Central European Journal of Economic Modelling and Econometrics

Estimating the Baumol-Bowen and Balassa-Samuelson Effects in the Polish Economy – a Disaggregated Approach

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Submitted: 9.08.2010, Accepted: 30.01.2011

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 $0.9 - 1.0$ percentage points in the short run. According to estimation results, the Balassa-Samuelson effect contributed around 0.9 to 1.0 percentage point per annum to the rate of relative price growth between Poland and the euro area and 1.0 to 1.2 p.p. to real exchange rate appreciation. The long-run effects are of an approximately twice larger magnitude. 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.

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

JEL Classification: C33, E31, F31, F41.

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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 Baumol and Bowen (1966), explaining the cross-sectoral inflation differential, and encompassing Balassa-Samuelson effect, additionally accounting for the real exchange rate appreciation.

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; former Articles 121-123 of the Treaty establishing the European Community). 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). 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. 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. After the euro adoption the real appreciation will not any more be channeled through the

nominal exchange rate adjustment, which 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, provided that consuming tradable goods requires some nontradable input (e.g. transport and distribution), higher price dynamics from of the nontradables could spill over into the tradable sector.

For both reasons, a quantitative, up-to-date assessment of the BS effect contribution to domestic inflation is crucial. Over the recent decade, both issues have been investigated in a wide range of empirical studies of catching-up countries, especially the NMS. 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 article attempts to address this need.

The article aims to contribute to the existing empirical literature by applying a disaggregated approach to estimating the Baumol-Bowen and Balassa-Samuelson effects in the Polish economy. The novelty of this approach consists in employing a multi-sector decomposition of the economy, that is the price, productivity and wage differentials are computed as a difference between the aggregated tradable sector and various branches of the nontradable sector. This approach has an advantage over the standard methods applied in the literature – both time-series approach and panel analysis for a group of countries – as it overcomes the problem of relatively short time span and low frequency of series available for NMS by applying panel econometric techniques, but at the same time enables to estimate both short-run and long-run effects for a particular economy.

On the basis of the obtained estimates we quantify the average magnitude of Baumol-Bowen and Balassa-Samuelson effects over the whole sample (years 1999 through 2008) and in the shorter sub-sample (1999-2008). We argue that, owing to the real convergence process towards the euro area, both effects add significantly to the Polish inflation rate and to the real appreciation of the Polish zloty. The magnitude of the effects should, however, fade away in time and, therefore, the estimates for the second sub-sample should be lower than for the whole period.

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 and literature overview

We derive the model of BB and BS effects, following the standard approach in the literature. 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:

1. A small open economy consists of two sectors – the tradable (T) and the non-tradable (NT) one.

2. 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 analysed small economy.

3. The capital is perfectly mobile between sectors and regions.

4. 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.

5. There is perfect competition in both sectors (in both regions).

6. 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:

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

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

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

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

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

where w_j ($j\epsilon\{T,N\}$) stands for the wage in respective sector and r is the cost of capital.
First order conditions for the above maximization problems (with respect to labour) are the following:

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

This implies the prices of labour as follows:

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

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

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

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

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

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

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 $j \in \{T,N\}$. 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 (the long-run and the short-run effect, respectively).

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

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

where E denotes nominal exchange rate.
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 tradable sectors at home and abroad are symmetrically sized ($\delta = \delta^*$), formula (17) collapses to:

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

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

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

2.4 Alternative explanations for relative price differentials

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 may potentially be more important in this respect. The higher pace of non-tradables' price growth may, namely, result from higher income elasticity of this sector's products, especially services. The rationale behind this hypothesis is that in a catching-up economy, consumers shift their demand from tradable to non-tradable sector as they become richer. This explanation is in line with the Linder (1961) hypothesis, according to which GDP *per capita* is the most important determinant of the demand structure.

Another factor which may potentially account for the relative price differentials between sectors and countries is the relative endowment in production inputs. Under the assumption of higher capital-intensity of production in the tradable sector, it is argued that prices of non-tradables will be higher in countries relatively more endowed in capital. This effect is referred to as the Heckscher-Ohlin hypothesis; see Jones (1956).

