Narodowy Bank Polski

NBP Working Paper No. 163

# The Balassa-Samuelson effect and the channels of its absorption in the Central and Eastern European Countries

Karolina Konopczak

NBP Working Paper No. 163

# The Balassa-Samuelson effect and the channels of its absorption in the Central and Eastern European Countries

Karolina Konopczak

Economic Institute
Warsaw, 2013

Karolina Konopczak – Warsaw School of Economics and Institute for Market, Consumption and Business Cycles Research, e-mail: karolina.konopczak@ises.edu.pl

This research project was conducted under the NBP Economic Research Committee's open competition for research projects to be carried out by the NBP staff and economists from outside NBP and was financed by Narodowy Bank Polski.

Print:
NBP Printshop

Published by:
Narodowy Bank Polski
Education & Publishing Department
ul. Świętokrzyska 11/21
00-919 Warszawa, Poland
phone +48 22 653 23 35
www.nbp.pl

ISSN 2084-624X

# Contents

**Abstract**

The aim of the study is to estimate the magnitude of the Balassa-Samuleson effect as well as the effectiveness of the labour and the product market in its absorption in Poland, the Czech Republic, Hungary and Slovakia. The obtained results allowed to determine the magnitude of the systematic component of inflation differentials relative to the euro area, hence to assess the risk of common monetary policy inadequacy with respect to these economies. The obtained estimates suggest that the catching-up driven inflationary pressure is a non-negligible issue in the context of the CEECs integration with the euro area, since the systematic inflation differentials were comparable in size to those experienced by the so-called peripheral member states in the first decade after the introduction of the euro. Moreover, in the case of Poland none of the potential absorption mechanisms of the Balassa-Samuelson effect seemed to mitigate the convergence-induced inflationary pressure over the sample period. The outcomes suggest that ignoring the non-fulfilment of theoretical model assumptions regarding wages and markups, which is common in the literature, distorts estimation results.



**JEL Classification**: F41, E31, C33

**Keywords**: Balassa-Samuelson hypothesis, monetary integration, real convergence, panel cointegration

# Introduction

Having joined the European Union in May 2004, the Central and Eastern European Countries (henceforth: CEECs) became member states with a derogation concerning the adoption of the single currency, i.e. they are obliged to join the third stage of the Economic and Monetary Union as soon as the convergence criteria stipulated in the Treaty on the Functioning of the EU are fulfilled. The inevitable corollary of the accession to the euro area is the resignation from autonomous monetary and exchange rate policy, both being major shock-absorbing instruments. In conducting common monetary policy, the European Central Bank takes into account average economic performance in the euro area, which may be sub-optimal from the perspective of an individual member state.

The inadequacy of common monetary policy arises from the heterogeneity of inflation within the monetary union. In the presence of persistent inflation differentials, the equality of nominal interest rates implies heterogeneity of real interest rates (leaving aside differences in the process of inflation expectations formation). Lower than average real interest rate in one country (in particular below the natural interest rate) boosts its economic activity relative to other member states and consequently the inflation rate (its cyclical component), hence amplifying divergences. This mechanism, lying at the heart of the so-called Walters critique of monetary unions (Walters, 1994), is thereby of procyclical nature and leads to the built-up of macroeconomic imbalances, both internal (the boom-bust cycles) and external (undermining of competitiveness and the accompanying current account deterioration). The resolution of the imbalances by means of nominal devaluation is not a feasible option within the monetary union, since nominal exchange rates are irrevocably fixed. The burden of adjustment rests, therefore, upon the real exchange rate channel, whose effectiveness depends on the elasticity of the domestic product and labour markets, i.e. the elasticity of prices and wages.

The functioning of the euro area hitherto suggests that, contrary to beliefs widely held prior to its launching, the effectiveness of the real exchange rate channel among some member states is very limited. Inflation differentials within the euro area, aggravated by the procyclical real interest rate mechanism, contributed to the built-up of considerable imbalances in some countries. The problem concerned primarily the so-called peripheral member states, namely Greece, Portugal, Spain and Ireland. As the result of the inflation rate being systematically above the average for the euro area (Table 1), in the years 1999 through 2007 these economies experienced a considerable deterioration of price competitiveness. Consequently, they developed persistently high current account deficits, which, in turn, accumulated into considerable stocks of net external liabilities.

Table 1: Inflation differentials between peripheral member states and the average for the euro area

| Country | 1999-2007 | 2008-2011 |
|---|---|---|
| Greece | 1.2 p.p | 1.5 p.p |
| Spain | 1.8 p.p | −0.1 p.p |
| Ireland | 2.0 p.p | −3.5 p.p |
| Portugal | 1.1 p.p | −0.1 p.p |

Note: Inflation differentials were computed on the basis of value-added deflators excluding agriculture and fishing (NACE rev.1.1. sections A and B).

These developments made the peripheral member states vulnerable to macroeconomic risk, which materialised during the recent financial crisis in various forms, i.a. the collapse of the real estate market (in Spain and Ireland), the sovereign debt crisis (in Greece, Spain and Portugal) and the banking crisis (in Ireland). What is more, despite a sharp and prolonged economic downturn which they experienced (cumulative output loss from 2008 to 2011 of 10% in Ireland, 14% in Greece and 3% in Spain and Portugal, the average for the euro area being 1%), a sizable internal devaluation has taken place only in the case of Ireland (Table 1), implying that, due to product and labour market rigidities, regaining competitiveness while in a currency union may be very difficult for the other countries.

The problems experienced by the peripheral member states suggest that inadequacy of common monetary policy may turn out to be a serious cost of joining a currency union. Therefore, a risk assessment of such inadequacy should become a major element of cost-benefit analysis of euro adoption for the prospective member states. The aim of the paper is to assess that risk with regard to four CEECs – Poland (PL), the Czech Republic( CZ), Hungary (HU) and Slovakia (SK). Due to considerably lower GDP per capita than in the euro area (in 2011 GDP per capita in Purchasing Power Parity amounted to 59% of the average for the euro area in the case of Poland, 73% in the Czech Republic, 60% in Hungary and 68% in Slovakia) it may be assumed that the inadequacy of ECB monetary policy with regard to the CEECs may be of structural nature, i.e. it may be determined by the catching-up process. The analysis focuses, therefore, on the inflationary pressure stemming from the real convergence process, i.e. the so-called Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964).

The rest of the paper is organised as follows. Section 1 derives the baseline Balassa-Samuelson model, as well as its extended version, taking into account possible invalidity of baseline model assumptions regarding the product and labour markets. In Section 2 we outline the empirical framework employed in the paper, i.a. a novel approach to estimating the Balassa-Samuelson effect aimed at overcoming the problem of relatively short time span and low frequency of series available for the CEECs. Next, we give a brief overview of panel econometrics methods, on the basis of which we choose the most appropriate estimation techniques. In Section 4 empirical results are presented. These include estimation results of the baseline and extended models, the results of tests of effectiveness of labour and product market in absorbing catching-up driven inflationary pressure, as well as the quantification of the Balassa-Samuelson effect. Next, on the basis of the obtained results conclusions are drawn concerning the risk of the ECB policy inadequacy with regard to the analysed CEECs.

# 1 Methodological issues

## 1.1 Theoretical framework

### 1.1.1 Baseline Balassa-Samuelson model

The cross-country inflation differentials may be the result of several factors, both of structural and transitory nature, the most important being: (1) the catching-up process, (2) phase shift in the business cycle, (3) asymmetric monetary and fiscal policy shocks, (4) changes in regulated prices and VAT or excise rates, (5) differences in the import-to-production ratio, and (6) relative endowment in production inputs. From the perspective of the CEECs in their run-up to the euro area the crucial factor are the inflation differentials adjusted for transitory effects, i.e. their structural component, which is determined by the catching-up process, called the Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964).

A starting point for the derivation of the Balassa-Samuelson model is an assumption of an economy consisting of two sectors – the tradable (T) and the non-tradable (NT) one – differing with respect to productivity growth (the tradable sector is the sole source of productivity growth or at least the growth rate is higher in the tradable than in the non-tradable one). Perfect competition in both sectors and perfect labour force mobility between sectors is also assumed. Labour is, however, immobile between countries.

Output in both sectors, home and abroad, is determined by the Cobb-Douglas production function with constant returns to scale:

$$Y_j(K_j, L_j) = A_j L_j^{\alpha_j} K_j^{1-\alpha_j}, \tag{1}$$

with $Y_j$ denoting output, $A_j$ – total factor productivity, $L_j$ – labour input $K_j$ – capital input in sector $j$ ($j\epsilon\{T, N\}$). $\alpha_j$ and $1-\alpha_j$ denote labour and capital elasticities of output, respectively.

In both sectors, producers maximise their profits by choosing the appropriate inputs of production factors:

$$\max_{L_j, K_j} \quad (P_j A_j K_j^{1-\alpha_j} L_j^{\alpha_j} - W_j L_j - RK_j) \tag{2}$$

where $W_j$ ($j\epsilon\{T, N\}$) stands for wages in respective sectors, $R$ is the cost of capital and $P_j$ ($j\epsilon\{T, N\}$) denotes price level of goods produced in respective sectors. First order conditions for the above maximisation problems (with respect to labour) are the following:

$$\alpha_j P_j A_j K_j^{1-\alpha_j} L_j^{\alpha_j - 1} - W_j = 0 \;, \quad j\epsilon\{T, N\}. \tag{3}$$

This implies the price of labour as follows:

$$W_j = \alpha_j P_j A_j (\frac{L_j}{K_j})^{\alpha_j - 1}, \quad j\epsilon\{T, N\}. \tag{4}$$

Perfect competition assumption implies price of production factors equal to their marginal revenue product. Therefore, in the long run an increase in labour productivity translates into a proportional increase in wages. On the other hand, perfect labour mobility assumption implies cross-sectoral equality of wages in the long run. Otherwise, the employees would be encouraged to change the sector until growing labour supply in the sector with higher earnings and falling labour supply in the other sector would level the wages in the entire economy.

