MF Working Paper Series

No 06-2010 September 2010

Karolina Konopczak, Andrzej Torój

Estimating the Baumol-Bowen and Balassa-Samuelson effects in the Polish economy – a disaggregated approach

Ministry of Finance in Poland

Financial Policy, Analyses and Statistics Department

Estimating the Baumol-Bowen and Balassa-Samuelson effects in the Polish economy – a disaggregated approach

Karolina Konopczak*, Andrzej Torój†

Abstract

This paper estimates the magnitude of the Baumol-Bowen and Balassa-Samuleson effects in the Polish economy. The purpose of the analysis is to establish to what extent the differential price dynamics in Poland and in the euro area and the real appreciation of PLN against EUR are explained by the differential in respective productivity dynamics. The historical contribution of the Baumol-Bowen effect to Polish inflation rate is estimated at $0.7-1.0$ percentage points in the short run. According to estimation results, the Balassa-Samuelson effect contributed around 0.9 to 1.3 percentage point per annum to the rate of relative price growth between Poland and the euro area and 0.9 to 1.6 p.p. to real exchange rate appreciation. Sub-sample calculations and productivity trends over the last decade suggest that this impact should be declining. However, its size is still non-negligible for policymakers in the context of euro adoption in Poland.

JEL Classification: C33, E31, F31, F41

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

*Ministry of Finance
†Ministry of Finance

1 Introduction

The Balassa-Samuelson hypothesis (Balassa, 1964; Samuelson, 1964) provides a framework which has become very popular in international macroeconomics to explain cross-country and cross-sector inflation differentials. The claim is that countries with relatively high productivity dynamics in the tradable sector face higher inflation rates than countries with a more balanced productivity growth. For this reason the effect should be of a higher magnitude in catching-up economies, such as the New Member States (NMS) of the European Union – including Poland. The additional inflation stems from the non-tradable sector, lagging behind the producers of tradable goods in terms of productivity, but facing the pressure of growing labour costs.

The economic reasoning behind this mechanism is sometimes decomposed into Baumol-Bowen effect, explaining the cross-sectoral inflation differential, and encompassing Balassa-Samuelson effect, addtionally accounting for the real exchange rate appreciation. Over the recent decade, both issues have been investigated in a wide range of empirical studies of catching-up countries, especially the NMS (see Appendix for details).

At the same time, all the NMS are obliged to adopt the euro as a common European currency as soon as they meet the criteria. Half of them (Slovenia, Malta, Cyprus, Slovakia, Estonia) will have joined the euro area by January 2011. It is predominantly their involvement in the process of European monetary integration that makes the Baumol-Bowen (henceforth: BB) and Balassa-Samuelson (BS) effect of particular interest for macroeconomists and policymakers. This is motivated by at least two main reasons.

The first one is the construction of the price stability criterion. According to Article 140 of the consolidated version of the Treaty on European Union and of the Treaty on the Functioning of the European Union (as resulting from the Treaty of Lisbon)[1] and the Protocol 13 on the convergence criteria, Member State should have *a price performance that is sustainable and an average rate of inflation, observed over a period of one year before the examination, that does not exceed by more than 1.5 percentage points that of, at most, the three best performing Member States in terms of price stability.* The relevant index is the Harmonized Index of Consumer Prices (HICP). This implies that (i) the price dynamics taken into account at the time of assessment covers the entire consumer basket, including both tradable and nontradable goods and (ii) the assessment is made in comparison to the best performing economies in the EU-27 group and the feasible disparity is precisely defined.

As high productivity dynamics in the NMS can be treated as an equilibrium phenomenon under the catching-up process, a significant Balassa-Samuelson effect boosts the equilibrium inflation rate (also as measured by HICP). From the Polish point of view, this hampers the feasibility of this criterion in a straightforward manner. It is highly probable that the group of best performers among EU-27 would contain advanced economies, with lower equilibrium inflation rates. The absence of steady-state inflation differentials, e.g. from welfare analyses of meeting the criteria (see Lipińska, 2008), has been subject to criticism as a potential source of underestimating the cost. Provided the admissible disparity of 1.5 p.p., it is therefore of crucial importance for domestic policymakers how much the BS effect contributes to the domestic inflation. Better understanding of the BS-induced inflation might also play a role in the assessment of convergence sustainability.

The second aspect are competitiveness considerations within the euro area. Not any more will the real appreciation be channelled through the nominal exchange rate adjustment, which – coupled with exogeneity of the big foreign economy – leaves the absorption of Balassa-Samuelson effect to the domestic deflator. The question is how this would affect the price competitiveness of the Polish economy. On the one hand, the price adjustment should be concentrated in the nontradable sector, which should have little direct impact on the prices of domestically produced tradable goods and hence on relative competitiveness of domestic producers. On the other hand, modelling the transport and distribution costs as a nontradable component of a tradable price (see e.g. Corsetti and Dedola, 2005) has got much support in the recent literature. Provided that consuming tradable goods requires some nontradable input, higher price dynamics from of the nontradables could spill over into the overall price level. Moreover, the role of the nontradable sector is restoring cross-country equilibrium in a monetary union is increasingly emphasized (European Commission, 2009). The flow of resources from tradable to nontradable production, encouraged by higher prices, could boost the latter sector and affect the effcient structure of the economy along with potential growth. Excess nontradable sector weakens the competitiveness channel as a mechanism of adjustment after asymmetric shocks and leaves the economy vulnerable to shocks. It is often argued that rebalancing the resources back to the tradable sector should be a critical element of the recovery in the Baltic economies after the 2004-2007 boom.

For both reasons, a quantitative, up-to-date assessment of the BS effect contribution to domestic inflation is crucial. Policymakers could take both issues into account, as long as the effects are quantified and regarded as significant. No clear consensus view emerges from the empirical literature for Poland. Moreover, due to strong disinflation in late 1990s and early 2000s, estimates over samples ending a few years ago might overestimate the impact and contributions based on more recent data would be more useful. This paper attempts to address this need.

The rest of the paper is organized as follows. Section 2 develops the standard Baumol-Bowen and Balassa-Samuelson modelling framework and reviews previous empirical literature. In Section 3, the model is estimated via panel techniques and in Section 4 the contribution of BS effect to Polish inflation is quantitatively assessed.

2 Theoretical framework

We derive the model of BB and BS effects, following the standard approach in the literature (see the Appendix for a number of references). Starting with production functions and standard firms' profit maximization conditions, we end up with equations that express relative (cross-sector) inflation as a function of relative productivity dynamics (BB). Also, we develop the relationship between real exchange rate dynamics and relative productivities, calculated jointly from cross-sector and cross-country perspective (BS).

In this analytical framework, we make use of the following economic assumptions:

- A small open economy consists of two sectors – the tradable (T) and the non-tradable (NT) one.
- The price of tradable goods, as well as the price of capital are set in international markets and hence exogenous from the point of view of the analyzed small economy.
- The capital is perfectly mobile between sectors and regions.
- The labour force is perfectly mobile between sectors but immobile between regions. Cross-sectoral labour mobility should imply equality of wages between sectors 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.
- There is perfect competition in both sectors (in both regions).
- Technology in both sectors is described by Cobb-Douglas production functions (for algebraic simplicity) with constant returns to scale $Y_t = A_t L_t^{\alpha} K_t^{1-\alpha}$, with Y_t denoting output at time t, A_t – total factor productivity, L_t – labour input K_t – capital input. α and $1-\alpha$ denote labour and capital elasticities of output, respectively.

2.1 Baumol-Bowen effect

The Baumol-Bowen effect explains cross-sector inflation differential by means of divergent productivity dynamics between T and NT sectors.

To see this, assume that producers of tradable (T) and nontradable (N) goods face the analogous technologies:

$$Y_T(K_T, L_T) = A_T K_T^{1-\alpha_T} L_T^{\alpha_T} \tag{1}$$

$$Y_N(K_N, L_N) = A_N K_N^{1-\alpha_N} L_N^{\alpha_N} \tag{2}$$

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

$$\max_{L_T, K_T} \quad P_T A_T K_T^{1-\alpha_T} L_T^{\alpha_T} - w_T L_T - r K_T \tag{3}$$

$$\max_{L_N, K_N} \quad P_N A_N K_N^{1-\alpha_T} L_N^{\alpha_T} - w_N L_N - r K_N \tag{4}$$

First order conditions for the above maximization problems (with respect to labour) are the following:

$$\alpha_j P_j A_j K_j^{1-\alpha_j} L_j^{\alpha_j - 1} - w_j = 0 \quad j \epsilon \{T, N\} \tag{5}$$

This implies the prices of labour as follows:

$$w_T = \alpha_T P_T A_T (\frac{L_T}{K_T})^{\alpha_T - 1} \tag{6}$$

$$w_N = \alpha_N P_N A_N (\frac{L_N}{K_N})^{\alpha_N - 1} \tag{7}$$

Assuming wage homogeneity across the sectors, $w_T = w_N \equiv w$, and using (6) and (7), we obtain:

$$\alpha_T P_T A_T (\frac{L_T}{K_T})^{\alpha_T - 1} = \alpha_N P_N A_N (\frac{L_N}{K_N})^{\alpha_N - 1} \tag{8}$$

Equation (8) can also be expressed as a formula for relative price of nontradable versus tradable production:

$$\frac{P_N}{P_T} = \frac{\alpha_T A_T (\frac{L_T}{K_T})^{\alpha_T - 1}}{\alpha_N A_N (\frac{L_N}{K_N})^{\alpha_N - 1}} = \frac{\alpha_T A_T L_T^{\alpha_T} K_T^{1-\alpha_T} L_T^{-1}}{\alpha_N A_N L_N^{\alpha_N} K_N^{1-\alpha_N} L_N^{-1}} = \frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}} \tag{9}$$

Let lowercase letters denote the natural logarithms of their uppercase counterparts. Let $l_j \equiv ln\left(\frac{Y_j}{L_j}\right)$ denote labour productivity in sector $j \epsilon \{T, N\}$. Taking logs of (9) yields

$$p_N - p_T = \alpha_T - \alpha_N + l_T - l_N \tag{10}$$

or, alternatively in log differences (versus previous period), (10) can be expressed as

$$\dot{p}_N - \dot{p}_T = \dot{l}_T - \dot{l}_N \tag{11}$$

with $\dot{x}$ denoting the growth rate (log-difference) of variable x.