The encompassing theoretical model, explaining relative price differentials by both supply-side (the Balassa-Samuelson hypothesis) and demand-side (the Linder hypothesis) effects, as well as relative factor endowment (the Heckscher-Ohlin hypothesis), was derived by Bergstrand (1991).

2.5 Overview of the empirical literature

The empirical investigation on the existence of the Baumol-Bowen and Balassa-Samuelson effects consists in regressing price level or growth rate differentials (cross-sector and cross-country, respectively) on productivity differentials. The hypothesis is verified on the basis of significance of the coefficient corresponding to the differential productivity variable, and the magnitude of the effects is calculated as the product of the estimated coefficient and the average value of productivity variable over the sample period.

The BB and BS effects in Poland and other Central European countries have been investigated in a number of empirical studies. Equations similar to those derived in subsections 2.1-2.3 were fit with various econometric approaches, including both time series and panel econometric techniques. Cipriani (2001) and Mihaljek and Klau (2003) run OLS regressions. Chmielewski (2003), Rawdanowicz (2002) and Égert (2002a) use the Johansen method for cointegrated time series. In a number of studies by García-Solanes, Sancho-Porteroa, Torrejón-Floresa (2008), Wagner and Hlouskova (2004), Wagner (2005), Lojschová (2003), Égert (2002b) and Halpern and Wyplosz (2001), a panel of Central European countries is examined. Finally, Breuss (2003) based his calculations on a computable general equilibrium model.

The magnitude of Baumol-Bowen effect in the case of Poland, measured as annual contribution to domestic price dynamics, differs across these studies from 1.5 to 2.9 - Chmielewski (2003) - or even 4.3 percentage points; Rawdanowicz (2002). The estimated contributions of Balassa-Samuelson effect to annual inflation rates range from 0.118 - Mihaljek and Klau (2003) - to 2.2; Égert (2002b).

Most of this literature is based on samples starting in mid-1990s and ending in mid-

2000s. However, more recent analyses – to the best of our knowledge – are missing, at least in the case of Poland.

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 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 (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 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. Due to a large number of missing values in the sectoral compensation of employees series (and virtually no data for the euro area), we decided to replace the sectoral wages with the data on a higher level of aggregation. That is instead of separate wage series for e.g. section G, H and I of NACE 1.1. classification (see Table

2) we imputed an average for these three, that is the wage series for non-financial market services. Likewise we substituted wage series for sections C and E with the average for non-manufacturing industry, for sections J and K with financial market services and for sections L, M, N, O with public services. In the case of substantial heterogeneity the averages on the higher aggregation level may not reflect the true sectoral wage trajectories. Therefore we treat the estimation results of the wage-augmented equations merely as a robustness check and interpret them with caution.
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, Austria, Portugal and partially Greece because of incomplete data). The exclusion from the analysis of the catching-up countries (Ireland, Portugal and partially Greece) may bias upward the magnitude of the Balassa-Samuelson effect with respect to the euro area as a whole.
Table 1 contains the definitions of the variables used in the empirical analysis.

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) and random events, such as weather conditions.

3.3 Methodological notes

3.3.1 Panel unit root tests

The inference on the stationarity of the analysed series is based both on the first generation unit root tests, assuming cross-sectional independence, as well as second generation test, allowing for cross-correlation.
According to the null hypothesis of all the applied tests, all cross-sectional processes contain a unit root:

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

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

Table 1: Definitions of variables

$p_{pl}^{diff} \equiv p_{N_j} - p_T$ – 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
$p_{pl_ea}^{diff} \equiv \left(p_{N_j} - p_T\right) - \left(p_{N_j}^{ea} - p_T^{ea}\right)$ – difference between differential price levels (non-tradables vs. tradables) in Poland and the euro area
$l_{pl_ea}^{diff} \equiv \left(l_T - l_{N_j}\right) - \left(l_T^{ea} - l_{N_j}^{ea}\right)$ – difference between differential productivity levels (tradables vs. non-tradables) in Poland and the euro area
$w_{pl_ea}^{diff} \equiv \left(w_{N_j} - w_T\right) - \left(w_{N_j}^{ea} - w_T^{ea}\right)$ – 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
$l_{pl_members}^{diff} \equiv (1-\delta)\left[\left(l_T - l_{\sum_j N_j}\right) - \left(l_T^{member_state_i} - l_{\sum_j N_j}^{member_state_i}\right)\right]$ – 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)
$l_{pl_members}^{*diff} \equiv (1-\delta)\left(l_T - l_{\sum_j N_j}\right) - \left(1-\delta^{member_state_i}\right)\left(l_T^{member_state_i} - l_{\sum_j N_j}^{member_state_i}\right)$ – 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
$w_{pl_members}^{diff} \equiv (1-\delta)\left[\left(w_{\sum_j N_j} - w_T\right) - \left(w_{\sum_j N_j}^{member_state_i} - w_T^{member_state_i}\right)\right]$ – 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)
$w_{pl_members}^{*diff} \equiv (1-\delta)\left(w_{\sum_j N_j} - w_T\right) - \left(1-\delta^{member_state_i}\right)\left(w_{\sum_j N_j}^{member_state_i} - w_T^{member_state_i}\right)$ – 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
	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)
Manufacturing (D)	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)

stationary.

From the first generation panel unit root tests we choose those 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, Shin (2003) test (IPS) and Fisher-type Dickey-Fuller test; Maddala and Wu (1999). The individual unit root processes assumption reduces the unobserved heterogeneity problem and, according to Monte Carlo simulations; Im, Pesaran, Shin (2003), results in higher power of the tests compared to those based on the supposition of common persistence parameter across cross-sectional units.

The Im, Pesaran, Shin (2003) statistics is obtained by a two-step procedure. In the first step a separate ADF regression is estimated for each cross-sectional unit:

$$\Delta y_{i,t} = \alpha_{0i} + (\alpha_i - 1)y_{i,t-1} + \sum_{k=1}^{K_i} \beta_{ik}\Delta y_{i,t-k} + \varepsilon_{i,t}. \tag{26}$$

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

$$IPS = N^{-1}\sum_{i=1}^{N} t_{(\alpha_i - 1)} \tag{27}$$

The standardized IPS statistics has an asymptotic (with $N \to \infty$) standard normal distribution. The values of the mean and the variance have been computed by means of Monte Carlo methods and presented in Im, Pesaran, Shin (2003).
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:

$$-2\sum_{i=1}^{N}\log(p_i), \tag{28}$$

where p_i denotes the p-value from the individual ADF test. In the limit (with $T \to \infty$), the test has a χ^2_{2N} distribution.
The Im, Pesaran and Shin as well as Maddala and Wu tests assume cross-sectional independence in the data. If this assumption is violated, however, both tests suffer from serious size distortions, which results in over-rejection of non-stationarity null; Banerjee, Marcellino, Osbat (2005). Owing to strong inter-economy linkages the risk of cross-correlation in our data is non-negligible. For this reason we additionally test for the unit root by applying the cross-sectionally augmented IPS (CIPS) test as proposed by Pesaran (2003). It is assumed that the cross-correlation can be ascribed to a single unobserved common factor (f_t):

$$\begin{gathered}\Delta y_{i,t} = \alpha_{0i} + (\alpha_i - 1)y_{i,t-1} + \textstyle\sum_{k=1}^{K_i}\beta_{ik}\Delta y_{i,t-k} + u_{i,t} \\ u_{i,t} = \lambda_i f_t + \varepsilon_{i,t}.\end{gathered} \tag{29}$$

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:

$$\Delta y_{i,t} = \alpha_{0i} + (\alpha_i - 1)y_{i,t-1} + \gamma_i\bar{y}_{t-1} + \delta_i\Delta\bar{y}_t + \sum_{k=1}^{K_i}\beta_{ik}\Delta y_{i,t-k} + \varepsilon_{i,t}, \tag{30}$$

