (11) in levels.²

Figure 3: Comparison of inflation in the NMS before and after joining the EU



Note: Figures report inflation evolution before and after joining the EU. Source: Eurostat.

5 Empirical results

As explained above, before estimating our PSTR model, the first step is to test whether the regimes-switching effect is statistically significant, using linearity and no remaining nonlinearity tests. The testing procedure is as follows. Given a PSTR model, we test the null hypothesis that the model is linear. If the null is rejected, we estimate a two-regime PSTR model. Then, we test the null hypothesis of no remaining nonlinearity in this model. If it is rejected, estimate a three regime model. The testing procedure continues until the first acceptance of the null hypothesis of no remaining linearity. Results from this analysis are presented in Table 1. The linearity tests results show that the null hypothesis that model is linear is rejected for all three tests, *i.e.* Wald, Fisher and Likelihood Ratio tests, implying that the extent of pass-through respond nonlinearly to the inflation regime. Regarding no remaining nonlinearity, the results indicate that the null hypothesis cannot be

² The presence of cointegrating relationship is robust to the use of popular first generation tests (residual-based tests for panel cointegration) such developed by Pedroni (1999, 2004). To save space, we do not list the testing result; however, the result is available upon request.

rejected, implying that the model has only one threshold/two regimes. This implies that in our PSTR, there is only one threshold level of inflation which indicates the presence of two extreme regimes, *i.e.* low-inflation regime and high-inflation regime.

Table 1: Tests for linearity and no remaining nonlinearity of the PSTR model

Specification tests	Statistic	<i>p</i> -value
Linearity test		
Lagrange Multiplier - Wald (LM_W)	26.969	0.002
Lagrange Multiplier - Fisher (LM_F)	2.771	0.003
Likelihood ratio (LR)	28.586	0.000
No remaining nonlinearity test		
Lagrange Multiplier - Wald (LM_W)	12.832	0.233
Lagrange Multiplier - Fisher (LM_F)	1.183	0.305
Likelihood ratio (LR)	13.183	0.214

Note: For linearity test, H_0 : linear model; H_1 : PSTR model with at least one threshold. For no remaining nonlinearity (test for the number of regimes), H_0 : PSTR with one threshold; H_1 : PSTR with at least two thresholds.

Next, Table 2 shows the estimated results of our PSTR pass-through equation (11) using Nonlinear Least Squares (NLS). We report long-run pass-through coefficient for the two extremes regimes, *i.e.* $g(q_{it}; \gamma, c) = 0$ and $g(q_{it}; \gamma, c) = 1$ as defined in (12). We compute sum of squared residuals ratio (SSR_{ratio}) between PSTR model and the linear specification which suggests a better fit for the nonlinear model. Following González et al. (2005), we also check the quality of the estimated PSTR specification by conducting the misspecification test of parameters constancy, in addition to non remaining nonlinearity.

From Table 2, we observe the presence of significant threshold inflation rate level in our panel of new EU countries, with $c = 4.56\%$. As expected, this an evidence of the existence of two inflation regimes in the ERPT mechanism: low-inflation environment when CPI inflation is below the threshold of $c = 4.56\%$ ($\pi_{it} < c$); and high-inflation environment when CPI inflation surpasses the value of 4.56% ($\pi_{it} > c$). This is consistent with our empirical priors that pass-through mechanisms may be different whether inflation rate is above or below a given threshold. Regarding the long-run ERPT, our results suggest a significant regime-dependence of the pass-through to inflation levels. We find that the sensitivity of import prices to exchange rate is about 0.78% when when CPI inflation is below $c = 4.56\%$, but beyond this threshold level, the degree of ERPT becomes higher and reaches a full ERPT (LR ERPT= 1.018%). To provide a statistical significance of our results, we display the point estimates of

long-run ERPT with 95% confidence intervals over the two regimes (see values between square brackets in the last row of Table 2). We see that the increase is robustly significant, with the rates of pass-through are strongly different between the two inflation regimes. Thereafter, we plot the estimated logistic transition functions and the ERPT as a function of the transition variable $q_{it} = \pi_{it}$ (see Figure). This sheds more light on the regime-dependence of ERPT to inflation environment. The positive connection between the degree of the ERPT and inflation is quite clear in our panel of the NMS.