Both (10) and (11) summarize the resulting relationship between relative prices and relative productivity between the tradable and nontradable sector, i.e. the Baumol-Bowen effect.

2.2 Balassa-Samuelson effect

The Balassa-Samuelson effect is an international extension of the Baumol-Bowen model. It describes the cross-country consequences of divergent productivity dynamics, expressed in terms of inflation differentials and the real exchange rate. To discuss this issue, let us denote the foreign counterparts to domestic variables with an asterisk * in superscripts.

Dividing (9) by its foreign counterpart leads to the following relationship, expressing relative price of nontradable goods in international comparison as a function of relative productivities in both sectors, both at home and abroad:

$$\frac{\frac{P_N}{P_T}}{\frac{P_N^*}{P_T^*}} = \frac{\frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}}}{\frac{\alpha_T^* \frac{Y_T^*}{L_T^*}}{\alpha_N^* \frac{Y_N^*}{L_N^*}}} \tag{12}$$

After taking log-differences of (12) we arrive at the following equation:

$$(\dot{p}_N - \dot{p}_T) - (\dot{p}_N^* - \dot{p}_T^*) = (\dot{l}_T - \dot{l}_N) - (\dot{l}_T^* - \dot{l}_N^*) \tag{13}$$

Define the aggregate price level at home P as a geometric average of tradable and nontradable prices, with δ being the share of the tradable sector in the home economy:

$$P \equiv P_T^\delta P_N^{1-\delta} \tag{14}$$

Taking log-differences of (14) yields

$$\dot{p} = \delta \dot{p}_T + (1-\delta)\dot{p}_N \tag{15}$$

Define the real exchange rate Q as

$$Q \equiv E\frac{P^*}{P} \tag{16}$$

Log-differencing (16) and using (15) (as well as its foreign counterpart) we obtain:

$$\dot{q} = \dot{e} + \dot{p}^* - \dot{p} = \dot{e} + \dot{p_T}^* - \dot{p_T} - (1-\delta)(\dot{p}_N - \dot{p}_T) + (1-\delta^*)(\dot{p}_N^* - \dot{p}_T^*) \tag{17}$$

Relative price dynamics can be replaced with productivity, according to (11), which finally leads to the real exchange rate dynamics as a function of relative productivity dynamics:

$$\dot{q} = \dot{e} + \dot{p_T}^* - \dot{p_T} - (1-\delta)\left(\dot{l}_T - \dot{l}_N\right) + (1-\delta^*)\left(\dot{l}_T^* - \dot{l}_N^*\right), \tag{18}$$

which is the Balassa-Samuelson effect with respect to real exchange rate.

If we assume that the purchasing power parity hypothesis holds in the tradable sector ($\dot{e} = \dot{p_T} - \dot{p_T}^*$) and that both sectors at home and abroad are symmetrically sized ($\delta = \delta^*$), formula (17) collapses to:

$$\dot{q} = -(1-\delta)[(\dot{l}_T - \dot{l}_N) - (\dot{l}_T^* - \dot{l}_N^*)] \tag{19}$$

Note that the real appreciation, implied by the right-hand side of equation (18) can be channeled in two ways. Firstly – via P (price level at home, composed by P_T and P_N) or via E (the nominal exchange rate). We assume here that a small open economy cannot influence the price level abroad, P^*. However, once the home and foreign economy share a common currency, the only possibility to appreciate Q is to raise P, as E is irrevocably fixed. This is the case when one of the NMS small economies integrates with the euro area.

2.3 Derivation under non-homogeneity of wages across sectors

Equation (8) was derived from first order conditions for producer maximization problems (3)-(4) under the assumption that wages are equal across sectors. Should this assumption be rejected, we proceed by dividing (6) by (7). After rearrangements, this leads to an analogue of (8), augmented with relative wages across sectors:

$$\frac{P_N}{P_T} = \frac{\alpha_T \frac{Y_T}{L_T}}{\alpha_N \frac{Y_N}{L_N}} \cdot \frac{w_N}{w_T} \tag{20}$$

The log and log-differenced version of (20) are, respectively,

$$p_N - p_T = (\alpha_T - \alpha_N) + (l_T - l_N) + (w_N - w_T) \tag{21}$$

$$\dot{p}_N - \dot{p}_T = \left(\dot{l}_T - \dot{l}_N\right) + (\dot{w}_N - \dot{w}_T) \tag{22}$$

The above equations generalize (10) and (11) to the case of non-homogenous wages across sectors. Following the steps (12) to (18) in a similar way, we finally arrive at an analogue of (18) which is

$$\dot{q} = -(1-\delta)\left[\left(\dot{l}_T - \dot{l}_N\right) - \left(\dot{l}_T^* - \dot{l}_N^*\right) + (\dot{w}_N - \dot{w}_T) - (\dot{w}_N^* - \dot{w}_T^*)\right]. \tag{23}$$

3 Empirical framework

Before turning to the estimation results, we first outline the empirical setup applied here. This includes (i) the mapping between the theoretical derivation of both BB and BS effects and the formulation of the equations that we estimate, (ii) description of data sources and definitions and (iii) technical aspects associated with the use of panel econometric methods.

In the literature, BB and BS equations are normally estimated via either pure time-series methods or panel methods with various countries as cross-sectional dimension of the panel. Both approaches have obvious drawbacks. Available time series for single NMS are still far too short to ensure efficient estimation of long-run relationships, and – working with annual data – some of them are virtually unavailable. Consequently, the relatively short time span and low frequency of the available series necessitates the use of panel econometrics techniques. In the literature the small sample problem is addressed by extending the analysis from one country to the group of relatively homogenous economies. However, turning to a multi-country panel alters the interpretation of the results. It is also unsatisfactory when we focus on a single country's policy objectives.

Our attempt to overcome this difficulty consists in designing the panel in a different manner. Namely, we propose to use a multi-sector decomposition of the economy to design a panel in which the relation between the tradable sector and various branches of the nontradable sector will serve as the unit dimension. This enables to concentrate on the Polish economy exclusively and, at the same time, improve the efficiency of the estimation.

3.1 Data source and definitions of variables

The data used in the analysis come from the Eurostat database. The sample covers years 1995 through 2008 and is of annual frequency. The source variables comprise sectoral[2] value added deflators as an approximation for price developments, sectoral labour productivity (value added over total employment), sectoral wages (compensation of employees over total employment) and the real exchange rate (deflated by GDP or value added in manufacturing). The application of NACE-based statistical concepts (value-added deflators instead of price indices) asserts the coherence of sectoral classification.

The cross-sectional dimension of the data is obtained either by means of sectoral disaggregation, i.e. the price, productivity and wage differentials are computed as a difference between the aggregated tradable sector and each non-tradable subsector, or country disaggregation, i.e. we include the real exchange rate of PLN against the incumbent euro area member states (without Ireland and Austria because of incomplete data).

Table 1 contains the definitions of the variables used in the empirical analysis.

[Table 1 about here]

Using various proxy variables and specifications, we end up with three alternative formulations of the BS effect equation:

1. Firstly, we use a multi-sectoral panel of cross-country inflation (or price level) differentials between Poland and the euro area, $p^{diff}_{pl\_ea}$, as the dependent variable, explained by the respective differential productivity dynamics (or levels) .
2. Secondly, we use the GDP-deflated real exchange rate (dynamics), Q_{GDP}, as implied by equation (18), explained with productivity index (or dynamics) in line with this equation, $l^{diff}_{pl\_members}$.
3. Thirdly, we refine the second approach by using $l^{*diff}_{pl\_members}$ instead of $l^{diff}_{pl\_members}$ as the explanatory variable, i.e. accounting for differences in size of T vs. NT sector in Poland and individual countries in the euro area.

3.2 Sectoral classification

The sectoral classification we decided on (Table 2) compromises 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. Although the products of this sub-sector are subject to international trade, both their prices and quantities are heavily distorted by administrative interventions (on both country- and the EU-level).