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:

$$CIPS = N^{-1}\sum_{i=1}^{N} t_{(\alpha_i - 1)}. \tag{31}$$

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

3.3.2 Panel cointegration tests

We test for the presence of cointegration by means of three panel cointegration tests – two Engle-Granger based tests, the Pedroni (2004) test and the Westerlund (2007)

test, and the Johansen-type Fisher test; Maddala and Wu (1999).
The Pedroni test is residual-based, that is 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

$$y_{i,t} = \alpha_{0i} + \beta_i x_{i,t} + \varepsilon_{i,t} \tag{32}$$

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. The ADF statistics performs also 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:

$$\Delta\hat{\varepsilon}_{i,t} = \alpha_{0i} + (\alpha_i - 1)\hat{\varepsilon}_{i,t-1} + \sum_{k=1}^{K_i} \beta_{ik}\Delta\hat{\varepsilon}_{i,t-k} + \vartheta_{i,t}, \tag{33}$$

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

$$\tilde{Z}_{ADF} = \sum_{i=1}^{N} \left(\hat{s}_i^2 \sum_{t=1}^{T} \hat{\varepsilon}_{i,t-1}^2 \right)^{-\frac{1}{2}} \sum_{t=1}^{T} \left(\hat{\varepsilon}_{i,t-1}\Delta\hat{\varepsilon}_{i,t} \right), \tag{34}$$

where $\hat{s}_i^2 = \frac{1}{T}\sum_{t=1}^{T} \hat{\vartheta}_{i,t}^2$. The standardised test statistics has an asymptotic standard normal distribution. The mean and variance adjustment terms for different number of regressors and deterministic components are tabulated in Pedroni (1999).
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:

$$\Delta y_{i,t} = \alpha_{0i} + \alpha_i \left(y_{i,t-1} - \beta_i x_{i,t-1} \right) + \sum_{j=1}^{p_i} \alpha_{ij}\Delta y_{i,t-j} + \sum_{j=1}^{p_i} \gamma_{ij}\Delta x_{i,t-j} + \epsilon_{i,t}. \tag{35}$$

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

$$\Delta y_{i,t} = \alpha_{0i} + \alpha_i y_{i,t-1} + \lambda_i x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij}\Delta y_{i,t-j} + \sum_{j=1}^{p_i} \gamma_{ij}\Delta x_{i,t-j} + \epsilon_{i,t}, \tag{36}$$

where $\lambda_i = -\alpha_i\beta_i$. Four tests based on the estimates of α_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). Additionally, Westerlund proposed a bootstrap-based critical values robust to cross-sectional dependence.
According to Monte Carlo experiments - Westerlund (2007) - two of those statistics – mean group τ (G_τ) and panel τ (P_τ) – seem to outperform the other two (mean-group α and panel α) 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 (36) is estimated separately for each i and the G_τ statistics is specified as follows:

$$G_\tau = N^{-1}\sum_{i=1}^{N} t_{\hat{\alpha}_i}. \tag{37}$$

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:

$$\Delta y_{i,t} = \alpha_{0i} + \lambda_i x_{i,t-1} + \sum_{j=1}^{p_i}\alpha_{ij}\Delta y_{i,t-j} + \sum_{j=1}^{p_i}\gamma_{ij}\Delta x_{i,t-j} + \epsilon_{i,t} \tag{38}$$

and

$$y_{i,t} = \alpha_{0i} + \lambda_i x_{i,t-1} + \sum_{j=1}^{p_i}\alpha_{ij}\Delta y_{i,t-j} + \sum_{j=1}^{p_i}\gamma_{ij}\Delta x_{i,t-j} + \epsilon_{i,t}. \tag{39}$$

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, α, is estimated. The P_τ 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:

$$-2\sum_{i=1}^{N}\log(p_i) \to \chi^2_{2,N}. \tag{40}$$

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:

$$\begin{cases} y_{it} = \alpha_i + \beta x_{it} + \mu_{it} \\ x_{it} = x_{it-1} + \varepsilon_{it} \end{cases} \qquad i = 1, ..., N \tag{41}$$

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.

$$\Omega_i = \sum_{k=-\infty}^{\infty} \Gamma_i^k = \Gamma_i^0 + \sum_{k=1}^{\infty} \left(\Gamma_i^k + \Gamma_i^{k\,T}\right), \tag{42}$$

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:

$$\hat{\Omega}_i = \hat{\Gamma^0}_i + \hat{\Gamma}_i + \hat{\Gamma}_i^T, \tag{43}$$

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

$$\hat{\Omega}_i = \hat{L}_i \hat{L}_i^T, \tag{44}$$

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:

$$y_{it}^* = y_{it} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} \Delta x_{it}, \tag{45}$$