Assuming wage homogeneity across the sectors, $W_T = W_N \equiv W$, the relationship between relative prices and relative productivity between the tradable and non-tradable sector can be derived:

$$\frac{P_N}{P_T} = \frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}}. \tag{5}$$

Dividing (5) by its foreign counterpart (foreign counterparts to domestic variables are denoted with an asterisk in superscripts) leads to the following relationship expressing relative price of non-tradable goods in international comparison as a function of relative productivities home and abroad, i.e. to the baseline Balassa-Samuelson model:

$$\frac{P_N}{P_T} / \frac{P_N^*}{P_T^*} = \frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}} / \frac{\alpha_T^* \frac{Y_T^*}{L_T^*}}{\alpha_N^* \frac{Y_N^*}{L_N^*}}. \tag{6}$$

### 1.1.2 Extended Balassa-Samuelson model

The baseline Balassa-Samuelson model implies that a differential productivity growth fully translates into inflation differentials between a catching-up and an advanced economy. This is asserted by a combined fulfilment of the following conditions: (1) an increase in productivity in the tradable sector fully translates into an increase in wages (perfect competition assumption implying price of production factors equal to their marginal revenue product), (2) an increase in wages in the tradable sector fully translates into an increase in wages in the non-tradable sector (perfect labour mobility assumption), (3) an increase in wages in the non-tradable sector which does not match productivity growth fully translates into an increase in prices (perfect competition assumption implying markups equal to zero). Implicitly, the model assumes also (4) identical trajectories of non-wage costs of production, such as raw materials, energy or external services.

The incomplete pass-through in the Balassa-Samuelson model may be the result of non-fulfilment of any of the above conditions. In particular, in the case of existing differences in raw material-, energy- or import-to-production ratio between the tradable and the non-tradable sector, production costs and, consequently, prices are differently affected by the developments of raw materials and energy costs, as well as the exchange rate. Therefore, the incomplete pass-through may be the consequence of random trajectories of these variables. It may, however, result from factors of systematic nature, reflecting product and labour market characteristics. Namely, labour market features can mitigate levelling of wages between sectors. On the other hand, product market characteristics may reduce the impact of wage increases on prices in the non-tradable sector. Thereby the labour and the product market may to some extent absorb inflationary pressure stemming from the real convergence process.

The effectiveness of the labour market in absorbing the Balassa-Samuelson effect depends on the magnitude of wage increase pass-through from the tradable to the non-tradable sector. Equation (6) was derived under the assumption that wages are equal across sectors. Should this assumption be rejected, the baseline Balassa-Samuelson model ought to be extended to accommodate the relative price trajectory:

$$\frac{P_N}{P_T} / \frac{P_N^*}{P_T^*} = \left( \frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}} / \frac{\alpha_T^* \frac{Y_T^*}{L_T^*}}{\alpha_N^* \frac{Y_N^*}{L_N^*}} \right) \cdot \left( \frac{W_N}{W_T} / \frac{W_N^*}{W_T^*} \right) \tag{7}$$

The ability of the product market to mitigate catching-up inflationary pressure hinges upon the extent to which fluctuations of production costs may be absorbed by markup adjustments. Perfect competition assumption implies markups in the economy equal to zero and, consequently, a complete pass-through of unit labour costs on prices. If, however, prices are set as a mark-up on production costs:

$$P_j = MC_j M_j, \tag{8}$$

where $MC_j$ stands for marginal costs, and $M_j$ for a markup in respective sectors $(j \epsilon \{T, N\})$, part of the unit labour costs increase may be absorbed by decreasing the profit margin.

Assuming identical trajectories of non-labour production costs in the tradable and non-tradable sector and proportionality of marginal and average costs, equation (7) can be further extended to accommodate the relative markups trajectory:

$$\frac{P_N}{P_T} / \frac{P_N^*}{P_T^*} = \left( \frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}} / \frac{\alpha_T^* \frac{Y_T^*}{L_T^*}}{\alpha_N^* \frac{Y_N^*}{L_N^*}} \right) \cdot \left( \frac{W_N}{W_T} / \frac{W_N^*}{W_T^*} \right) \cdot \left( \frac{M_N}{M_T} / \frac{M_N^*}{M_T^*} \right). \tag{9}$$

## 1.2 Empirical framework

The data used in the analysis come from the Eurostat database. The sample covers years 1995 through 2010 and is of annual frequency. The source variables comprise sectoral (according to NACE rev. 1.1 classification) value added deflators as an approximation for price developments, sectoral labour productivity (value added over total employment in each sector), sectoral wages (compensation of employees over total employment in each sector) and an approximation for sectoral markups (nominal output over variable costs, i.e. the sum of compensation of employees and intermediate inputs). The application of NACE-based statistical concepts (value-added deflators instead of price indices) asserts the coherence of sectoral classification. Table 2 contains definitions of variables used in the empirical analysis.

An important issue when estimating the Balassa-Samuelson effect is the classification of sectors as tradable or non-tradable. The sectoral classification applied in this study (Table 3) achieves two goals: firstly, it is in line with the main strand of the literature; secondly, it maximises the cross-sectional dimension of the panel, which enhances the effectiveness of the estimation. The only sub-sector we excluded from the analysis is agriculture and fishing (NACE rev.1.1. sections A and B). Although the products of these sub-sectors are subject to international trade, both their prices and quantities are heavily distorted by administrative interventions (on both country- and the EU-level) and random events, such as weather conditions.

Table 3: Composition of the tradable and non-tradable sector

| Sectoral classification according to NACE rev. 1.1 | |
|---|---|
| Tradable sector | Non-tradable sector |
| Manufacturing (D) | Mining and quarrying (C)<br>Electricity, gas and water supply (E)<br>Construction (F)<br>Wholesale and retail trade; repair of motor vehicles, motorcycles, personal and household goods (G)<br>Hotels and restaurants (H)<br>Transport, storage and communication (I)<br>Financial intermediation (J)<br>Real estate, renting and business activities (K)<br>Public administration and defence; compulsory social security (L)<br>Education (M)<br>Health and social work (N)<br>Other community, social, personal service activities (O) |

Table 2: Definitions of variables

| Variable | Description |
|---|---|
| $p_i^K$ | logarithm of value-added deflator in sector $i$ of country $K$ |
| $\tilde{l}_i^K$ | logarithm of labour productivity (value added over total employment) in sector $i$ of country $K$ |
| $w_i^K$ | logarithm of wage (compensation of employees over total employment) in sector $i$ of country $K$ |
| $m_i^K$ | logarithm of markup (nominal output over variable costs, i.e. the sum of compensation of employees and intermediate inputs) in sector $i$ of country $K$ |
| $p_{diff_j}^{CEEC\_EA} \equiv (p_{N_j}^{CEEC} - p_T^{CEEC}) - (p_{N_j}^{EA} - p_T^{EA})$ | difference between differential price levels (non-tradables vs. tradables) in a CEEC and the euro area |
| $\tilde{l}_{diff_j}^{CEEC\_EA} \equiv (\tilde{l}_T^{CEEC} - \tilde{l}_{N_j}^{CEEC}) - (\tilde{l}_T^{EA} - \tilde{l}_{N_j}^{EA})$ | difference between differential productivity levels (tradables vs. non-tradables) in a CEEC and the euro area |
| $w_{diff_j}^{CEEC\_EA} \equiv (w_{N_j}^{CEEC} - w_T^{CEEC}) - (w_{N_j}^{EA} - w_T^{EA})$ | difference between differential wage levels (non-tradables vs. tradables) in a CEEC and the euro area |
| $m_{diff_j}^{CEEC\_EA} \equiv (m_{N_j}^{CEEC} - m_T^{CEEC}) - (m_{N_j}^{EA} - m_T^{EA})$ | difference between differential markup levels (non-tradables vs. tradables) in a CEEC and the euro area |

The Balassa-Samuelson effect is a long-run phenomenon and, therefore, necessitates the use of cointegration analysis for the purpose of estimation. In the literature the model is estimated via either time series methods applied to a single economy, or panel econometric techniques applied to a group of economies. Both approaches have obvious drawbacks when applied to the CEECs. Accessible time series for these countries, mostly dating back to 1995, are far too short to ensure efficient estimation of long-run relationships. Moreover, sectoral disaggregated data available for the CEECs are of annual frequency, which makes the application of cointegration analysis virtually impossible. In previous studies the problem was addressed either by estimating the model in first differences (e.g. Cipriani, 2001), or by extending the sample to a group of CEECs which, under the assumption of their relative homogeneity, allowed for the application of panel methods (e.g. Égert et al., 2002; Wagner and Hlouskova, 2004; Wagner, 2005). Due to the fact that the mechanism outlined by Balassa and Samuelson is of a long-run nature, the former solution potentially leads to the underestimation of the effect, as only its short-run component can be captured. Turning to a multi-country panel, on the other hand, alters the interpretation of the results and hence it is not possible to develop policy recommendations for a single country. What is more, due to limited cross-sectional dimension of the panel (such analyses comprise usually only a few economies) the application of some macroeconometric panel techniques may result in biased estimates.

Our attempt to overcome this difficulty consists in designing the panel in a different manner (see Konopczak and Torój, 2010). Namely, its cross-sectional dimension is obtained by means of sectoral disaggregation, i.e. the price, productivity, wage and markup differentials are computed as a difference between the tradable sector (NACE rev.1.1. section D) and non-tradable subsectors (NACE rev 1.1. sections C, E, F, G, H, I, J, K, L, M, N, O). Thus, the relation between the tradable sector and various sub-sectors of the non-tradable sector serves as the unit dimension. This allows to concentrate on the single economy perspective and, at the same time, improve the efficiency of the estimation by applying panel econometric techniques.