[Table 2 about here]

3.3 Methodological notes

3.3.1 Panel unit root tests

The inference on the stationarity of the analysed series is based on the panel unit root tests which allow for heterogeneity across the cross-sectional dimension (in terms of the autoregressive coefficient and the number of dependent variable lags in the test regression), namely the Im, Pesaran and Shin test (IPS, Im et al. 2003) and Fisher augmented Dickey-Fuller and Phillips-Perron tests (Maddala and Wu, 1999). The individual unit root processes assumption reduces the unobserved heterogeneity problem and, according to Monte Carlo simulations', results in higher power of the tests compared to those based on the supposition of common persistence parameter across cross-sectional units.

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

$$H_0 : \alpha_i - 1 = 0, \tag{24}$$

where α_i denotes the persistence parameter, while the alternative hypothesis is given by:

$$H_1 : \alpha_i - 1 < 0 \tag{25}$$

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.

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:

$$\Delta y_{i,t} = (\alpha_i - 1)y_{i,t-1} + \sum_{k=1}^{K_i} \beta_{ik} \Delta y_{i,t-k} + \varepsilon_{i,t} \tag{26}$$

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

$$\bar{t} = \frac{1}{N} \sum_{i=1}^{N} t_{\alpha_i} \tag{27}$$

The possible problem of the cross-sectional dependence of the test regressions' residuals is utilized by demeaning of the IPS statistics. The demeaned standardized average t-statistics calculated according to the following formula

$$\sqrt{N}\frac{(\bar{t}-\frac{\sum_{i=1}^{N}E(t_i)}{N})}{\sqrt{\frac{\sum_{i=1}^{N}Var(t_i)}{N}}} \tag{28}$$

has an asymptotic (with $N \to \infty$) standard normal distribution.

Owing to the fact that the pair-wise correlation coefficients of the disturbances may differ across the individual units, the de-meaning may be insufficient to eliminate the cross-sectional dependence problem. For this reason (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:

$$-2\sum_{i=1}^{N}log(p_i), \tag{29}$$

where p_i denotes the p-value from the individual unit root test. In the limit (with $T \to \infty$) , the test has a χ^2_{2N} distribution. The Fisher statistics is applied both to the ADF and PP test.

The inclusion of tests with asymptotic properties relying on the cross-sectional (the IPS test) and temporal (the Fisher-type tests) dimension, respectively, serves for the purpose of asserting the robustness of the results.

3.3.2 Panel cointegration tests

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

The Pedroni test consists in applying a unit root test to the residuals of the regression of the analysed variables. 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 data set we confine our attention to mean group statistics. In the case of those statistics the cointegration equation

$$y_{i,t} = \alpha_i + \delta_i t + \beta_i x_{i,t} + \varepsilon_{i,t} \tag{30}$$

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. For this reason 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:

$$\Delta\hat{\varepsilon}_{i,t} = (\alpha_i - 1)\hat{\varepsilon}_{i,t-1} + \sum_{k=1}^{K_i} \beta_{ik}\Delta\hat{\varepsilon}_{i,t-k} + \vartheta_{i,t}, \tag{31}$$

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

$$\tilde{Z}_{ADF} = \sum_{i=1}^{N} (\hat{s}_i^2 \sum_{t=1}^{T} \hat{\varepsilon}_{i,t-1}^2)^{-\frac{1}{2}} \sum_{t=1}^{T} (\hat{\varepsilon}_{i,t-1}\Delta\hat{\varepsilon}_{i,t}), \tag{32}$$

where $\hat{s}_i^2 = \frac{1}{T}\sum_{t=1}^{T} \hat{\vartheta}_{i,t}^2$.

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

$$-2\sum_{i=1}^{N} log(p_i) \rightarrow \chi^2_{2,N} \tag{33}$$

3.3.3 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 on 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:

$$\begin{cases} y_{it} = \alpha_i + \beta x_{it} + \mu_{it} \\ x_{it} = x_{it-1} + \varepsilon_{it} \end{cases} \quad i = 1, ..., N \tag{34}$$

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.

$$\Omega_i = \sum_{k=-\infty}^{\infty} \Gamma_i^k = \Gamma_i^0 + \sum_{k=1}^{\infty}(\Gamma_i^k + \Gamma_i^{k\,T}), \tag{35}$$

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:

$$\hat{\Omega}_i = \hat{\Gamma^0}_i + \hat{\Gamma}_i + \hat{\Gamma}_i^T, \tag{36}$$

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

$$\hat{\Omega}_i = \hat{L}_i \hat{L}_i^T, \tag{37}$$

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

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

$$y_{it}^* = y_{it} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} \Delta x_{it}, \tag{38}$$

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

$$\hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Gamma^0}_{21i} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}}(\hat{\Gamma}_{22i} + \hat{\Gamma^0}_{21i}) \tag{39}$$

The corrections are applied to the OLS estimator in the following manner:

$$\hat{\beta}^{FMOLS} = \frac{1}{N}\sum_{i=1}^{N}(\sum_{t=1}^{T}(x_{it} - \bar{x}_i)^2)^{-1}(\sum_{t=1}^{T}(x_{it} - \bar{x}_i)y_{it}^* - T\hat{\gamma}_i) \tag{40}$$

and the t-statistics for the long-run coefficient β 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:

$$y_{it} = \alpha + \beta x_{it} + \sum_{p=-P}^{P} \delta_p \Delta x_{it-p} + u_i + \varepsilon_{it} \tag{41}$$

The estimator is given as:

$$\hat{\beta}^{DOLS} = \frac{1}{n}\sum_{i=1}^{n}(\sum_{t=1}^{T} z_{it}z_{it}^{'})^{-1}(\sum_{t=1}^{T} z_{it}y_{it}), \tag{42}$$

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

4 Empirical results

In this section we report the results of the empirical investigation of the existence and the magnitude of the Baumol-Bowen and Balassa-Samuelson effects in the Polish economy. The presentation of the estimates and the quantification of the effects is preceded by the analysis of the variables' properties that could possibly shed some light on the validity of model's assumptions and hence on the interpretation of the results.

4.1 Validity of assumptions

The decision on the empirical specifications of the Baumol-Bowen and Balassa-Samuelson equations is conditonal upon the validity of the theoretical model assumptions. For this reason, in the beginning of the empirical investigation we assess the validity of the two underlying suppositions – the wage homogeneity and the prevalence of purchasing power parity (PPP) in the tradable sector. The empirical verification of these hypotheses consists in applying the stationarity test to either relative wages or real exchange rate (Poland vs. each euro area member state) deflated by the price index of value added in manufacturing. The stationarity test is a weak econometric formulation of wage homogeneity and PPP hypothesis, as it allows for substantial and persistent differences in the level of sectoral wages or price levels in the tradable sector of individual countries.

Both IPS and Fisher ADF tests clearly reject the null hypothesis of non-stationarity of real exchange rate deflated by the deflator of gross value added in manufacturing (Table 3). However, the Fisher PP test points out to the unit root in the data generating process. Also, the conclusions of two previous tests are mainly due to a strong effect of economic downturn in the final year of the sample and the resulting nominal depreciation. Owing to this result and the numerous theoretical arguments[3] and empirical investigations in the literature, we assume that purchasing power parity does not hold for the tradable sector.

As a result, we cannot skip the real exchange rate deflated by a proxy of tradable price deflators when moving from (18) into (19) and include this term on the right-hand side of the estimated equations.

The assumption of wage homogeneity also seems not to be fulfilled, although – in some cases – by a slight margin. For this reason, we additionally estimate the augmented specifications of the model with sectoral wage differentials, based upon equations (22) and (23) instead of (11) and (18), respectively. In the following Subsections, we report both sets of results (i.e. with and without wage homogeneity assumed), but – given majority of the tests provided – we tend to treat the results derived with rejected wage homogeneity assumption as more reliable. There are, however, only marginal differences in terms of final BB and BS effects quantification.

[Table 3 about here]

Recall that transforming equation (18) into (19), we assumed that sector sizes are equal across home and foreign economy. However, owing to the fact that there are substantial differences in the share of the tradable sector in the economy (value added in manufacturing over the overall gross value added) in Poland and in the euro area member states, we also correct for the difference in the sectoral composition of the economies in the real exchange rate equations. This technical correction strengthens the interpretation of the results, moving away the considerations of variable scaling and its influence on the magnitude of the estimated coefficients.

4.2 Stationarity testing

All the "level" variables (difference in log-indices) seem to be non-stationary (Table 4), which allows us to apply panel cointegration techniques and explore the long-run relationships (LR) between price-level and productivity differentials. The "growth rate" variables (difference in growth rates) are all stationary. The only possibly vague case here is the real exchange rate deflated by the GDP deflator, for which the 3 tests applied indicate various conclusions. In line with a similar situation of $Q_{manufacturing}$ in Subsection 4.1, and taking economic plausibility considerations into account, we also conclude nonstationarity of Q_{GDP}.