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

$$\hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Gamma^0}_{21i} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} \left(\hat{\Gamma}_{22i} + \hat{\Gamma^0}_{21i}\right) \tag{46}$$

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

$$\hat{\beta}^{FMOLS} = \frac{1}{N} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} (x_{it} - \bar{x}_i)^2 \right)^{-1} \left(\sum_{t=1}^{T} (x_{it} - \bar{x}_i)\, y_{it}^* - T\hat{\gamma}_i \right) \tag{47}$$

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:

$$y_{it} = \alpha + \beta x_{it} + \sum_{p=-P}^{P} \delta_p \Delta x_{it-p} + u_i + \varepsilon_{it} \tag{48}$$

The estimator is given as:

$$\hat{\beta}^{DOLS} = \frac{1}{n} \sum_{i=1}^{n} \left(\sum_{t=1}^{T} z_{it} z_{it}^{'} \right)^{-1} \left(\sum_{t=1}^{T} z_{it} y_{it} \right), \tag{49}$$

where $z_{it} = (x_{it} - \overline{x_i}, \Delta x_{it-P}, \ldots, \Delta x_{it+P})$ 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 conditional upon the validity of the theoretical model assumptions. For this reason, at 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 a unit root 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 Pesaran CIPS test points out to the unit root in the data generating process. Owing to a high risk of cross-sectional dependence (nominal exchange rates on the basis of which the real exchange rates are computed have very similar or – after the euro adoption –

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

	p-value		test statistics
Variable	Im-Pesaran-Shin	Fisher ADF	Pesaran CIPS[1]
$Q_{manufacturing}$	0.00	0.00	-1.35
$d(Q_{manufacturing})$	0.00	0.00	-4.16***
w_{pl}^{diff}	0.10	0.08	-1.96
$d(w_{pl}^{diff})$	0.00	0.00	-2.91**
$w_{pl_ea}^{diff}$	0.99	0.81	-1.76
$d(w_{pl_ea}^{diff})$	0.00	0.00	-3.12***
$w_{pl_members}^{diff}$	0.30	0.24	-1.34
$d(w_{pl_members}^{diff})$	0.00	0.00	-2.54**

[1]The critical values for CIPS test are tabulated in Pesaran (2003). One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

identical trajectories across cross-sectional units) and consequently size distortions in the case of first generation unit root tests as well as numerous theoretical arguments (as 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) and empirical results 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 tradable price deflators when moving from (18) into (19) and have to 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), since the sectoral wage series are proxied by the averages on the higher level of aggregation and therefore wage-augmented equations serve merely as a robustness check.

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, that is we use $l^{*diff}_{pl_members}$ instead of $l^{diff}_{pl_members}$ as the explanatory variable. 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 and cointegration testing

All the "level" variables (difference in log-indices) seem to be non-stationary (see Table 11 in Appendix), 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 first generation tests reject the non-stationarity null. 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}.

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 level equations as long-run 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 and Westerlund tests, as well as Fisher-Johansen statistics clearly indicate the existence of a long-run equilibrium (see Table 12 in Appendix). On the other hand, the existence of wage-augmented long-run BS effect seems to be backed only by the Fisher-type statistics. In the case of trivariate systems (wage-augmented equations) 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.

A more nuanced picture emerges from the cointegration tests 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 13 in Appendix). 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.

Despite the absence of a clear conclusions whether (and how many) cointegrating relations exist in the case of some variables' sets, we proceed estimating both short-

run and long-run equations. Much caution is however given to the interpretation of the results obtained.

4.3 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 unreported results of models' estimation are available from authors upon request). The estimates of β parameter in all the empirical specifications are statistically significant, albeit substantially less than unity (see Table 4) – 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 somewhat 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 p.p. 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 the 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.

The comparison between the coefficients in the BS equation for dual inflation differential and real exchange rate 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 channeled via the nominal exchange 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 pairs, whereas in the inflation differential equations – as sectors. Due to the exclusion of some catching-up countries from the former panel (see Subsection 3.1) the Balassa-Samuelson effect with respect to the real exchange rate may be upward biased.