# 2 Econometric issues

The inference on the stationarity of the analysed series and the existence of the long-run equilibrium among them is based on the results of both first and second generation tests. On the one hand, the former tests are widely-spread in the literature and, therefore, their application makes it possible to compare the obtained results with the previous studies. On the other hand, they assume cross-sectional independence. If this assumption is violated, however, first generation tests suffer from serious size distortions, which results in over-rejection of non-stationarity null (Banerjee et al., 2005). Owing to strong inter-economy linkages, the risk of cross-correlation in our data is non-negligible. For this reason we additionally apply second generation tests allowing for cross-sectional dependence.

## 2.1 Panel unit root tests

From the available unit root tests we choose those that allow for heterogeneity across the cross-sectional dimension in terms of the autoregressive coefficient, since according to Monte Carlo simulations (Im et al., 2003) such tests yield higher power compared to those based on the supposition of common persistence parameter across cross-sectional units.

According to the null hypothesis of all the applied tests, all cross-sectional processes contain a unit root:

$$H_0 : \alpha_i - 1 = 0, \tag{10}$$

where $\alpha_i$ denotes the persistence parameter, while the alternative hypothesis is given by

$$H_1 : \alpha_i - 1 < 0 \tag{11}$$

for at least one $i$ ($i$=1,...,$N$), where N is the number of cross-sectional units. The alternative hypothesis may be interpreted as a non-zero fraction of the processes being stationary.

From the first generation panel unit root tests we choose the Im, Pesaran and Shin test (IPS, Im et al., 2003) and Fisher-type Dickey-Fuller test (Maddala and Wu, 1999). The Im, Pesaran and Shin statistics is obtained by a two-step procedure. In the first step a separate ADF regression is estimated for each cross-sectional unit:

$$\Delta y_{i,t} = \alpha_{0i} + (\alpha_i - 1)y_{i,t-1} + \sum_{k=1}^{K_i} \beta_{ik}\Delta y_{i,t-k} + \varepsilon_{i,t}. \tag{12}$$

In the second step the average t-statistics for the persistence coefficient is computed:

$$IPS = N^{-1}\sum_{i=1}^{N} t_{(\alpha_i - 1)} \tag{13}$$

The standardised IPS statistics[1] has an asymptotic (with $N \rightarrow \infty$) standard normal distribution.

Maddala and Wu (1999) propose an alternative test based on the Fisher's (1932) method of combining significance levels of the independent tests with the same set of hypotheses. The test statistics is given by:

$$-2\sum_{i=1}^{N} log(p_i), \tag{14}$$

where $p_i$ denotes the p-value from the individual ADF test. In the limit (with $T \rightarrow \infty$) , the test has a $\chi^2_{2N}$ distribution.

From the second generation tests we choose the cross-sectionally augmented IPS (CIPS) test as proposed by Pesaran (2003), which assumes that the cross-correlation can be ascribed to a single unobserved common factor ($f_t$):

$$\begin{gathered}\Delta y_{i,t} = \alpha_{0i} + (\alpha_i - 1)y_{i,t-1} + \textstyle\sum_{k=1}^{K_i} \beta_{ik}\Delta y_{i,t-k} + u_{i,t} \\ u_{i,t} = \lambda_i f_t + \varepsilon_{i,t}.\end{gathered} \tag{15}$$

Pesaran proposes to proxy the common factor by cross-sectional mean of levels and differences of the series of interest and their lagged values. The cross-sectionally augmented ADF regression takes the following form:

$$\Delta y_{i,t} = \alpha_{0i} + (\alpha_i - 1)y_{i,t-1} + \gamma_i \bar{y}_{t-1} + \delta_i \Delta\bar{y}_t + \sum_{k=1}^{K_i} \beta_{ik}\Delta y_{i,t-k} + \varepsilon_{i,t}, \tag{16}$$

where $\bar{y}_t = N^{-1}\sum_{i=1}^{N} y_{i,t}$, and $\Delta\bar{y}_t = N^{-1}\sum_{i=1}^{N} \Delta y_{i,t}$.

Like in the case of the IPS test, the CIPS statistics is obtained by cross-sectionally averaging the t-statistics for the persistence coefficient:

$$CIPS = N^{-1}\sum_{i=1}^{N} t_{(\alpha_i - 1)}. \tag{17}$$

The CIPS statistics has a non-standard limiting distribution and the critical values are taken from Pesaran (2003).

Other second generation panel unit root tests proposed in the literature, e.g. by Bai and Ng (2004) or Moon and Perron (2004), allow for a multifactor error structure, which may better reflect the true data generating process. Those tests are, however, dedicated to panels with a considerable time and cross-sectional dimension of the panel, which is not the case in our empirical analysis.

## 2.2 Panel cointegration tests

We test for the presence of cointegration by means of three panel cointegration tests − two Engle-Granger-based tests, the Pedroni test (Pedroni, 2004) and the Westerlund test (Westerlund, 2007), and the Johansen-type Fisher test (Maddala and Wu, 1999).

The Pedroni test is residual-based, in that it consists in applying a unit root test to the residuals of the cointegration regression in order to verify the null hypothesis of no cointegration against the alternative hypothesis of cointegration in all cross-sectional units. There are seven test statistics available, four of which assume homogenous persistence parameters of the residuals series across the cross-sectional units (panel statistics) and three allow for heterogeneity in this respect (mean group statistics). Owing to the considerable risk of heterogeneity bias in the case of the analysed dataset we confine our attention to mean group statistics. In the case of those statistics the cointegration equation:

$$y_{i,t} = \alpha_{0i} + \beta_i x_{i,t} + \varepsilon_{i,t} \tag{18}$$

is estimated separately for each cross-sectional unit by means of the ordinary least squares. According to the results of Monte Carlo experiments (Pedroni, 2004), the group ADF statistics is the most powerful test for small temporal dimension of the panel (T inferior to 20), which is the case in our analysis. The ADF statistics also performs best in the presence of cross-correlation (Wagner and Hlouskova, 2010). For these reasons the statistical inference will be based on this statistics solely.

The group ADF statistics is computed on the basis of the estimates of the following equation:

$$\Delta\hat{\varepsilon}_{i,t} = \alpha_{0i} + (\alpha_i - 1)\hat{\varepsilon}_{i,t-1} + \sum_{k=1}^{K_i} \beta_{ik}\Delta\hat{\varepsilon}_{i,t-k} + \vartheta_{i,t}, \tag{19}$$

where $\hat{\varepsilon}_{i,t}$ denotes series of estimated residuals from the cointegration equation. The formula for the statistics is given by:

$$\tilde{Z}_{ADF} = \sum_{i=1}^{N}(\hat{s}_i^2 \sum_{t=1}^{T} \hat{\varepsilon}_{i,t-1}^2)^{-\frac{1}{2}} \sum_{t=1}^{T}(\hat{\varepsilon}_{i,t-1}\Delta\hat{\varepsilon}_{i,t}), \tag{20}$$

where $\hat{s}_i^2 = \frac{1}{T}\sum_{t=1}^{T}\hat{\vartheta}_{i,t}^2$. The standardized[2] test statistics has an asymptotic standard normal distribution.

Westerlund (2007) proposed an alternative approach to residual-based tests for panel cointegration. It consists in testing the significance of the error correction term $(y_{i,t-1} - \beta_i x_{i,t-1})$ within the error correction model:

$$\Delta y_{i,t} = \alpha_{0i} + \alpha_i(y_{i,t-1} - \beta_i x_{i,t-1}) + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + \epsilon_{i,t}. \tag{21}$$

The model is estimated by means of least squares in a reparameterised form:

$$\Delta y_{i,t} = \alpha_{0i} + \alpha_i y_{i,t-1} + \lambda_i x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + \epsilon_{i,t}, \tag{22}$$

where $\lambda_i = -\alpha_i \beta_i$. Four tests based on the estimates of $\alpha_i$ and its t-ratio are available, two of which are mean group statistics (the alternative hypothesis assumes that a non-zero fraction of units is cointegrated) and two are panel statistics (alternative hypothesis of cointegration in all cross-sectional units). According to Monte Carlo experiments (Westerlund, 2007) two of those statistics – mean group $\tau$ ($G_\tau$) and panel $\tau$ ($P_\tau$) – seem to outperform the other two (mean-group $\alpha$ and panel $\alpha$) in terms of power, size and robustness to cross-sectional dependence. For this reason we base our analysis on the results of these two tests.

In the case of mean group statistics equation (22) is estimated separately for each $i$ and the $G_\tau$ statistics is specified as follows:

$$G_\tau = N^{-1} \sum_{i=1}^{N} t_{\hat{\alpha}_i}. \tag{23}$$

The panel statistics is obtained by means of a three-step procedure. In the first step two regressions are run separately for each cross-sectional unit:

$$\Delta y_{i,t} = \alpha_{0i} + \lambda_i x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + \epsilon_{i,t} \tag{24}$$

and

$$y_{i,t} = \alpha_{0i} + \lambda_i x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + \epsilon_{i,t}. \tag{25}$$

Subsequently, the projection errors, $\Delta \tilde{y}_{i,t} = \Delta y_{i,t} - \Delta \hat{y}_{i,t}$ and $\tilde{y}_{i,t} = y_{i,t} - \hat{y}_{i,t}$, are computed and on their basis the common error correction parameter, $\alpha$, is estimated. The $P_\tau$ statistics is the t-ratio for $\hat{\alpha}$.