[Table 4 about here]

4.3 Cointegration testing

In the next step, the existence of cointegrating relationships between the I(1) variables needs to be examined. Should the relationships be confirmed, one will be able to proceed to estimation and interpret the estimates for levels as long-run equations describing Baumol-Bowen and Balassa-Samuelson effects.

In the case of the price differentials-productivity differentials relationships (both within the Polish economy and between Poland and the euro area) as well as in the case of wage-augmented equation for the Polish economy, both the Pedroni ADF statistics and Fisher-Johansen statistics clearly indicate the existence of a long-run equilibrium (Table 5). This result implies that the existence of wage-augmented long-run BS effect seems to be backed only by the Fisher-type statistics. At the same time, the cointegration of variable set corresponding to BS effect without correction for nonhomogeneity of wages (column 2), as well as both versions of the variable set representing the Baumol-Bowen effect (column 1 and 3), is confirmed by both tests applied here.

In the case of trivatiate systems the eigenvalue analysis suggests the existence of two cointegration vectors. This would imply that the estimated parameters of the single long-run equation with all three variables (which is the only possibility given the FMOLS and DOLS estimators) could be merely a linear combination of the "true" cointegration vectors and, therefore, should be interpreted with caution.

[Table 5 about here]

A more nuanced picture emerges from both test statistics applied to the Balassa-Samuelson equations with the real exchange rate as the dependent variable. For all variable sets under consideration, the ADF group statistics does not reject the null hypothesis of no cointegrating relations within this set (see Table 6). This result is contradicted by the Johansen-Fisher tests. Both the trace and the maximum eigenvalue statistics strongly reject the null of zero cointegrating relations in favour of at least one. Also, both versions of the test suggest the existence of two cointegrating equations, consistently across the variable sets.

In the absence of a clear conclusion whether (and how many) cointegrating relations exist, we proceed estimating a single equation, which is the only feasible option given FMOLS and DOLS estimators. This is in line with the discussion of variable sets presented in Table 5, whereby even more caution will be needed when interpreting the results.

[Table 6 about here]

Having confirmed the I(1)-ness and the cointegration of the "level" variables, we proceed by presenting the estimation results. We report multiple estimates, along the following dimensions:

- short-run and long-run estimates, depending whether the coefficients are estimated in the equation in log-differenced variables or as long-run relations in log-levels;
- with and without wage homogeneity assumption, according to the results in Subsection 4.1;
- with different proxies for dependent and independent variables in the case of the BS effect, as discussed in Subsection 3.1;
- estimation method in the case of long-run estimates.

4.4 Short-run estimates

The short-run estimates are obtained on the basis of the equations specified on the differences in growth rates of the variables. The estimates of β parameter in all the empirical specifications are statistically significant, albeit substantially less than unity – contrary to the prediction of the theoretical model.

The internal mechanism (BB effect) appears to be relatively week. According to the estimation results, the increase in the difference between productivity growth rate in the tradable and non-tradable sector by 1 percentage point translates on impact merely into 0.14-0.17 percentage point raise in the relative inflation, depending on whether wage homogeneity assumption is relaxed or not.

The external mechanism (BS effect) seems to be stronger, judging merely by the estimated coefficients. Namely, the increase in relative productivity differential growth rate in Poland *versus* the euro area results in 0.18-0.21 hike in dual inflation differential. The magnitude is even higher when we look at the estimates of equations with the real exchange rate (GDP-deflated) as a dependent variable. The estimated real exchange rate appreciation (on impact) due to a 1 p.p. differential in sectoral productivity growth rates, in relation to the euro area countries, ranges from 0.52 to 0.62 p.p. These results seem to be insensitive to the selection of proxy variables that account or not for the cross-country differences in sector sizes.

A comparison between the coefficients in the BS equation for dual inflation dfferential and real exchange rate appreciation suggests that there is a considerable discrepancy between the two estimates. One of the possible explanations is that a significant portion of relative productivities' impact on the real exchange rate was channelled via the nominal interest rate, which is absent from the left-hand side of the BS equation on inflation differentials. Another difference between the two approaches is the definition of cross-sectional units in both panels. In the real exchange rate equations, the units are defined as country pair, whereas in the inflation differential equations – as sectors. In this case, the difference in the results could be explained by the presence of either cross-country or cross-sector heterogeneity in the strength of the effect.

Note that these results capture only the transmission of the relative productivity growth to relative price growth *on impact*, i.e. within the same time period. Although the relatively low, annual frequency implies that a significant portion of the adjustment process might be taking place in the same period, we cannod exclude the existence of some lagged adjustment that could be captured in the long-term specification.

[Table 7 about here]

4.5 Long-run estimates

Having established the cointegration relationships between differential price levels and productivity (as well as wages), we can estimate the long-run relationships by means of the FMOLS and DOLS estimator. Table 8 presents the long-run versions of the estimates in Table 7.

[Table 8 about here]

All the variables are significant in all the equations and the parameters are signed in line with theoretical priors. Unlike the short-run results, in all the long-run equations the estimates obtained from the wage-augmented equations are lower than their reduced-form counterparts and all the wage terms are highly significant. It seems therefore that in the Polish economy the productivity-induced inflation pressure is mitigated in the long run by lower wage growth in the non-tradable sector.

The estimated long-run impact of 1% relative productivity growth on relative prices (Baumol-Bowen effect) ranges from 0.60% to 0.65%, depending on the estimator and cross-sectoral wage homogeneity assumption. The Balassa-Samuelson effect ranges from 0.33% to 0.55% in response to a 1% growth of relative productivity when we consider the relative prices as a dependent variable, and from 0.55% to even 0.86% when we take into account the GDP-deflated real exchange rate. However, taking into account our previous results of stationarity testing (non-homogeneity of wages), we could expect the lower bound to be relatively more plausible as a value range for the true parameter.

The results seem to be relatively robust to the choice of the estimation method. Also, the construction of the relative productivity proxy in the real exchange rate equation does not affect the estimates in a considerable manner. However, in line with the short-term results for the BS effect, the coefficients for relative productivities are generally higher in absolute terms when the real exchange rate is the dependent variable rather than relative prices. Consequently, the possible explanation for this discrepancy also applies to the long-run conclusions.

Finally, we specify an error correction model (see Table 9), comprising the short-run formulation from Table 7 and an error correction term, i.e. the lagged residual of the corresponding cointegration regression in Table 8. Again, the short-run parameters in question are correctly signed and significant in both the Baumol-Bowen and Balassa-Samuelson equations. What is more, the relative wage dynamics is significant in the equations with inflation on the left hand side. In all the specifications, the error correction parameter is significantly lower than 0 and ranges from -0.22 to -0.53. The relatively strong error corrections are consistent with the annual frequency of the data and imply half-life parameters from 0.9 to 2.8 years.

[Table 9 about here]

More importantly, the extension of the short-run model to error correction specification has allowed to obtain estimates of β that are more robust across model specifications and estimation methods. The estimated, short-run impact of additional 1 p.p. relative productivity growth on relative price growth across sectors (Baumol-Bowen equations) ranges between 0.19 and 0.21 p.p. Relative productivity growth between Poland and the euro area of the same magnitude leads to an increase in relative inflation differential of 0.20 to 0.24 and a real appreciation of 0.28 to 0.43, depending on the specification and estimation method.

4.6 Demand side effects

The Baumol-Bowen and Balassa-Samuelson effects hinge upon the assumption of full utilization of production factors. This leads to a supply-side based explanation of inflation and real exchange rate developments. However, in the short run the demand side effects are potentially more important in this respect. The higher pace of non-tradables' price growth might, namely, result from a positive income elasticity of this sector's products, especially services. The rationale behind this hypothesis is that in a catching-up economy (like Poland in the sample period), consumers shift their demand from tradable to non-tradable sector as they become richer. This could explain the relative price dynamics instead of productivity differentials. Therefore, we check for the presence of those phenomena by augmenting the short-run equations (Tables 7 through 9) with a regressor being a proxy for demand-side developments, i.e. GDP *per capita* (Table 10).

[Table 10 about here]

In most cases GDP per capita is significant and correctly signed. What is more, the estimates of β parameter obtained from the GDP-augmented equations are lower than their counterparts from a purely supply-driven specifications, which additionally supports the existence of demand-side effects. However, this additional control variable does not substantially affect the estimation results and in most cases the productivity-differential terms are still highly significant.

4.7 Quantification of the effects

Both short-run and long-run results can be seen as a confirmation of the existence of Baumol-Bowen and Balassa-Samuelson effect in the Polish economy. The question now is how strong both effects are in quantitative terms, i.e. how many percentage points did they add to Polish inflation rate and to the real appreciation in the sample period.

The quantification of the Baumol-Bowen effect is given by the product of (1) the estimated coefficient, corresponding to the differential productivity variable, and (2) the average value of this variable over the sample period, and (3) the share of the non-tradable sector in the economy[4].