Table 2: Estimation results of PSTR pass-through equation for the NMS over 1996:1-2013:4

$$\pi_t = \underset{(0,000)}{0.779} e_{it} + \underset{(0,000)}{0.190} y_{it} + \underset{(0,000)}{0.580} w_{it}^* + \underset{(0,000)}{0.238} e_{it} \times g(q_{it}; \hat{\gamma}, \hat{c}) + \hat{\varepsilon}_t$$

$$g(q_{it}; \hat{\gamma}, \hat{c}) = \left(1 + \exp \left\{ \underset{(0,07)}{-57.852} (\pi_{it} - \underset{(0,012)}{4.566}) \right\} \right)^{-1}$$

$$NT = 560, \quad R^2 = 0.839, \quad SSR_{ratio} = 0.623, \quad pLM_F^C = 0.305, \quad pLM_F^N = 0.259$$

ERPT Elasticities

Low-inflation regime: $g = 0$	High-inflation regime: $g = 1$
LR ERPT= 0.779 (0,000) [0.723 0.843]	LR ERPT= 1.018 (0,000) [0.937 1.099]

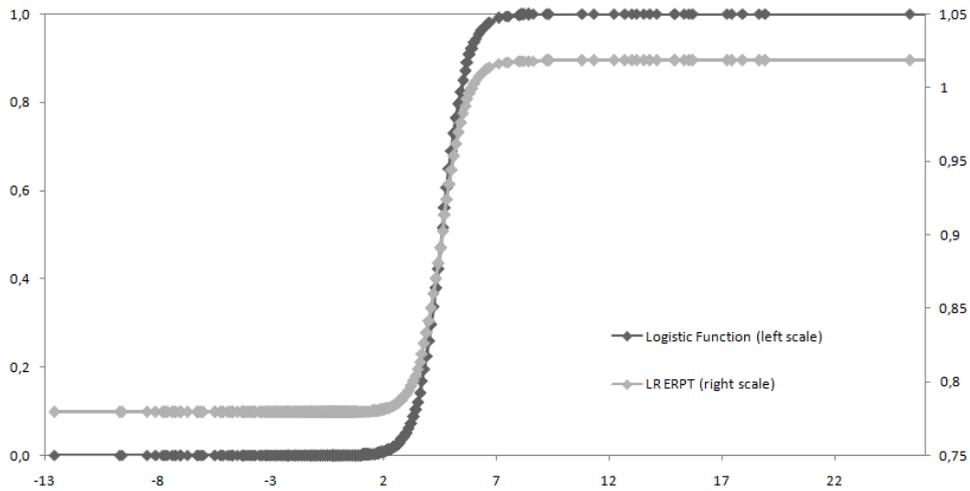
Note: In this table are reported the estimation results of the PSTR pass-through equation (11) using Nonlinear Least Squares (NLS). LR ERPT are long-run pass-through elasticities for low- and high-inflation regimes. Numbers in parentheses are p -values of estimates. For ERPT elasticities, we also report 95% confidence intervals between square brackets. R^2 denotes the coefficient of determination and SSR_{ratio} is the ratio of sum of squared residuals between PSTR model and the linear specification. Also, misspecification tests are conducted using the F -version of the LM-type test: pLM_F^C is the p -values of the LM test of parameter constancy and pLM_F^N is the p -values of the LM test of no remaining nonlinearity.

Broadly speaking, our results are in line with Taylor's hypothesis, i.e. the responsiveness of prices to exchange rate fluctuations depends positively on inflation environment. The intuition behind this phenomenon may be due to the foreign firms' behavior. The latter are more willing to set their prices in the currency of importing countries where inflation environment is stable (*local-currency pricing* or *LCP* setting). In such case ERPT would be lower. However, when exporters perceive a higher inflation level, they may shift away from local-currency pricing by passing exchange rate changes through the

prices in importer's currency (*producer-currency pricing* or *PCP* strategy). This behavior would entail a higher degree of pass-through.