Applying the Fisher's (1932) method of combining p-values of independent tests, Maddala and Wu (1999) proposed a test based on Johansen's trace and maximum eigenvalue statistics:

$$-2 \sum_{i=1}^{N} log(p_i) \rightarrow \chi^2_{2,N}. \tag{26}$$

## 2.3 Estimation of the cointegration vectors

As proven by Kao and Chiang (2000) the least squares estimator is inconsistent when applied to cointegrated panel variables. For this reason the cointegration vectors of the long-run relationships are estimated by means of the fully-modified ordinary least squares (FMOLS) proposed by Phillips and Moon (1999), building upon Phillips and Hansen (1990) and the dynamic ordinary least squares estimators (DOLS) proposed by Kao and Chiang (2000), basing on Saikkonen (1991). Both estimators are asymptotically efficient and allow for serial correlation and endogeneity of regressors in the cointegration equation. In the limit both estimators are equivalent (Banerjee, 1999).

The FMOLS estimator involves a two-step procedure. In the first stage the long-run covariance is estimated on the basis on the OLS-regression estimates and subsequently the OLS estimator is corrected by factors derived in the first step. Let us consider the following panel system:

$$\begin{cases} y_{it} = \alpha_i + \beta x_{it} + \mu_{it} \\ x_{it} = x_{it-1} + \varepsilon_{it} \end{cases} \quad i = 1, ..., N \tag{27}$$

Vector error process $\xi_{it} = [\mu_{it}, \varepsilon_{it}]^T$ is stationary, which is equivalent to cointegration of the analysed variables.

We denote by $\Omega_i = \begin{bmatrix} \Omega_\mu & \Omega_{\mu\varepsilon} \\ \Omega_{\varepsilon\mu} & \Omega_\varepsilon \end{bmatrix}$ the long-run covariance matrix of the error process, i.e.

$$\Omega_i = \sum_{k=-\infty}^{\infty} \Gamma_i^k = \Gamma_i^0 + \sum_{k=1}^{\infty} (\Gamma_i^k + \Gamma_i^{k\,T}), \tag{28}$$

where $\Gamma_i^k = E(\xi_i^k \xi_i^{0\,T})$ is the autocovariance matrix of order $k$. The consistent estimator of long-run covariance matrix is given by:

$$\hat{\Omega}_i = \hat{\Gamma^0}_i + \hat{\Gamma}_i + \hat{\Gamma}_i^T, \tag{29}$$

where $\hat{\Gamma}_i$ is a weighted sum of estimated autocovariances obtained by means of kernel estimation. The estimated matrix may be Cholesky-decomposed:

$$\hat{\Omega}_i = \hat{L}_i \hat{L}_i^T, \tag{30}$$

where $\hat{L}_i = \begin{bmatrix} \hat{L}_{11i} & 0 \\ \hat{L}_{21i} & \hat{L}_{22i} \end{bmatrix}$ is the lower triangular decomposition of $\hat{\Omega}_i$, normalised so that $\hat{L}_{22i} = \hat{\Omega}_{22i}^{-\frac{1}{2}}$.

The endogeneity correction is achieved by means of the following transformation:

$$y_{it}^{*} = y_{it} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} \Delta x_{it}, \tag{31}$$

while the serial correlation correction term is given by the following formula:

$$\hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Gamma^0}_{21i} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} (\hat{\Gamma}_{22i} + \hat{\Gamma^0}_{21i}) \tag{32}$$

The correction terms are applied to the OLS estimator in the following manner:

$$\hat{\beta}^{FMOLS} = \frac{1}{N} \sum_{i=1}^{N} (\sum_{t=1}^{T} (x_{it} - \bar{x}_i)^2)^{-1} (\sum_{t=1}^{T} (x_{it} - \bar{x}_i) y_{it}^{*} - T\hat{\gamma}_i) \tag{33}$$

and the t-statistics for $\hat{\beta}$ has an asymptotical standard normal distribution.

The DOLS estimator, on the other hand, corrects for the endogeneity problem by augmenting the regression with leads and lags of first difference of independent variables. The estimation equation has the following specification:

$$y_{it} = \alpha + \beta x_{it} + \sum_{p=-P}^{P} \delta_p \Delta x_{it-p} + u_i + \varepsilon_{it} \tag{34}$$

The estimator is given as:

$$\hat{\beta}^{DOLS} = \frac{1}{n} \sum_{i=1}^{n} (\sum_{t=1}^{T} z_{it} z_{it}^{'})^{-1} (\sum_{t=1}^{T} z_{it} y_{it}), \tag{35}$$

where $z_{it} = (x_{it} - \overline{x_i}, \Delta x_{it-P}, ..., \Delta x_{it+P})$ constitutes a vector of regressors.

# 3 Empirical findings

## 3.1 Baseline model

The specification of the model used for the estimation of the Balassa-Samuelson effect is conditional upon the results of integration and cointegration tests. In all analysed economies the results of unit root tests indicate non-stationarity of relative prices and relative productivities (Table 4 and 5). First-generation tests unequivocally point to I(1)-ness of these variables. CIPS test, however, suggests that some variables may be of higher degree of integration. Nevertheless, due to the lack of any theoretical underpinnings supporting their integration of order 2, it was assumed that all variables are difference-stationary. Hence the application of panel cointegration tests was possible.

The applied cointegration tests in the case of all countries strongly reject the null hypothesis of no cointegration between relative prices and relative productivities, and thereby suggest the existence of long-run relationship implied by the Balassa-Samuelson hypothesis. Therefore, the following panel error correction model constitutes the basis for the estimation of the Balassa-Samuelson effect:

$$\begin{gathered}\Delta p_{diff_{jt}}^{CEEC\_EA} = \beta_0 + \beta_1 \Delta \tilde{l}_{diff_{jt}}^{CEEC\_EA} + \gamma E\hat{C}T_{jt-1} + u_j + \varepsilon_{jt} \\ E\hat{C}T_{jt} = p_{diff_{jt}}^{CEEC\_EA} - \hat{\delta}_0 - \hat{\delta}_1 \tilde{l}_{diff_{jt}}^{CEEC\_EA},\end{gathered} \tag{36}$$

where $CEEC = \{PL, CZ, HU, SK\}$.

Table 4: Panel unit root tests of $p_{diff_j}^{CEEC\_EA}$

| Variable | IPS | ADF | CIPS |
|---|---|---|---|
| $p_{diff_{jt}}^{PL\_EA}$ | $-1.15$ | 28.29 | $-1.53$ |
| $\Delta p_{diff_{jt}}^{PL\_EA}$ | $-2.71^{***}$ | $43.61^{***}$ | $-3.07^{**}$ |
| $p_{diff_{jt}}^{CZ\_EA}$ | 5.41 | 11.37 | $-1.32$ |
| $\Delta p_{diff_{jt}}^{CZ\_EA}$ | $-8.67^{***}$ | $112.35^{***}$ | $-2.98^{**}$ |
| $p_{diff_{jt}}^{HU\_EA}$ | $-0.42$ | 22.29 | $-1.54$ |
| $\Delta p_{diff_{jt}}^{HU\_EA}$ | $-3.45^{***}$ | $52.16^{***}$ | $-2.04$ |
| $p_{diff_{jt}}^{SK\_EA}$ | 1.90 | 10.13 | $-1.10$ |
| $\Delta p_{diff_{jt}}^{SK\_EA}$ | $-8.17^{***}$ | $101.80^{***}$ | $-3.15^{**}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 5: Panel unit root tests of $\tilde{l}^{CEEC\_EA}_{diff_j}$

| Variable | IPS | ADF | CIPS |
|---|---|---|---|
| $\tilde{l}^{PL\_EA}_{diff_{jt}}$ | 1.45 | 19.07 | $-1.41$ |
| $\Delta\tilde{l}^{PL\_EA}_{diff_{jt}}$ | $-2.28^{***}$ | $45.45^{***}$ | $-1.99$ |
| $\tilde{l}^{CZ\_EA}_{diff_{jt}}$ | 1.60 | 22.68 | $-1.83$ |
| $\Delta\tilde{l}^{CZ\_EA}_{diff_{jt}}$ | $-8.12^{***}$ | $104.05^{***}$ | $-2.67^{*}$ |
| $\tilde{l}^{HU\_EA}_{diff_{jt}}$ | $-0.42$ | 25.47 | $-1.16$ |
| $\Delta\tilde{l}^{HU\_EA}_{diff_{jt}}$ | $-5.30^{***}$ | $69.80^{***}$ | $-2.66^{*}$ |
| $\tilde{l}^{SK\_EA}_{diff_{jt}}$ | 1.07 | 21.05 | $-1.63$ |
| $\Delta\tilde{l}^{SK\_EA}_{diff_{jt}}$ | $-10.99^{***}$ | $135.31^{***}$ | $-2.44^{*}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 6: Panel cointegration tests between $p^{CEEC\_EA}_{diff_{jt}}$ and $\tilde{l}^{CEEC\_EA}_{diff_{jt}}$

| Country | PL | CZ | HU | SK |
|---|---|---|---|---|
| **Pedroni Panel Cointegration Test** | | | | |
| $PG_{ADF}$ | $-3.64^{***}$ | $-4.13^{***}$ | $-2.40^{**}$ | $-6.05^{***}$ |
| **Johansen-Fisher Cointegration Test** | | | | |
| **trace statistics** | | | | |
| $r = 0$ | $44.14^{***}$ | $58.03^{***}$ | $48.87^{***}$ | $55.60^{***}$ |
| $r \leq 1$ | 31.94 | 27.30 | 26.01 | 11.82 |
| **maximum eigenvalue statistics** | | | | |
| $r = 0$ | $38.52^{***}$ | $54.27^{***}$ | $46.62^{***}$ | $60.42^{***}$ |
| $r \leq 1$ | 31.94 | 27.30 | 26.01 | 11.82 |
| **Westerlund Panel Cointegation Tests** | | | | |
| $G_{\tau}$ | $-13.35^{***}$ | $-10.33^{**}$ | $-12.30^{***}$ | $-9.70^{*}$ |
| $P_{\tau}$ | $-27.11^{***}$ | $-6.77^{**}$ | $-14.16^{***}$ | $-13.99^{***}$ |

Notes: Westerlund statistics were computed via the *xtwest* procedure implemented in STATA (see Persyn and Westerlund (2008) for reference). One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

The coefficients were estimated by applying the two-step Engle-Granger procedure. In the first step the cointegration vector coefficients were obtained using FMOLS and DOLS estimators[3]. Those estimates were next used to compute the error correction term. The ECM model was estimated by means of the least squares dummy variable (LSDV) estimator.