According to the short-run estimates, the magnitude of Baumol-Bowen effect in the Polish economy (as a contribution to Polish inflation rate) amounted to $0.7 - 1.0$ percentage points *per annum* on average in the sample period (0.6-0.9 in the shorter sub-sample 1999-2008). This is relatively small, compared to the average growth rate of inflation[5] in this period, which amounted to 6.0% (2.0% in years 1999 through 2008; see Table 12). The long-run estimates of the Baumol-Bowen effect are of higher magnitude: 2.8-3.0 contribution to country-specific inflation (2.6-2.7 in years 1999 through 2008). These results are also too low to explain the average difference in log-indices of price levels between non-tradable sub-sectors and the tradable sector.

In line with expectations, the results obtained for the more recent sub-sample are lower than for the entire sample. This is a straightforward consequence of the dampened trend in Polish non-tradable sector's relative productivity. Its growth rate has been gradually decreasing over the last decade and one can probably expect the Baumol-Bowen effect to stay at or below the lower bound of the estimates for the subsample 1999-2008. On the other hand, in the shorter sub-sample the relative contribution of this effect to Polish inflation was much higher and amounted almost to 50%.

The long-run estimates clearly outperform the short-run impact. This can be explained in at least 2 manners. Firstly, the relative productivity shifts are not immediately mirrored in relative price developments, but they also continue to affect price indices in the subsequent years. This is additionally confirmed by the significance, correct sign and reasonable magnitude of error correction parameters in the error correction models. One possible explanation for that are labour and product market rigidities. Secondly, the short-run specification might underestimate the parameter for econometric reasons. If the relative productivity and relative price growth are relatively stable and smooth processes, the stable relationship between annual growth rates on both sides of the estimated equation can be captured by the constant to a dominant extent.

The Balassa-Samuelson effect can be quantified in a very similar fashion, i.e. as a product of the parameter β of the respective relative productivity level (or dynamics) and the average of this relative productivity dynamics in the sample period.

This calculation leads us to an estimate of 0.8-1.0 additional percentage point in differential between Poland's and euro area's relative cross-sectoral price dynamics and 1.5-2.2 additional percentage point in short-run real exchange rate appreciation that can be attributed to Balassa-Samuelson effect. If we consider the error correction specification, this short-run contribution can be limited to 1.0-1.3 percentage points of real exchange rate appreciation (while the contribution to relative inflation rate remains broadly unchanged). In the long run, the estimates of the effect range from 1.5 to 2.1 (percentage point contribution to relative inflation rate) or 1.4-2.3 (contribution to real exchange rate appreciation), which is in line with generally higher estimates of β in the cointegrating relations than in difference equations.

Like in the case of Baumol-Bowen effect, the results for the sub-sample 1999-2008 are lower than in the entire sample 1995-2008 because the productivity growth differentials were more moderate towards the end of the sample. Narrowing the sample limits the estimated short-run contribution of BS to real appreciation to 1.2-1.7 percentage point (0.8-1.1 in the ECM version). In the long run, this contribution amounts to 1.4-2.3 percentage points. All these intervals are narrower and lie closer to zero than their counterparts based on the sample 1995 through 2008. The only exception is the estimated contribution to relative inflation, which is relatively robust with respect to sample length.

[Table 11 about here]

[Table 12 about here]

Also, we note that the long-run estimates of the BS effect depend to a large extent on whether we accept or reject the wage homogeneity assumption. These estimates are systematically lower in absolute terms (0.4 to 0.6 percentage point) when we reject this assumption, which seems to be more plausible given the outcomes of stationarity tests (Table 3). Taking into account the estimates without wage homogeneity assumption only, we could narrow the estimated range of long-run BS effect magnitudes to 1.8-2.0 percentage point of annual real exchange rate appreciation in the longer sample or even 1.4-1.6 in the shorter sample. The estimated BS effect measured with contribution to relative inflation would decline from 2.1 to 1.5 percentage point.

[Table 13 about here]

These results explain more of the respective cross-region and cross-sector price level differential than in the case of Baumol-Bowen effect. The exact assessment depends, however, on the horizon of the analysis and the choice of the dependent variable. Taking difference between differential price levels (non-tradables vs. tradables) in Poland and the euro area with its annual growth rate of 4.9% (3.8% in the shorter sample), both the short-run and the long-run estimates of BS effect are far too low to account for this. However, when we consider the GDP-deflated real exchange rates instead (appreciation of 1.8% p.a. over the period 1995-2008 and 1.1% in the more recent sub-sample), the estimated BS contributions to Poland's real appreciation are of comparable magnitude.

However, in general, these results are insufficient to conclude that the Balassa-Samuelson effect could be a dominant contributor to the Polish real appreciation observed in the sample period.

5 Conclusions

This paper revisits the Baumol-Bowen and Balassa-Samuelson effects in the Polish economy. Both mechanisms are of highest interest for policymakers. Poland, as a country with derogation, needs to take steps to adopt the euro, which requires i.a. to fulfil the price stability criterion. This stability will be assessed in comparison with three best-performing EU countries. In this context, the factors responsible for low-frequency inflation movements, which are specific for catching-up economies, should be investigated in detail, quantified and compared with the admissible difference of 1.5 percentage point between Polish and EU best performers' consumer inflation rate.

The empirical strategy adopted here uses techniques of panel econometrics. For the assessment of Baumol-Bowen effect, we propose a novel approach that defines the spatial dimension of the panel as individual sectors of the economy. The unit dimension contains variables defined in relative terms between single tradable sector (manufacturing) and various non-tradable branches, according to NACE rev. 1.1 classification. This is also the case in one of the equations testing the Balassa-Samuelson effect. To verify the latter effect, variables are expressed in relative terms between Poland and most of the EA-12 countries. We provide both short-run and long-run estimates, using alternative specifications, proxies and estimation methods (including fully-modified OLS and dynamic OLS for panel cointegration).

The estimated historical contribution of the Baumol-Bowen effect to Polish inflation rate is $0.7-1.0$ percentage points in the short run and 2.8-3.0 in the long run. These results are slightly lower when we consider only the average relative productivity dynamics in a more recent sub-sample. These results are broadly in line with a relatively broad spectrum of estimates in the literature, although the short-run estimates are close to the lower bound of this range. Moreover, most of the previous literature did not provide explicit differentiation between short-run and long run effects.

The Balassa-Samuelson effect is quantified in two manners: (i) as a contribution to annual difference in relative price level growth (non-tradables vs. tradables) in Poland and the euro area and (ii) as a contribution to GDP-deflated annual real exchange rate appreciation, both in terms of average over the sample period. The results, respectively, amount to (i) 0.8-1.0 p.p. (short-run) and 1.5-2.1 p.p. (long-run) (ii) 1.0-2.2 (short-run) and 1.8-2.8 (long-run). However, when we focus on the more recent subsample 1999-2008 and the most plausible specifications (error-correction model, without cross-sector wage homogeneity assumption), we could narrow these ranges to (i) 0.9 (short-run) and 1.3 (long-run) (ii) 0.9-1.1 (short-run) and 1.4-1.6 (long-run).

The above results suggest that the Baumol-Bowen and Balassa-Samuelson effects should be treated by policymakers as a non-negligible issue in the context of Poland's integration with the euro area, but not as an obstacle. One needs to stress that the results discussed above are historical and their direct extrapolation into the future would be misleading. The productivity gap between Poland and the euro area has been trending down over the last decade, along with productivity growth rate differential. Hence, the estimates discussed here – even for the sub-sample 1999-2008 – should be treated as an upper bound for analogous estimates in the future rather than a benchmark.

On the other hand, the estimated impacts of BS effect on relative price growth are significant, compared with the feasible difference of 1.5 percentage point between 12-month average annual HICP growth rate in Poland and 3 'best performers' in the EU. In particular, even when the pressure on real appreciation against the euro area within ERM II is channelled fully through the domestic price growth, an annual appreciation of arount 1% would leave relatively little room for manoeuvre to policymakers if there are clearly outstanding countries in the reference group for evaluating this criterion.

Finally, the analysis does not seem to provide strong evidence against Poland's ability to maintain competitiveness after the integration with the euro area. The estimated historical impacts of BS effect on relative inflation rates are comparable, and in many cases even lower, than cross-country inflation differentials between euro area countries over the first decade of the common currency. Moreover, the additional price growth would mainly be concentrated in the non-tradable sector.

Notes

[1] Former Articles 121-123 of the Treaty establishing the European Community.

[2] According to NACE rev. 1.1 classification.

[3] E.g. pricing-to-market practices, difference in the quality of goods consumed at home and abroad, local consumers' tastes, local non-tradable inputs in tradable goods, differences in tax systems.

[4] To see this, rearrange equation (11) to $\dot{p} = (1-\delta)(\dot{p}_N - \dot{p}_T) + \dot{p}_T$ and substitute the right-hand side of (15). In the resulting expression, $\dot{p} = (1-\delta)\left(\dot{l}_T - \dot{l}_N\right) + \dot{p}_T$, treat $\dot{p}_T = 0$ as a 'numeraire'. This allows to interpret the result as the contribution of Baumol-Bowen effect to overall inflation rate, provided that we multiply the productivity growth differential by the non-tradable sector size.

[5] By inflation we mean artificial value-added deflator, composed solely of the NACE sectors C through O.

References

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, 607–629.