Furthermore, our results corroborate those obtained for the NMS in the sense that falling inflation rates are closely linked to the decline in pass-through estimates since the late 1990s, although ERPT in central and eastern Europe remains on average larger than for industrial countries (Bitans, 2004). Nevertheless, as mentioned above, previous studies on the new EU member states have suggested a high degree of sensitivity of results which are mainly due to differences in empirical methodologies and data periods. For instance, María-Dolores (2010) found long-run ERPT coefficient roughly 0.63% as the average a sample of 9 NMS EU countries. The hypothesis of *PCP* (complete pass-through) is clearly rejected for all the countries except for Slovenia and Cyprus. Similarly, in a panel of 10 new EU member states, Jimborean (2013) reported a more pronounced degree of pass-through to import prices at about 0.90-1.13%. However, the major drawback of the latter study is that the time-series properties of the data, *i.e.* non-stationarity and cointegration issues in a panel-data framework, are neglected.³

Figure 4: Estimated transition function and long-run ERPT



Source: Personal calculation. Note: Estimated transition functions and long-run ERPT as function of inflation level. Results are from PSTR equation (11) with $q_{it} = \pi_{it}$.

Another important drawback of the previous studies is the failure to account for policy shifts could result in biased and underestimated ERPT

³ Jimborean (2013) computed the long-run ERPT using the lagged adjustment of import prices (as an explanatory variable) in the pass-through equation.

estimates. This is of notable importance due to the great diversity in exchange rate regimes and inflation environment the central and eastern European countries. Although previous empirical works have put forth the role of macroeconomic factors in influencing the exchange rate transmission, they fail to recognize the nonlinear behaviour the pass-through mechanism. Given the fact that most of the NMS, on one hand, commuted from a high in inflation environment to a relatively moderate inflation rate by the late 90s, and on the other hand, have adopted different exchange systems over time, it is crucial to take into account these changing macroeconomic conditions by implementing relevant econometric approaches.

6 Conclusion

This paper studies the existence of nonlinear behaviour in exchange rate pass-through (ERPT) for a set of new EU Member States (NMS). To account for nonlinearities, we implement a Panel Smooth Transition Regression (PSTR) approach, where the extent of pass-through is allowed to respond nonlinearly to the inflation environment. The main advantage of PSTR framework is the use of a grid search to select endogenously threshold level(s) allowing to identify different inflation regimes. Using quarterly data for ten NMS over 1996-2013, our results suggest a significant regime-dependence of the pass-through to inflation levels. We find that the sensitivity of import prices to exchange rate is about 0.78% when when CPI inflation is below $c = 4.56\%$, but beyond this threshold level, the degree of ERPT becomes higher and reaches a full ERPT (LR ERPT= 1.018%). Furthermore, our findings corroborate those obtained for the NMS in the sense that falling inflation rates are closely linked to the decline in pass-through estimates since the late 1990s, although ERPT in the NMS remains on average larger than for industrial countries. These results contribute to lively debate on the path towards the Euro and how ECB could reduce the inflation differentials with the enlargement of the monetary union.

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Appendix A. Panel unit root and cointegration Tests

Table 3: Pesaran's (2007) test for panel unit root

Variables	Level		First difference	
	Intercept	Intercept & trend	Intercept	Intercept & trend
p_{it}	-0.179 (0.359)	-0.538 (0.367)	-7.334 (0.000)	-5.440 (0.000)
e_{it}	-0.631 (0.176)	-0.954 (0.283)	-8.226 (0.000)	-9.024 (0.000)
y_{it}	-0.758 (0.217)	-0.843 (0.256)	-7.887 (0.000)	-8.501 (0.000)
w_{it}^*	-0.521 (0.298)	-0.726 (0.267)	-6.281 (0.000)	-6.933 (0.000)

Note: p -values for the null hypothesis of non stationarity are reported between parentheses. Also, the empirical statistics can be compared to the critical value from Pesaran (2007) which are -2.15 for specification with an intercept and -2.65 for specification with intercept and linear time trend, at 5% level. Individual lag lengths are based on Akaike Information Criteria (AIC).

Table 4: Westerlund's (2007) Panel Cointegration Test

Statistic	Value	p -value	Robust p -value
<i>Group-mean statistics</i>			
G_τ	-9.386	0.000	0.016
G_α	-13.399	0.000	0.000
<i>Panel statistics</i>			
P_τ	-18.641	0.000	0.004
P_α	-23.109	0.000	0.000

Note: G_τ and G_α are group mean statistics that test the null of no cointegration for the whole panel against the alternative of cointegration for some countries in the panel. P_τ and P_α are the panel statistics that test the null of no cointegration against the alternative of cointegration for the panel as a whole. Optimal lag and lead lengths are determined by Akaike Information Criterion (AIC). In the last column, we present the bootstrapped p -values which are robust against cross-sectional dependencies. Number of bootstraps is set to 800..