The estimates of $\beta_1$ and $\delta_1$, i.e. the short- and long-run elasticity, indicate the extent to which relative productivity growth translates into relative price growth in CEECs versus the euro area. Thereby they reflect the magnitude of the pass-through in the Balassa-Samuelson model. Table 7 presents the estimates of elasticities ($\hat{\delta}_1$ and $\hat{\beta}_1$), as well as error correction coefficients ($\hat{\gamma}$) obtained by means of Engle-Granger procedure.

In all economies under consideration both short- and long-run elasticities are highly significant and correctly signed. Negative error correction coefficients provide strong support for the existence of the long-run equilibrium and imply half-life of the adjustment from less than two years in the case of the Czech Republic to over three years in Slovakia. Contrary to the predictions of the theoretical model, in all analysed economies long-run elasticities are substantially less than unity, implying incomplete pass-through in the Balassa-Samuelson model. However, the obtained results point to considerable heterogeneity in this respect among the considered CEECs. The weakest mechanism seems to be present in Poland and Hungary, where only about half of relative productivity growth passes through to relative price growth versus the euro area (47% and 53% respectively). In Slovakia the pass-through reaches 60%. The mechanism is strongest in the Czech Republic – the estimated long-run elasticity amounts to 0.81 and almost half of the differences in relative productivity growth is reflected in the differential inflation rate versus the euro area already in the first year (the short-run elasticity equal to 0.49).

Table 7: Estimation results of the Balassa-Samuelson model – baseline version

| Country | $\hat{\delta}_1$ | $\hat{\beta}_1$ | $\hat{\gamma}$ | $R^2$ |
|---|---|---|---|---|
| PL | 0.47*** | 0.11*** | −0.26*** | 0.33 |
| CZ | 0.81*** | 0.49*** | −0.32*** | 0.39 |
| HU | 0.53*** | 0.24*** | −0.27*** | 0.26 |
| SK | 0.60*** | 0.29*** | −0.18*** | 0.33 |

Notes: Cointegration vector coefficients were obtained using DOLS estimator. $R^2$ denotes the 'within' coefficient of determination. One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

The results indicating incomplete pass-through in the Balassa-Samuelson model in the case of the CEECs are consistent with numerous previous studies (Table 8). The question of factors which contributed to the weakening of the mechanisms, however, so far has been neglected in the literature. The remainder of the paper aims to contribute to filling this gap.

Table 8: Estimation results of the Balassa-Samuelson model – overview of the empirical literature

| Study | Country / group of countries | Period | Method | Long-run estimates | Short-run estimates |
|---|---|---|---|---|---|
| Wagner and Hlouskova (2004) | CEEC-8 | 1993 - 2001 | techniques for stationary panel data | – | 0.17 – 0.26 |
| Chmielewski (2003) | Poland | 1995 - 2002 | time series cointegration | 0.18 – 0.87 | – |
| Lojschová (2003) | Poland | 1995 - 2002 | OLS | – | 0.39 – 0.42 |
| MacDonald and Wójcik (2003) | Estonia, Hungary, Slovakia, Slovenia | 1995 - 2001 | panel cointegration | 0.41 | – |
| Mihaljek and Klau (2003) | Poland | 1995 - 2001 | OLS | 0.66 | 0.29 |
| Égert (2002) | Poland | 1991 - 2000 | time series cointegration | 0.46 – 0.48 | – |
| Égert et al. (2002) | CEEC-8, Croatia | 1990 - 2000 | panel cointegration | 0.89 | – |
| Halpern and Wyplosz (2001) | CEEC-8 excl. Slovakia, Romania, Russia | 1991 - 1999 | techniques for stationary panel data | – | 0.17 – 0.24 |
| Cipriani (2001) | Poland | 1995 - 1999 | OLS | – | 0.33 |

Notes: CEEC-4 denotes Poland, the Czech Republic, Hungary and Slovakia; CEEC-8 denotes CEEC-4, Slovenia, Lithuania, Latvia and Estonia.

## 3.2 Extended model

The empirical operationalisation of the extended Balassa-Samuelson model is the following wage- and markup-augmented ECM:

$$
\begin{gathered}
\Delta p_{diff_{jt}}^{CEEC\_EA} = \beta_0^{'} + \beta_1^{'} \Delta \tilde{l}_{diff_{jt}}^{CEEC\_EA} + \beta_2^{'} \Delta w_{diff_{jt}}^{CEEC\_EA} + \beta_3^{'} \Delta m_{diff_{jt}}^{CEEC\_EA} \\
+ \gamma^{'} E\hat{C}T_{jt-1} + u_j + \varepsilon_{jt} \\
E\hat{C}T_{jt} = p_{diff_{jt}}^{CEEC\_EA} - \hat{\delta}_0^{'} - \hat{\delta}_1^{'} \tilde{l}_{diff_{jt}}^{CEEC\_EA} - \hat{\delta}_2^{'} w_{diff_{jt}}^{CEEC\_EA} - \hat{\delta}_3^{'} m_{diff_{jt}}^{CEEC\_EA}.
\end{gathered}
\tag{37}
$$

The estimation of the coefficients is preceded by testing the degree of integration of relative wages and markups and the existence of long-run relationship within the extended set of variables. The results of panel unit root tests indicate difference-stationarity of relative wages (Table 9), hence cross-sectoral non-homogeneity of wages. In the case of relative markups the results are mixed (Table 10). First generation tests reject the null hypothesis of non-stationarity at the 10% significance level in the case of Poland, the Czech Republic and Slovakia. However, considering the size distortion of first generation tests, as well as the results of the CIPS test, further analysis assumes that these variables are integrated of order one, which allows to conduct cointegration analysis.

The results of panel cointegration tests (Table 11) consistently indicate the existence of a long-run relationship between relative prices, productivity, wages and markups. Johansen tests suggest that there is more than one cointegrating vector, which may be due to their size distortion. However, should the test results correctly identify the underlying data generating process, the estimation results obtained using Engle-Granger procedure ought to be interpreted with caution.

The results of the extended model estimation (Table 12) point to an improvement in the goodness of fit relative to the baseline model. Moreover, in all cases augmenting the baseline model with trajectories of relative wages and markups significantly increases the estimated magnitude of the Balassa-Samuelson mechanism. This suggests that failing to fulfil the theoretical model assumptions regarding the labour and the product markets was one of factors contributing to the weakening of the pass-through.

Table 9: Panel unit root tests of $w_{diff_{jt}}^{CEEC\_EA}$

| Variable | IPS | ADF | CIPS |
|---|---|---|---|
| $w_{diff_{jt}}^{PL\_EA}$ | $-0.15$ | $24.26$ | $-1.43$ |
| $\Delta w_{diff_{jt}}^{PL\_EA}$ | $-7.70^{***}$ | $97.92^{***}$ | $-2.57^{*}$ |
| $w_{diff_{jt}}^{CZ\_EA}$ | $-0.56$ | $30.67$ | $-1.58$ |
| $\Delta w_{diff_{jt}}^{CZ\_EA}$ | $-6.78^{***}$ | $90.00^{***}$ | $-3.27^{**}$ |
| $w_{diff_{jt}}^{HU\_EA}$ | $-1.38^{*}$ | $31.78$ | $-1.71$ |
| $\Delta w_{diff_{jt}}^{HU\_EA}$ | $-4.42^{***}$ | $62.64^{***}$ | $-2.57^{*}$ |
| $w_{diff_{jt}}^{SK\_EA}$ | $0.24$ | $21.59$ | $-0.61$ |
| $\Delta w_{diff_{j}}^{SK\_EA}$ | $-9.35^{***}$ | $116.37^{***}$ | $-3.13^{**}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 10: Panel unit root tests of $m_{diff_{jt}}^{CEEC\_EA}$

| Variable | IPS | ADF | CIPS |
|---|---|---|---|
| $m_{diff_{jt}}^{PL\_EA}$ | $-0.96$ | $33.78^{*}$ | $-1.70$ |
| $\Delta m_{diff_{jt}}^{PL\_EA}$ | $-5.63^{***}$ | $75.03^{***}$ | $-2.32^{*}$ |
| $m_{diff_{jt}}^{CZ\_EA}$ | $-1.63^{*}$ | $31.96$ | $-2.12$ |
| $\Delta m_{diff_{jt}}^{CZ\_EA}$ | $-4.55^{***}$ | $45.10^{***}$ | $-3.11^{**}$ |
| $m_{diff_{jt}}^{HU\_EA}$ | $-0.59$ | $19.97$ | $-1.37$ |
| $\Delta m_{diff_{jt}}^{HU\_EA}$ | $-6.56^{***}$ | $83.39^{***}$ | $-2.87^{**}$ |
| $m_{diff_{jt}}^{SK\_EA}$ | $-1.43^{*}$ | $34.20^{*}$ | $-2.02$ |
| $\Delta m_{diff_{jt}}^{SK\_EA}$ | $-3.76^{***}$ | $54.58^{***}$ | $-2.35^{*}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 11: Panel cointegration tests between $p_{diff_{jt}}^{CEEC\_EA}$, $l_{diff_{jt}}^{CEEC\_EA}$, $w_{diff_{jt}}^{CEEC\_EA}$ and $m_{diff_{jt}}^{CEEC\_EA}$