Breuss F. (2003): *Balassa-Samuelson Effects in the CEEC: Are they Obstacles for Joining the EMU?*, Wirtschaftsuniversit at Wien IEF Working Paper, 52.

Chmielewski T. (3003): *Znaczenie efektu Balassy-Samuelsona dla polskiej polityki pieniężnej*, Materiały i Studia NBP, 163.

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

Corsetti G., Dedola L. (2005): *A macroeconomic model of international price discrimination*, Journal of International Economics, 67, 129–155.

Égert B. (2002a): *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. (2002b): *Investigating the Balassa-Samuelson hypothesis in transition: Do we understand what we see?*, Economics of Transition, 10(2), 273–309.

European Commission (2009): *Competitiveness developments within the euro area*, Quarterly Report on the Euro Area, 8(1).

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

García-Solanes J., Sancho-Porteroa F.I., Torrejón-Floresa F. (2008): *Beyond the Balassa-Samuelson effect in some new member states of the European Union*, Economic Systems, 32(1), 17–32.

Halpern L., Wyplosz C. (2001): *Economic Transformation and Real Exchange Rates in the 2000s: The Balassa-Samuelson Connection*, Economic Survey Of Europe, Chapter 6, United Nations Economic Commission for Europe.

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

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

Lipińska A. (2008): *The Maastricht Convergence Criteria and Optimal Monetary Policy for the EMU Accession Countries*, ECB Working Paper.

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

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, 631–652.

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

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, 597–625.

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

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

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

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

Łukasz Rawdanowicz (2002): *Badanie wpływu efektu Harroda-Balassy-Samuelsona na ceny relatywne w Polsce*, Fundacja CASE.

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

Wagner M., Hlouskova J. (2004): *What's Really the Story with this Balassa-Samuelson Effect in the CEEC?*, Universit at Bern Volkswirtschaftliches Institut Diskussionsschriften, 04-16.

Tables

Table 1: Definitions of variables

Variable	Description
$p_{pl}^{diff} \equiv p_{N_j} - p_T$	difference between the logarithm of value-added deflator index in each non-tradable subsector and the tradable sector
$l_{pl}^{diff} \equiv l_T - l_{N_j}$	difference between the logarithm of productivity index in the tradable sector and each non-tradable subector
$w_{pl}^{diff} \equiv w_{N_j} - w_T$	difference between the logarithm of average wage index in each non-tradable subsector and the tradable sector
$p_{pl\_ea}^{diff} \equiv (p_{N_j} - p_T) - (p_{N_j}^{ea} - p_T^{ea})$	difference between differential price levels (non-tradables vs. tradables) in Poland and the euro area
$l_{pl\_ea}^{diff} \equiv (l_T - l_{N_j}) - (l_T^{ea} - l_{N_j}^{ea})$	difference between differential productivity levels (tradables vs. non-tradables) in Poland and the euro area
$w_{pl\_ea}^{diff} \equiv (w_{N_j} - w_T) - (w_{N_j}^{ea} - w_T^{ea})$	difference between differential wage index (non-tradables vs. tradables) in Poland and the euro area
$Q_{GDP} \equiv E\frac{P_{GDP}^{member\_state_i}}{P_{GDP}}$	real exchange rate (Poland vs. each euro area member state) deflated by the GDP deflator
$Q_{manufacturing} \equiv E\frac{P_{manufacturing}^{member\_state_i}}{P_{manufacturing}}$	real exchange rate (Poland vs. each euro area member state) deflated by the value added in manufacturing deflator
$l_{pl\_members}^{diff} \equiv (1-\delta)[(l_T - l_{\sum_j N_j}) - (l_T^{member\_state_i} - l_{\sum_j N_j}^{member\_state_i})]$	difference between differential productivity levels (tradables vs. aggregated non-tradables) in Poland and each euro area member state multiplied by the share of non-tradable sector in the economy (homogenity assumed)
$l_{pl\_members}^{*diff} \equiv (1-\delta)(l_T - l_{\sum_j N_j}) -$ $(1-\delta^{member\_state_i})(l_T^{member\_state_i} - l_{\sum_j N_j}^{member\_state_i})$	difference between differential productivity levels (tradables vs. aggregated non-tradables) in Poland and each euro area member state weighted by the share of non-tradable sector in each economy
$w_{pl\_members}^{diff} \equiv$ $(1-\delta)[(w_{\sum_j N_j} - w_T) - (w_{\sum_j N_j}^{member\_state_i} - w_T^{member\_state_i})]$	difference between differential wage levels (aggregated non-tradables vs. tradables) in Poland and each euro area member state multiplied by the share of non-tradable sector in the economy (homogeneity assumed)
$w_{pl\_members}^{*diff} \equiv (1-\delta)(w_{\sum_j N_j} - w_T) -$ $(1-\delta^{member\_state_i})(w_{\sum_j N_j}^{member\_state_i} - w_T^{member\_state_i})$	difference between differential wage levels (aggregated non-tradables vs. tradables) in Poland and each euro area member state weighted by the share of non-tradable sector in each economy

Table 2: The 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) Electricity, gas and water supply (E) Construction (F) Wholesale and retail trade; repair of motor vehicles, motorcycles, personal and household goods (G) Hotels and restaurants (H) Transport, storage and communication (I) Financial intermediation (J) Real estate, renting and business activities (K) Public administration and defence; compulsory social security (L) Education (M) Health and social work (N) Other community, social, personal service activities (O)

Table 3: The results of panel unit root tests – assessment of BS model assumptions validity

	p-value		
Variable	**Im-Pesaran-Shin**	**Fisher ADF**	**Fisher PP**
$Q_{manufacturing}$	0.00	0.00	0.85
$d(Q_{manufacturing})$	0.00	0.00	0.00
w_{pl}^{diff}	0.10	0.08	0.30
$d(w_{pl}^{diff})$	0.00	0.00	0.00
$w_{pl\_ea}^{diff}$	0.99	0.81	0.56
$d(w_{pl\_ea}^{diff})$	0.00	0.00	0.00
$w_{pl\_members}^{diff}$	0.69	0.10	0.01
$d(w_{pl\_members}^{diff})$	0.00	0.00	0.00
$w_{pl\_members}^{diff}$	0.30	0.24	0.00
$d(w_{pl\_members}^{diff})$	0.00	0.00	0.00

Table 4: The results of panel unit root tests (equation variables)

	p-value		
Variable	**Im-Pesaran-Shin**	**Fisher ADF**	**Fisher PP**
p_{pl}^{diff}	0.60	0.16	0.21
$d(p_{pl}^{diff})$	0.00	0.00	0.00
l_{pl}^{diff}	0.97	0.98	0.89
$d(l_{pl}^{diff})$	0.00	0.00	0.00
$p_{pl\_ea}^{diff}$	0.13	0.11	0.07
$d(p_{pl\_ea}^{diff})$	0.00	0.00	0.00
$l_{pl\_ea}^{diff}$	0.76	0.67	0.81
$d(l_{pl\_ea}^{diff})$	0.00	0.00	0.00
Q_{GDP}	0.00	0.00	0.56
$d(Q_{GDP})$	0.00	0.00	0.00
$l_{pl\_members}^{diff}$	0.81	0.76	0.32
$d(l_{pl\_members}^{diff})$	0.00	0.00	0.00
$l_{pl\_members}^{*diff}$	0.33	0.19	0.98
$d(l_{pl\_members}^{*diff})$	0.05	0.00	0.00

Table 5: Panel cointegration tests (1)

Variables		p_{pl}^{diff}, l_{pl}^{diff}	$p_{pl\_ea}^{diff}$, $l_{pl\_ea}^{diff}$	p_{pl}^{diff}, l_{pl}^{diff}, w_{pl}^{diff}	$p_{pl\_ea}^{diff}$, $l_{pl\_ea}^{diff}$, $w_{pl\_ea}^{diff}$
		Pedroni Panel Cointegration Test			
ADF group statistics		-5.54 (0.00)	-4.53 (0.00)	-5.65 (0.00)	-1.09 (0.14)
number of cointegrating vectors		**Johansen-Fisher Cointegration Test**			
trace statistics	**none**	55.39 (0.00)	29.47 (0.00)	129.3 (0.00)	121.2 (0.00)
	at most 1	30.21 (0.18)	10.55 (0.45)	53.20 (0.00)	52.40 (0.00)
	at most 2	-	-	30.62 (0.10)	25.58 (0.38)
maximum eigenvalue	**none**	52.79 (0.00)	20.91 (0.00)	102.0 (0.00)	93.99 (0.00)
	at most 1	30.21 (0.18)	-5.8 (0.45)	44.98 (0.00)	53.08 (0.00)
	at most 2	-	-	30.62 (0.10)	25.58 (0.38)

Table 6: Panel cointegration tests (2)

Variables		Q_{GDP}, $Q_{manufacturing}$, $l^{diff}_{pl\_members}$	Q_{GDP}, $Q_{manufacturing}$, $l^{diff}_{pl\_members}$, $w^{diff}_{pl\_members}$	Q_{GDP}, $Q_{manufacturing}$, $l^{*diff}_{pl\_members}$	Q_{GDP}, $Q_{manufacturing}$, $l^{*diff}_{pl\_members}$, $w^{*diff}_{pl\_members}$
		Pedroni Panel Cointegration Test			
ADF group statistics		1.39 (0.92)	−0.11 (0.61)	0.25 (0.60)	−0.11 (0.61)
number of cointegrating vectors		**Johansen-Fisher Cointegration Test**			
trace statistics	**none**	142.0 (0.00)	38.23 (0.00)	112.50 (0.00)	38.23 (0.00)
	at most 1	56.24 (0.00)	32.62 (0.00)	32.96 (0.00)	32.62 (0.00)
	at most 2	14.14 (0.59)	8.20 (0.22)	9.21 (0.81)	8.20 (0.22)
	at most 3	-	2.55 (0.86)	-	2.55 (0.86)
maximum eigenvalue	**none**	112.5 (0.00)	38.23 (0.00)	101.20 (0.00)	38.23 (0.00)
	at most 1	57.41 (0.00)	33.03 (0.00)	34.38 (0.00)	33.03 (0.00)
	at most 2	14.14 (0.59)	9.51 (0.15)	9.26 (0.81)	9.51 (0.15)
	at most 3	-	2.55 (0.86)	-	2.55 (0.86)

Table 7: The estimation results of the equations in growth rates (short-run estimates)

Effect	**Specification**	$\hat{\beta}$	$\hat{\gamma}$
Baumol-Bowen (wage homogeneity assumption)	$d(p_{pl}^{diff})_{it} = \alpha + \beta d(l_{pl}^{diff})_{it} + u_i + \varepsilon_{it}$	0.17 (0.02)	-
Baumol-Bowen (without wage homogeneity assumption)	$d(p_{pl}^{diff})_{it} = \alpha + \beta d(l_{pl}^{diff})_{it} + \gamma d(w_{pl}^{diff})_{it} + u_i + \varepsilon_{it}$	0.14 (0.03)	0.57 (0.01)
Balassa-Samuelson [inflation] (wage homogeneity assumption)	$d(p_{pl\_ea}^{diff})_{it} = \alpha + \beta d(l_{pl\_ea}^{diff})_{it} + u_i + \varepsilon_{it}$	0.21 (0.00)	-
Balassa-Samuelson [inflation] (without wage homogeneity assumption)	$d(p_{pl\_ea}^{diff})_{it} = \alpha + \beta d(l_{pl\_ea}^{diff})_{it} + \gamma d(w_{pl\_ea}^{diff})_{it} + u_i + \varepsilon_{it}$	0.18 (0.03)	0.19 (0.09)
Balassa-Samuelson [rer] (wage homogeneity assumption)	$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{diff})_{it} + \theta d(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.52 (0.00)	-
Balassa-Samuelson [rer] (wage homogeneity assumption, different share of tradables)	$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{*diff})_{it} + \theta d(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.53 (0.00)	-
Balassa-Samuelson [rer] (without wage homogeneity assumption)	$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{diff})_{it} + \gamma d(w_{pl\_members}^{diff})_{it} + \theta d(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.62 (0.00)	-0.01 (0.95)
Balassa-Samuelson [rer] (without wage homogeneity assumption, different share of tradables)	$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{*diff})_{it} + \gamma d(w_{pl\_members}^{*diff})_{it} + \theta d(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.60 (0.00)	-0.01 (0.94)

Table 8: The estimation results of the equations in log-levels (long-run estimates)

Specification	FMOLS		DOLS	
	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\beta}$	$\hat{\gamma}$
$(p_{pl}^{diff})_{it} = \alpha + \beta(l_{pl}^{diff})_{it} + u_i + \varepsilon_{it}$	0.65 (0.00)	-	0.60 (0.00)	-
$(p_{pl}^{diff})_{it} = \alpha + \beta(l_{pl}^{diff})_{it} + \gamma(w_{pl}^{diff})_{it} + u_i + \varepsilon_{it}$	0.60 (0.00)	1.38 (0.00)	0.60 (0.00)	1.08 (0.00)
$(p_{pl\_ea}^{diff})_{it} = \alpha + \beta(l_{pl\_ea}^{diff})_{it} + u_i + \varepsilon_{it}$	0.55 (0.00)	-	0.43 (0.00)	-
$(p_{pl\_ea}^{diff})_{it} = \alpha + \beta(l_{pl\_ea}^{diff})_{it} + \gamma(w_{pl\_ea}^{diff})_{it} + u_i + \varepsilon_{it}$	0.38 (0.00)	0.59 (0.00)	0.33 (0.00)	0.56 (0.00)
$(Q_{GDP})_{it} = \alpha + \beta(l_{pl\_members}^{diff})_{it} + \theta(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.76 (0.00)	-	-0.86 (0.00)	-
$(Q_{GDP})_{it} = \alpha + \beta(l_{pl\_members}^{*diff})_{it} + \theta(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.82 (0.00)	-	-0.86 (0.00)	-
$(Q_{GDP})_{it} = \alpha + \beta(l_{pl\_members}^{diff})_{it} + \gamma(w_{pl\_members}^{diff})_{it} + \theta(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.55 (0.00)	-0.44 (0.00)	-0.60 (0.00)	-0.29 (0.00)
$(Q_{GDP})_{it} = \alpha + \beta(l_{pl\_members}^{*diff})_{it} + \gamma(w_{pl\_members}^{*diff})_{it} + \theta(Q_{manufacturing})_{it} + u_i + \varepsilon_{it}$	-0.62 (0.00)	-0.42 (0.00)	-0.63 (0.00)	-0.28 (0.00)

Table 9: The estimation results of the Error Correction Models

Specification	FMOLS-estimated ECT			DOLS-estimated ECT		
	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
$d(p_{pl}^{diff})_{it} = \alpha + \beta d(l_{pl}^{diff})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	0.20 0.01	-	-0.28 0.00	0.19 0.01	-	-0.29 0.00
$d(p_{pl}^{diff})_{it} = \alpha + \beta d(l_{pl}^{diff})_{it} + \gamma d(w_{pl}^{diff})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	0.20 0.01	1.00 0.00	-0.24 0.00	0.21 0.01	0.96 0.00	-0.26 0.00
$d(p_{pl\_ea}^{diff})_{it} = \alpha + \beta d(p_{pl\_ea}^{diff})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	0.22 0.00	-	-0.23 0.00	0.20 0.00	-	-0.22 0.00
$d(p_{pl\_ea}^{diff})_{it} = \alpha + \beta d(l_{pl\_ea}^{diff})_{it} + \gamma d(w_{pl\_ea}^{diff})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	0.24 0.00	0.31 0.00	-0.26 0.00	0.23 0.00	0.31 0.00	-0.26 0.00
$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{diff})_{it} + \theta d(Q_{manufacturing})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	-0.28 0.01	-	-0.47 0.01	-0.29 0.01	-	-0.41 0.00
$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{*\,diff})_{it} + \theta d(Q_{manufacturing})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	-0.38 0.00	-	-0.51 0.00	-0.38 0.00	-	-0.44 0.00
$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{diff})_{it} + \theta d(Q_{manufacturing})_{it} + \gamma d(w_{pl\_members}^{diff})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	-0.39 0.00	-0.17 0.02	-0.50 0.00	-0.37 0.00	-0.09 0.21	-0.47 0.00
$d(Q_{GDP})_{it} = \alpha + \beta d(l_{pl\_members}^{*\,diff})_{it} + \theta d(Q_{manufacturing})_{it} + \gamma d(w_{pl\_members}^{*\,diff})_{it} + \delta ECT_{it-1} + u_i + \varepsilon_{it}$	-0.43 0.00	-0.21 0.01	-0.53 0.00	-0.41 0.00	-0.12 0.11	-0.50 0.00

Table 10: The estimation results of GDP-augmented equations

Effect	SR			ECM [FMOLS-estimated ECT]		
	$\hat{\beta}^{GDP-augumented}$	$\hat{\beta}$	$\hat{\theta}^{*}$	$\hat{\beta}^{GDP-augumented}$	$\hat{\beta}$	$\hat{\theta}^{*}$
Baumol-Bowen (wage homogeneity assumption)	0.12 (0.10)	0.17 (0.02)	1.07 (0.00)	0.18 (0.01)	0.20 0.01	0.84 (0.00)
Baumol-Bowen (without wage homogeneity assumption)	0.12 (0.13)	0.14 (0.03)	1.21 (0.03)	0.18 (0.01)	0.20 0.01	1.14 (0.03)
Balassa-Samuelson [inflation] (wage homogeneity assumption)	0.22 (0.01)	0.21 (0.00)	0.33 (0.24)	0.23 (0.00)	0.22 0.00	−0.11 (0.68)
Balassa-Samuelson [inflation] (without wage homogeneity assumption)	0.25 (0.00)	0.18 (0.03)	0.55 (0.07)	0.24 (0.00)	0.24 0.00	0.09 (0.76)
Balassa-Samuelson [rer] (wage homogeneity assumption)	−0.33 (0.01)	−0.52 (0.00)	−1.05 (0.00)	−0.26 (0.01)	−0.28 0.01	−0.27 (0.10)
Balassa-Samuelson [rer] (wage homogeneity assumption, different share of tradables)	−0.32 (0.02)	−0.53 (0.00)	−1.04 (0.00)	−0.39 (0.00)	−0.38 0.00	−0.59 (0.01)
Balassa-Samuelson [rer] (without wage homogeneity assumption)	−0.45 (0.00)	−0.62 (0.00)	−1.36 (0.00)	−0.35 (0.00)	−0.39 0.00	−0.42 (0.05)
Balassa-Samuelson [rer] (without wage homogeneity assumption, different share of tradables)	−0.48 (0.00)	−0.60 (0.00)	−1.60 (0.00)	−0.35 (0.00)	−0.43 0.00	−0.18 (0.27)