| Country | PL | CZ | HU | SK |
|---|---|---|---|---|
| **Pedroni Panel Cointegration Test** | | | | |
| $PG_{ADF}$ | $-2,87^{**}$ | $-3,54^{***}$ | $-3,18^{***}$ | $-1,80^{**}$ |
| **Johansen-Fisher Cointegration Test** | | | | |
| **trace statistics** | | | | |
| $r=0$ | $241,60^{***}$ | $263,40^{***}$ | $232,60^{***}$ | $207,20^{***}$ |
| $r\leq 1$ | $108,20^{***}$ | $92,77^{***}$ | $96,86^{***}$ | $84,84^{***}$ |
| $r\leq 2$ | $43,91^{***}$ | $31,23$ | $44,62^{***}$ | $28,10$ |
| $r\leq 3$ | $25,94$ | $18,56$ | $26,34$ | $19,99$ |
| **maximum eigenvalue statistics** | | | | |
| $r=0$ | $193,10^{***}$ | $230,00^{***}$ | $189,10^{***}$ | $178,40^{***}$ |
| $r\leq 1$ | $90,90^{***}$ | $89,37^{***}$ | $78,89^{***}$ | $79,86^{***}$ |
| $r\leq 2$ | $41,18^{**}$ | $29,33$ | $40,53^{***}$ | $26,12$ |
| $r\leq 3$ | $25,94$ | $18,56$ | $26,34$ | $19,99$ |
| **Westerlund Panel Cointegation Tests** | | | | |
| $G_\tau$ | $-8,34^{**}$ | $-12,35^{**}$ | $-5,84$ | $-5,27$ |
| $P_\tau$ | $-7,99^{***}$ | $-8,69^{**}$ | $-10,07^{***}$ | $-6,43^{**}$ |

Note: Westerlund statistics were computed via the *xtwest* procedure implemented in STATA (see Persyn and Westerlund (2008) for reference). One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 12: Estimation results of the Balassa-Samuelson model – extended version

| Country | $\hat{\delta}_1'$ | $\hat{\delta}_2'$ | $\hat{\delta}_3'$ | $\hat{\beta}_1'$ | $\hat{\beta}_2'$ | $\hat{\beta}_3'$ | $\hat{\gamma}'$ | $R^2$ |
|---|---|---|---|---|---|---|---|---|
| PL | $0.73^{***}$ | $0.59^{***}$ | $1.07^{***}$ | $0.56^{***}$ | $0.51^{***}$ | $1.02^{***}$ | $-0.45^{***}$ | 0.60 |
| CZ | $0.88^{***}$ | $0.65^{***}$ | $0.77^{***}$ | $0.80^{***}$ | $0.63^{***}$ | $0.95^{***}$ | $-0.34^{***}$ | 0.84 |
| HU | $0.70^{***}$ | $0.84^{***}$ | $1.04^{***}$ | $0.66^{***}$ | $0.60^{***}$ | $0.94^{***}$ | $-0.37^{***}$ | 0.85 |
| SK | $0.77^{***}$ | $0.85^{***}$ | $1.04^{***}$ | $0.65^{***}$ | $0.64^{***}$ | $0.89^{***}$ | $-0.35^{***}$ | 0.71 |

Notes: Cointegration vector coefficients were obtained using DOLS estimator. $R^2$ denotes the 'within' coefficient of determination. One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

## 3.3 The effectiveness of labour and product market in absorbing the Balassa-Samuelson effect

The incomplete transmission of relative productivity growth to relative price growth in CEECs, as compared to the euro area, may to some extent be due to the relative wages and relative margins trajectory being inconsistent with the theoretical assumptions, as the estimated magnitude of the Balassa-Samuelson mechanism is greater when controlled for these factors. The important question, however, is whether relative wages and markups can be treated as mechanisms absorbing the Balassa-Samuelson effect. Below we propose tests which address this question.

The labour market in a given economy mitigates the convergence-induced inflationary pressure, provided that three conditions are fulfilled: (1) there is a long-run relationship between wages in the tradable and non-tradable sectors (operationalised using cointegration tests), (2) the direction of this relationship is from wages in the tradable sector to wages in the non-tradable sector (operationalised using causality tests), (3) the pass-through is incomplete, thus the long-term elasticity is significantly lower than unity (operationalised using the test of restrictions on the cointegrating vector). The labour market acts as the Balassa-Samuelson effect absorption channel, if the mechanism of incomplete levelling of wages described above is stronger in a developing than in a developed economy. Therefore the above mentioned conditions pertain to relative wages, i.a. the difference between wages in sector $j$ home and abroad, $w_j^{CEEC\_EA}$.

Table 13: Panel unit root tests of $w_{N_j}^{CEEC\_EA}$

| Variable | IPS | ADF | CIPS |
|---|---|---|---|
| $w_{N_{jt}}^{PL\_EA}$ | $-1.14$ | 31.62 | $-1.94$ |
| $\Delta w_{N_{jt}}^{PL\_EA}$ | $-3.48^{***}$ | $51.86^{***}$ | $-2.77^{*}$ |
| $w_{N_{jt}}^{CZ\_EA}$ | 0.17 | 27.95 | $-1.64$ |
| $\Delta w_{N_{jt}}^{CZ\_EA}$ | $-6.77^{***}$ | $83.63^{***}$ | $-3.14^{**}$ |
| $w_{N_{jt}}^{HU\_EA}$ | 3.61 | 13.48 | $-1.60$ |
| $\Delta w_{N_{jt}}^{HU\_EA}$ | $-7.54^{***}$ | $90.81^{***}$ | $-2.80^{**}$ |
| $w_{N_{jt}}^{SK\_EA}$ | 1.30 | 24.09 | $-1.99$ |
| $\Delta w_{N_{jt}}^{SK\_EA}$ | $-6.73^{***}$ | $85.04^{***}$ | $-3.57^{**}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 14: ADF and KPSS tests of $w_T^{CEEC\_EA}$

| Country | PL | CZ | HU | SK |
|---|---|---|---|---|
| | | ADF | | |
| $w_T^{CEEC\_EA}$ | $-2.63$ | $0.23$ | $-1.15$ | $-2.37$ |
| $\Delta w_T^{CEEC\_EA}$ | $-2.52$ | $-4.58^{***}$ | $-3.77^{**}$ | $-4.06^{**}$ |
| | | KPSS | | |
| $w_T^{CEEC\_EA}$ | $0.15^{**}$ | $0.18^{**}$ | $0.17^{**}$ | $0.16^{**}$ |
| $\Delta w_T^{CEEC\_EA}$ | $0.31$ | $0.41^{*}$ | $0.46^{*}$ | $0.44^{*}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 15: Test for the effectiveness of the labour market (1): the cointegration test

| Country | PL | CZ | HU | SK |
|---|---|---|---|---|
| | | **Pedroni Panel Cointegration Test** | | |
| $PG_{ADF}$ | $-1.98^{**}$ | $-2.66^{***}$ | $-6.17^{***}$ | $-2.17^{**}$ |
| | | **Johansen-Fisher Cointegration Test** | | |
| | | **trace statistics** | | |
| $r = 0$ | $62.67^{***}$ | $54.08^{***}$ | $86.89^{***}$ | $77.04^{***}$ |
| $r \leq 1$ | $25.82$ | $16.24$ | $19.94$ | $17.32$ |
| | | **maximum eigenvalue statistics** | | |
| $r = 0$ | $60.39^{***}$ | $56.92^{***}$ | $89.12^{***}$ | $80.62^{***}$ |
| $r \leq 1$ | $25.82$ | $16.24$ | $19.94$ | $17.32$ |
| | | **Westerlund Panel Cointegration Tests** | | |
| $G_\tau$ | $-11.56^{***}$ | $-10.95^{***}$ | $-10.54^{**}$ | $-10.55^{**}$ |
| $P_\tau$ | $-14.18^{***}$ | $-7.68^{***}$ | $-12.49^{***}$ | $-12.50^{***}$ |

Note: Westerlund statistics were computed via the *xtwest* procedure implemented in STATA (see Persyn and Westerlund (2008) for reference). One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Panel unit root tests unequivocally indicate that wages in the non-tradable sectors in CEECs relative to the euro area (variable $w_{N_j}^{CEEC\_EA}$) are integrated of order one (Table 13). The results of ADF and KPSS tests for relative wages in the tradable sector (variable $w_T^{CEEC\_EA}$), however, are mixed (Table 14). While both point to the non-stationarity of series, in the case of Poland the ADF test fails to reject the null hypothesis of non-stationarity of first differences, whereas for all other countries the KPSS test rejects the null hypothesis of stationarity of first differences at the 10% significance level. Nevertheless, due to the lack of theoretical underpinnings for the non-stationarity of growth rate of wages and considering the shortness of the series, we assumed that all relative wages in the tradable sector are I(1), which allows to apply cointegration tests.

The results of cointegration tests (Table 15) indicate that in all cases there exists a long-run relationship between wages in the tradable and the non-tradable sector. This implies causality in at least one direction.

Let us consider the following system of error-correction equations:

$$\begin{aligned} \Delta w_{N_j,t}^{CEEC\_EA} &= \beta_0^w + \beta_1^w \Delta w_{T,t}^{CEEC\_EA} + \gamma^w E\hat{C}T_{jt-1} + u_j + \varepsilon_{jt} \\ \Delta w_{T,t}^{CEEC\_EA} &= \beta_0^{w'} + \beta_1^{w'} \Delta w_{N_j,t}^{CEEC\_EA} + \gamma^{w'} E\hat{C}T_{jt-1} + v_j + \epsilon_{it} \\ E\hat{C}T_{jt} &= w_{N_j,t}^{CEEC\_EA} - \hat{\delta_0^w} - \hat{\delta_1^w} w_{T,t}^{CEEC\_EA}. \end{aligned} \tag{38}$$

Should the null hypothesis of insignificance of $\gamma^w$ be rejected, wages in the tradable sector cause wages in the non-tradable sector in the long-run. The causality is reverse (wages in the non-tradable sector cause wages in the tradable sector), if the hypothesis can be rejected with respect to $\gamma'$.