*The GDP per capita parameter estimates

Table 11: Estimates of Baumol-Bowen and Balassa-Samuelson effects (1995–2008)

	Estimates			Average of the independent variable in the sample period	Share of the non-tradable sector	Effects estimates – contribution of productivity differential to inflation[b] or RER appreciation[a]		
Model specification	LR^{*}	SR^{ECM*}	SR	$d(l^{\bar{diff}})$		LR^{*}	SR^{ECM*}	SR
Baumol-Bowen (wage homogeneity assumption)	0.63	0.20	0.17	5.9	0.8	3.0	0.9	0.8
Baumol-Bowen (without wage homogeneity assumption)	0.60	0.21	0.14	5.9	0.8	2.8	1.0	0.7
Balassa-Samuelson [inflation] (wage homogeneity assumption)	0.49	0.21	0.21	4.3	-	2.1	0.9	0.9
Balassa-Samuelson [inflation] (without wage homogeneity assumption)	0.36	0.24	0.18	4.3	-	1.5	1.0	0.8
Balassa-Samuelson [RER] (wage homogeneity assumption)	-0.81	-0.29	-0.52	3.5	-	-2.8	-1.0	-1.8
Balassa-Samuelson [RER] (wage homogeneity assumption, different share of tradables)	-0.84	-0.38	-0.53	2.8	-	-2.4	-1.1	-1.5
Balassa-Samuelson [RER] (without wage homogeneity assumption)	-0.58	-0.38	-0.62	3.5	-	-2.0	-1.3	-2.2
Balassa-Samuelson [RER] (without wage homogeneity assumption, different share of tradables)	-0.63	-0.42	-0.60	2.8	-	-1.8	-1.2	-1.7

[a] In percentage points.

[b] By inflation we mean artificial value-added deflator, composed solely of the NACE sectors C through O.

* The average of the FMOLS and DOLS estimate.

Table 12: Estimates of Baumol-Bowen and Balassa-Samuelson effects (1999–2008)

	Estimates			**Average of the independent variable in the sample period**	**Share of the non-tradable sector**	**Effects estimates – contribution of productivity differential to inflation[b] or RER appreciation[a]**		
Model specification	LR^*	SR^{ECM*}	SR	$d(l\bar{diff})$		LR^*	SR^{ECM*}	SR
Baumol-Bowen (wage homogeneity assumption)	0.63	0.20	0.17	5.4	0.8	2.7	0.9	0.7
Baumol-Bowen (without wage homogeneity assumption)	0.60	0.21	0.14	5.4	0.8	2.6	0.9	0.6
Balassa-Samuelson [inflation] (wage homogeneity assumption)	0.49	0.21	0.21	3.7	-	1.8	0.8	0.8
Balassa-Samuelson [inflation] (without wage homogeneity assumption)	0.36	0.24	0.18	3.7	-	1.3	0.9	0.7
Balassa-Samuelson [RER] (wage homogeneity assumption)	-0.81	-0.29	-0.52	2.8	-	-2.3	-0.8	-1.5
Balassa-Samuelson [RER] (wage homogeneity assumption, different share of tradables)	-0.84	-0.38	-0.53	2.2	-	-1.8	-0.8	-1.2
Balassa-Samuelson [RER] (without wage homogeneity assumption)	-0.58	-0.38	-0.62	2.8	-	-1.6	-1.1	-1.7
Balassa-Samuelson [RER] (without wage homogeneity assumption, different share of tradables)	-0.63	-0.42	-0.60	2.2	-	-1.4	-0.9	-1.3

[a] In percentage points.

[b] By inflation we mean artificial value-added deflator, composed solely of the NACE sectors C through O.

* The average of the FMOLS and DOLS estimate.

Table 13: Contributions of the BB and BS effects to the inflation and real exchange rate development

	Average		**Baumol-Bowen / Balassa-Samuelson effect estimates**			
			SR		**LR**	
Variable	**1995–2008**	**1999–2008**	**1995–2008**	**1999–2008**	**1995–2008**	**1999–2008**
$d(p_{pl})^*$	6.0	2.0	0.7 – 1.0	0.6 – 0.9	2.8 – 3.0	2.6 – 2.7
$d(p_{pl\_ea}^{diff})$	4.9	3.8	0.8 – 1.0	0.7 – 0.9	1.5 – 2.1	1.3 – 1.8
$d(Q_{GDP})$	-1.8	-1.1	-1.0 – -2.2	-0.8 – -1.7	-1.8 – -2.8	-1.4 – -2.3

* The dynamics of an artificial value-added deflator, composed solely of the NACE sectors C through O.

Appendix: Literature overview

Author and year	**Title**	**Method**	**Sample**	**Tradable sector (T)**	**Nontradable sector (NT)**	**Results**
García-Solanes et al. (2008)	Beyond the Balassa-Samuelson effect in some new member states of the European Union	panel cointegration	1995-2004	UNO classification (without agriculture)	construction, market services	1% productivity growth in Poland against Germany causes ca. 1,3% growth of difference in relative prices
Wagner (2005)	The Balassa-Samuelson Effect in 'East & West'. Differences and Similarities.	panel cointegration (bootstrapping)	1994-2001	NACE C, D, E	NACE F- K	BS effect: 0.80 p.p. per year
Wagner and Hlouskova (2004)	What's Really the Story with this Balassa-Samuelson Effect in the CEEC?	panel cointegration (bootstrapping)	1993-2001	NACE C, D, E	NACE F- K	BB effect: 1.219 p.p. per year BS effect for inflation: 0.893 pp. per year
Breuss (2003)	Balassa-Samuelson Effects in the CEEC: Are they Obstacles for Joining the EMU?	CGE model (2 sectors, 5 factors, 17 regions)	1997	mainly manufacturing, some financial services and agriculture	mainly services	1 per cent productivity growth in the tradable sector causes 2.4% growth of relative prices (T against NT) and 0.89% real exchange rate appreciation
Chmielewski (3003)	Od kursu płynnego do unii monetarnej. Znaczenie efektu Balassy-Samuelsona dla polskiej polityki pieniężnej	Johansen method	1995-2002	manufacturing (industry)	market services (market services and construction)	BB effect: 1.12-2.26 (over 1999-2002) and 1.46-2.94 (in the entire sample) BS effect: 1.36-1.62 and 1.82-2.17 respectively

Lojschová (2003)	Estimating the Impact of the Balassa-Samuelson Effect in Transition Economies	panel econometrics	1995-2002	manufacturing	services, construction	BB effect: 1.3 p.p. per year BS effect: 2 p.p. per year
Mihaljek and Klau (2003)	The Balassa-Samuelson effect in central Europe: a disaggregated analysis	OLS	1995-2001	manufacturing, mining, transport, communication, tourism	energy, construction, retail and wholesale trade, real estate renting and business activities, financial services, education, health, peronal services	BB effect: 1.196 p.p .per year BS effect: 0.118 p.p. per year
M. Dubravko	The Balassa-Samuelson effect in Central Europe: a disaggregated analysis	OLS	1994-2001	manufacturing, mining and quarrying, transport, tourism	energy, water and gas supply, construction, wholesale and retail trade, financial intermediation, real estate, renting and business activities, repair, education, health and other personal services	BB effect: 1.41 p.p. per year BS effect (CPI): 0.118 p.p. per year

Égert (2002a)	Estimating the impact of the Balassa-Samuelson effect on inflation and the real exchange rate during the transition	Johansen method	1999-2000	industry (manufacturing)	services	BS effect (CPI): 0,882-1,505 p.p. per year
Égert (2002b)	Investigating the Balassa-Samuelson hypothesis in transition: Do we understand what we see?	panel cointegration	1991-2001	industry	services	BS effect: 1,5-2,2%
Łukasz Rawdanowicz (2002)	Badanie wpływu efektu Harroda-Balassy-Samuelsona na ceny relatywne w Polsce	Johansen method	1996-2001	agriculture and industry	construction, services	BB effect: 4,3 p.p.
Halpern and Wyplosz (2001)	Economic Transformation and Real Exchange Rates in the 2000s: The Balassa-Samuelson Connection	panel econometrics	1991-1999	industry	services	productivity growth in T causes growth of relative NT prices by 2.4% in the short run and 4.4% in the long run
Cipriani (2001)	The Balassa-Samuelson Effect in Transition Economies	OLS	1995-1999	industry, construction	market services	BB effect: 1,5 p.p. per year