Table 16: Test for the effectiveness of the labour market (2): the causality test

| Country | $\hat{\gamma^w}$ | $\hat{\gamma^{w'}}$ |
|---|---|---|
| **PL** | $-0.23^{***}$ | 0.03 |
| **CZ** | $-0.27^{***}$ | 0.01 |
| **HU** | $-0.38^{***}$ | $0.19^{***}$ |
| **SK** | $-0.30^{***}$ | $-0.05$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

The results of causality tests (Table 16) indicate that in all cases wages in the tradable sector cause wages in the non-tradable sector in the long-run. Only Hungary exhibits reverse long-run causality. In the case of Poland, the Czech Republic and Slovakia wages in the tradable sector are therefore weakly exogenous, hence constitute a stochastic trend driving wages in the non-tradable sector.

Table 17: Test for the effectiveness of the labour market (3): elasticity estimates

| Country | $\hat{\delta_1^w}$ | $H_0 : \hat{\delta_1^w} = 1$ |
|---|---|---|
| **PL** | $1.18^{***}$ | $21.3^{***}$ |
| **CZ** | $1.13^{***}$ | $8.26^{***}$ |
| **HU** | $1.36^{***}$ | $39.84^{***}$ |
| **SK** | $0.86^{***}$ | $9.69^{***}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

The third element of the labour market effectiveness test is the estimate of the long-run elasticity of wages in non-tradable sectors with respect to wages in the tradable sector (Table 17). The estimation results indicate that all long-run elasticities are significantly different from unity, which is consistent with the results of relative wages integration tests (Table 9). However, the pass-through is incomplete only in the case of Slovakia. In all other countries wages not only level off across sectors, but even overshoot, which is evidenced by the long-run elasticity significantly greater than unity.

Joint results of the test for the effectiveness of the labour market indicate that in the analysed period it mitigated the Balassa-Samuelson effect only in Slovakia, as the transmission of wage increases across sectors in the long-run was incomplete (long-run elasticity was significantly lower than unity) and the direction thereof was consistent with the theoretical assumptions. In the case of Poland and the Czech Republic the direction of the transmission was also correct, meaning that the trajectory of relative wages constitute an element of the Balassa-Samuelson mechanism. In those two economies the labour market, however, had an amplifying effect with respect to catching-up induced inflationary pressure. In the case of Hungary, in turn, the causality tests indicate that the trajectory of relative wages cannot be attributed solely to the behaviour of wages in the tradable sector. Hence, there exists a different mechanism of cross-sectoral wage formation than this described by Balassa and Samuelson.

Table 18: Panel unit root tests of $m_{N_j}^{CEEC\_EA}$

| Variable | IPS | ADF | CIPS |
|---|---|---|---|
| $m_{N_{jt}}^{PL\_EA}$ | $-1.25$ | $34.7^{*}$ | $-1.94$ |
| $\Delta m_{N_{jt}}^{PL\_EA}$ | $-5.59^{***}$ | $71.57^{***}$ | $-2.93^{**}$ |
| $m_{N_{jt}}^{CZ\_EA}$ | $-1.06$ | $33.03$ | $-2.19$ |
| $\Delta m_{N_{jt}}^{CZ\_EA}$ | $-6.14^{***}$ | $78.94^{***}$ | $-2.84^{**}$ |
| $m_{N_{jt}}^{HU\_EA}$ | $-0.15$ | $21.96$ | $-1.48$ |
| $\Delta m_{N_{jt}}^{HU\_EA}$ | $-3.89^{***}$ | $93.46^{***}$ | $-2.43^{*}$ |
| $m_{N_{jt}}^{SK\_EA}$ | $-0.91$ | $26.14$ | $-2.05$ |
| $\Delta m_{N_{jt}}^{SK\_EA}$ | $-7.57^{***}$ | $93.32^{***}$ | $-2.72^{*}$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

In a given economy the product market dampens the Balassa-Samuelson mechanism, if an increase in wages in the non-tradable sector which is not compensated with a respective increase in productivity does not transmit in full to an increase in prices owing to a decrease in a profit margin. By analogy to the labour market, the conditions for the effectiveness of the product market in absorbing convergence-driven inflationary pressure can be defined as follows: (1) there exists a long-run relationship between wages and markups in the non-tradable sector, (2) the direction of this relationship is from wages to markups, (3) elasticity of markups with respect to wages is negative. Sectoral markups have been defined as markups over variable costs (i.e. the sum of compensation of employees and intermediate inputs). This may be the source of endogeneity in the model where markups are regressed against wages. Therefore productivity in the tradable sector has been used as an instrument for wages in non-tradable sectors.

Panel unit root tests indicate that the markups in non-tradable sectors in the CEECs relative to the euro area are I(1) (Table 18 ), as are the relative wages (Table 13). The results of cointegration tests for this set of variables, in turn, are mixed: first generation tests suggest that there exists a long-run relationship, but for Poland and Slovakia the Westerlund test fails to reject the null hypothesis of no cointegration. In this case rejecting the null hypothesis may result in an error of the first kind, since size distortions of the IPS test and the Fisher-based ADF test are possible in the light of potential cross-sectional dependence.

Table 19: Test for the effectiveness of the product market (1): the cointegration test

| Country | PL | CZ | HU | SK |
|---|---|---|---|---|
| **Pedroni Panel Cointegration Test** | | | | |
| $PG_{ADF}$ | $-1.77^{**}$ | $-3.98^{***}$ | $-3.34^{***}$ | $-5.65^{***}$ |
| **Johansen-Fisher Cointegration Test** | | | | |
| **trace statistics** | | | | |
| $r = 0$ | $56.88^{***}$ | $66.29^{***}$ | $66.75^{***}$ | $41.35^{**}$ |
| $r \leq 1$ | $35.65^{*}$ | $34.22^{*}$ | $36.03^{*}$ | 29.63 |
| **maximum eigenvalue statistics** | | | | |
| $r = 0$ | $51.89^{***}$ | $61.79^{***}$ | $62.50^{***}$ | $37.89^{**}$ |
| $r \leq 1$ | $35.65^{*}$ | $34.22^{*}$ | $36.03^{*}$ | 29.63 |
| **Westerlund Panel Cointegation Tests** | | | | |
| $G_\tau$ | $-2.93$ | $-12.17^{***}$ | $-10.01^{***}$ | $-5.45$ |
| $P_\tau$ | $-5.84$ | $-5.90^{*}$ | $-5.98^{*}$ | $-4.54$ |

Note: Westerlund statistics were computed via the *xtwest* procedure implemented in STATA (see Persyn and Westerlund (2008) for reference). One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

By analogy to the labour market, we estimate the following system of error-correction equations:

$$\begin{aligned} \Delta m_{N_j,t}^{CEEC\_EA} &= \beta_0^m + \beta_1^m \Delta w_{N_j,t}^{CEEC\_EA} + \gamma^m E\hat{C}T_{jt-1} + u_j + \varepsilon_{jt} \\ \Delta w_{N_j,t}^{CEEC\_EA} &= \beta_0^{m'} + \beta_1^{m'} \Delta m_{N_j,t}^{CEEC\_EA} + \gamma^{m'} E\hat{C}T_{jt-1} + v_j + \epsilon_{it} \\ E\hat{C}T_{jt} &= m_{N_j,t}^{CEEC\_EA} - \hat{\delta_0^m} - \hat{\delta_1^m} w_{N_j,t}^{CEEC\_EA}. \end{aligned} \tag{39}$$

The estimation results of the cointegrating vector (Table 21) confirm the indication of the Westerlund test. In the case of Poland and Slovakia there exists no long-run relationship between markups and wages in non-tradable sectors. For the Czech Republic and Hungary, in turn, long-run elasticities of markups to wages are significantly different from zero, and the direction of the long-run causality (Table 20) is consistent with the theoretical predictions (from wages to markups). Hence, in the case of the latter economies their product markets to some extent absorb the inflationary pressure resulting from the convergence process.

Table 20: Test for the effectiveness of the product market (2): the causality test

| Country | $\hat{\gamma^m}$ | $\hat{\gamma^{m'}}$ |
|---|---|---|
| **PL** | – | – |
| **CZ** | $-0.28^{***}$ | 0.14 |
| **HU** | $-0.39^{***}$ | $-0.12$ |
| **SK** | – | – |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

Table 21: Test for the effectiveness of the product market (3): elasticity estimates

| Country | $\hat{\delta_1^m}$ |
|---|---|
| **PL** | $-0.03$ |
| **CZ** | $-0.06^{**}$ |
| **HU** | $-0.09^{***}$ |
| **SK** | $-0.01$ |

Notes: One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

## 3.4 Quantification of the Balassa-Samuelson effect

The quantification of the Balassa-Samuelson effect is given by the product of (1) the estimated elasticities in the Balassa-Samuelson model, (2) the average growth rate of relative productivity, wages and markups over the sample period[4], and (3) the share of the non-tradable sector in the economy. The obtained quantification results indicate that the highest magnitude of the Balassa-Samuelson effect over the sample period (1995-2010) was to be observed in the case of Poland – on average the inflationary pressure driven by the catching-up process amounted to 3.3% per annum, hence it contributed to over 80% of inflation differentials versus the euro area. In Hungary the effect was the weakest (on average 1.2 p.p. per annum) while the inflation differentials vis-a-vis the euro area were the highest of all the considered economies (on average almost 7 p.p.). Thereby, their considerable proportion was determined by factors other than the real convergence process, one of them being arguably considerable dynamics of labour costs, which however was not driven by the catching-up process (basing on the results of causality tests, Table 16). In Slovakia and the Czech Republic the magnitude of the Balassa-Samuelson effect was moderate (on average 2.2 and 2.4 p.p., respectively) and to a large extent accounted for the observed inflation differentials relative to the euro area (over 90%).

Table 22: Quantification of the Balassa-Samuelson effect – baseline vs. extended model

| | | estimated Balassa-Samuelsn effect | | |
|---|---|---|---|---|
| Country | average inflation differentials versus the euro area | baseline model | extended model (excl. potential absorption channels) | extended model (incl. potential absorption channels) |
| PL | 4.0 p.p | 1.9 p.p (48%) | 2.9 p.p (72%) | 3.3 p.p (83%) |
| CZ | 2.3 p.p | 1.6 p.p (76%) | 1.9 p.p (81%) | 2.2 p.p (94%) |
| HU | 6.9 p.p | 1.2 p.p (17%) | 1.6 p.p (24%) | 1.2 p.p (18%) |
| SK | 2.5 p.p | 1.9 p.p (76%) | 2.5 p.p (99%) | 2.3 p.p (92%) |

Notes: Inflation differentials were computed on the basis of value-added deflators excluding agriculture and fishing (NACE rev.1.1. sections A and B). Contributions of the Balassa-Samuelson effect to inflation differentials are presented in brackets.

The comparison of the effects quantified using estimation results of the baseline and the extended model (Table 22) suggests that the baseline model specification – widely used in the literature – leads to biased estimates. There are two reasons for the bias. First, ignoring the non-fulfilment of theoretical model assumptions regarding wages and markups distorts estimation results of elasticity of relative prices with respect to relative productivity (ommitted variable bias). Secondly, the effect quantified on the basis of the baseline model does not take into account the possible mitigating or amplifying impact of the labour and product markets with respect to the Balassa-Samuelson mechanism. The problem is particularly evident in the case of Poland, where the effect quantified according to the baseline model estimates is approximately 40% (1.4 p.p.) lower than according to the wage- and markup-augmented model.

# Conclusions

The aim of this study is to determine the magnitude of the structural component of inflation differentials in the CEECs relative to the euro area which can be attributed to the catching-up process of these economies, i.e. the magnitude of the Balassa-Samuelson effect. Due to relatively short time span and low frequency of series available for the CEECs, which virtually prohibits the application of cointegration analysis, we propose a novel approach to estimating the Balassa-Samuelson model. It consists in designing the panel so that the relation between the tradable sector and various sub-sectors of the non-tradable sector constitutes the cross-sectional dimension. This allows to apply panel econometric techniques, but at the same time to concentrate on the single economy perspective.

Consistently with earlier studies, the obtained results indicate that the pass-through in the Balassa-Samuelson model is incomplete, i.e. the long-run elasticity of relative prices with respect to relative productivity is significantly less than unity. An important, though thus far neglected in the literature, question is which factors contribute to the weakening of the pass-through and can they be treated as mechanisms absorbing the catching-up induced inflationary pressure. This paper aims to contribute to filling this gap by proposing tests for the effectiveness of the labour and product markets in absorbing the Balassa-Samuelson effect. The results thereof indicate that only in the case of Slovakia the convergence-driven inflationary pressure is to some extent mitigated by the incomplete levelling of wages across sectors. For Poland and the Czech Republic the labour market acts as an amplifier to the Balassa-Samuelson effect. In Hungary, on the other hand, sectoral wage developments are determined by mechanisms other than those outlined by Balassa and Samuelson. In the case of the Czech Republic and Hungary an increase in wages in the non-tradable sector is to some extent absorbed by the corresponding decrease in markups, hence the product market acts as an absorption mechanism to the Balassa-Samuelson effect.

The obtained results suggest that ignoring the non-fulfilment of theoretical model assumptions regarding the product and labour markets, which is common in the literature, causes the estimates of the Balassa-Samuelson effect to be biased. In the case of Poland the resulting error amounts to 1.4 percentage points, which is equivalent to a 40-percent underestimation of this effect. Therefore we propose to extend the baseline model to encompass relative wages and markups developments.

In the years 1995-2010 average inflation differentials between the CEECs and the euro area which resulted from the real convergence process amounted to 1.2 percentage points in Hungary, slightly above 2 percentage points in the Czech Republic and Slovakia, and 3.3 percentage points in Poland. The systematic inflation differentials relative to the euro area were therefore comparable in size (even slightly greater in the case of Poland) to those experienced by the peripheral member states (Table 1). Thus the results of the study suggest that in the case of the considered CEECs the risk of common monetary policy inadequacy may be serious, which negatively affects the balance between benefits and costs of their integration with the euro area.

# References

**Bai J., Ng S. (2004)**: *A PANIC Attack on Unit Roots and Cointegration*, Econometrica, 72(4), 1127–1177.

**Balassa B. (1964)**: *The Purchasing-Power Parity Doctrine: A Reappraisal*, The Journal of Political Economy, 72(6), 584–596.

**Banerjee A. (1999)**: *Panel Data Unit Roots and Cointegration: An Overview*, Oxford Bulletin of Economics and Statistics, 61(0), 607–629.

**Banerjee A., Marcellino M., Osbat C. (2005)**: *Testing for PPP: Should We Use Panel Methods?*, Empirical Economics, 30(1), 77–91.

**Chmielewski T. (2003)**: *Znaczenie efektu Balassy-Samuelsona dla polskiej polityki pieniężnej*, Materiały i Studia NBP, 163.

**Cipriani M. (2001)**: *The Balassa-Samuelson Effect in Transition Economies*, mimeo.

**Égert B. (2002)**: *Estimating the impact of the Balassa-Samuelson effect on inflation and the real exchange rate during the transition*, Economic Systems, 26(1), 1–16.

**Égert B., Drine I., Lommatzsch K., Raut C. (2002)**: *The Balassa-Samuelson effect in Central and Eastern Europe: Myth or reality?*, William Davidson Institute Working Paper Series, 483.

**Fisher R. (1932)**: *Statistical Methods for Research Workers*, Oliver & Boyd, Edinburgh.

**Halpern L., Wyplosz C. (2001)**: *Economic Transformation and Real Exchange Rates in the 2000s: The Balassa-Samuelson Connection*, ECE Discussion Paper Series, 1.

**Im K., Pesaran M., Shin Y. (2003)**: *Testing for Unit Roots in Heterogenous Panels*, Journal of Econometrics, 115(1), 53–74.

**Kao C., Chiang M. (2000)**: *On the estimation and inference of a cointegrated regression in panel data*, Advanced in Economics, 15(1), 179–222.

**Konopczak K., Torój A. (2010)**: *Estimating the Baumol-Bowen and Balassa-Samuelson effects in the Polish Economy - a Disaggregated Approach*, Central European Journal of Economic Modelling and Econometrics, 2(2), 117–150.

**Lojschová A. (2003)**: *Estimating the Impact of the Balassa-Samuelson Effect in Transition Economies*, Institute for Advanced Studies Vienna Economic Series, 140.

**MacDonald R., Wójcik C.** (**2003**): *Catching Up: The Role of Demand, Supply and Regulated Price Effects on the Real Exchange Rates of Four Accession Countries*, CESifo Working Paper Series, 899.

**Maddala G., Wu S.** (**1999**): *A Comparative Study of Unit Root Tests with Panel Data a New Simple Test*, Oxford Bulletin of Economics and Statistics, 61(1), 631–652.

**Mihaljek D., Klau M.** (**2003**): *The Balassa-Samuelson effect in central Europe: a disaggregated analysis*, BIS Working Paper Series, 143.

**Moon R., Perron B.** (**2004**): *Testing for a unit root in panels with dynamic factors*, Journal of Econometrics, 122(1), 81–126.

**Pedroni P.** (**2004**): *Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis*, Econometric Theory, 20(3), 597–625.

**Persyn D., Westerlund J.** (**2008**): *Error Correction Based cointegration Tests for Panel Data*, Stata Journal, 8(2), 232–241.

**Pesaran M.** (**2003**): *A Simple Panel Unit Root Test in the Presence of Cross Section Dependence*, Cambridge Working Papers in Economics Series, 346.

**Phillips P., Hansen B.** (**1990**): *Statistical Inference in Instrumental Variables Regression with I(1) Processes*, The Review of Economic Studies, 57(1), 99–125.

**Phillips P., Moon H.** (**1999**): *Linear Regression Limit Theory for Nonstationary Panel Data*, Econometrica, 67(5), 1057–1111.

**Saikkonen P.** (**1991**): *Asymptotically effcient estimation of cointegration regressions*, Econometric Theory, 7(1), 1–21.

**Samuelson P.** (**1964**): *Theoretical Notes on Trade Problems*, The Review of Economics and Statistics, 46(2), 145–154.

**Wagner M.** (**2005**): *The Balassa-Samuelson Effect in 'East & West'. Differences and Similarities*, Institute for Advanced Studies Vienna Economic Series, 180.

**Wagner M., Hlouskova J.** (**2004**): *What's Really the Story with this Balassa-Samuelson Effect in the CEEC?*, Universiteat Bern Volkswirtschaftliches Institut Diskussionsschriften, 16.

**Wagner M., Hlouskova J.** (**2010**): *The Performance of Panel Cointegration Methods: Results from a Large Scale Simulation Study*, Econometric Reviews, 29(2), 182–223.

**Walters A.** (**1994**): *Walters Critique*, w: Matthews, K. (red.) The Economics and Politics of Money: The Selected Essays of Alan Walters, Elgar Publishing, London.

**Westerlund J.** (**2007**): *Testing for Error Correction in Panel Data*, Oxford Bulletin of Economics and Statistics, 69(6), 709–748.

www.nbp.pl