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 cannot exclude the existence of some lagged adjustment that could be captured in the long-term specification.

Table 4: The estimation results of the equations in growth rates (short-run estimates)

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

4.4 Long-run estimates

Having established the cointegration relationships between differential price levels and productivity, we can estimate the long-run relationships by means of the FMOLS and DOLS estimator. Table 5 presents the long-run versions of the estimates in Table 4.

Table 5: The estimation results of the equations in log-levels (long-run estimates)

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

All the variables are significant in all the equations and the parameters are signed in line with theoretical priors. 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 0.43% to 0.55% in response to a 1% growth of relative productivity when we consider the relative prices as a dependent variable, and from 0.76% to even 0.86% when we take into account the GDP-deflated real exchange rate. The estimates obtained from the wage-augmented equations are lower than their reduced-form counterparts, which may imply that the productivity-induced inflation pressure is mitigated in the long run by lower wage growth in the non-tradable sector.

The results seem to be relatively robust to the choice of the estimation method. 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 6), comprising the short-run formulation from Table 4 and an error correction term, i.e. the lagged residual of the corresponding cointegration regression in Table 5. In the presence of cointegration, the estimates presented in Table 4 are biased due to the omitted error correction term. Nevertheless, owing to possible distortions of cointegration tests applied to small-size (both N and T) panels (e. g. see Pedroni, 2004), we take into account the short-run estimates from both Table 4 and Table 6. Again, the short-run parameters in

Table 6: The estimation results of the Error Correction Models

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

question are correctly signed and significant in both the Baumol-Bowen and Balassa-

Samuelson equations. In all the specifications, the error correction parameter is significantly lower than 0 and ranges from -0.22 to -0.29 in the case of inflation equations and form -0.41 to -0.53 in the case of real exchange rate equations. 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.
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.

4.5 Robustness of the results

As presented in Bergstrand (1991) the Balassa-Samuelson effect is not the only possible explanation for relative price differentials (see Subsection 2.4). Another two factors which may contribute in this respect are relative factor endowments and demand-side developments. Therefore estimating BS equations augmented with relative capital-labour ratio and GDP *per capita* as a proxy for demand effects may serve for the purpose of robustness check. Unfortunately, owing to the lack of data on capital stock it is virtually impossible to compute sectoral capital-labour ratios for the Polish economy. For this reason we restrain the robustness check solely to short-run phenomena by augmenting the short-run equations with additional regressor, i.e. GDP *per capita* (Table 7).
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 slightly 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.

4.6 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

Table 7: The estimation results of GDP-augmented equations

Effect	SR			ECM [FMOLS-estimated ECT]		
	$\hat{\beta}^{GDP-aug.}$	$\hat{\beta}$	$\hat{\theta}^*$	$\hat{\beta}^{GDP-aug.}$	$\hat{\beta}$	$\hat{\theta}^*$
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

of the non-tradable sector in the economy. 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.

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; see Table 8 and 9). This is relatively small, compared to the average rate of inflation (by inflation we mean here an artificial value-added deflator, composed solely of the NACE sectors C through O) in this period, which amounted to 6.0% (2.0% in years 1999 through 2008). The long-run estimates of the Baumol-Bowen effect are of higher magnitude: 2.8-3.0 percentage points contribution to country-specific inflation (2.6-2.7 in years 1999 through 2008).

In line with expectations, the results obtained for the more recent sub-sample

Table 8: Estimates of Baumol-Bowen and Balassa-Samuelson effects (1995–2008)

Model specification	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)]		
	LR^*	SR^{ECM*}	SR	$d\left(\bar{l^{diff}}\right)$		LR^*	SR^{ECM*}	SR
BB (+WHA)	0.63	0.20	0.17	5.9	0.8	3.0	0.9	0.8
BB (-WHA)	0.60	0.21	0.14	5.9	0.8	2.8	1.0	0.7
BS [inflation] (+WHA)	0.49	0.21	0.21	4.3	-	2.1	0.9	0.9
BS [inflation] (-WHA)	0.36	0.24	0.18	4.3	-	1.5	1.0	0.8
BS [RER] (+WHA)	-0.81	-0.29	-0.52	3.5	-	-2.8	-1.0	-1.8
BS [RER] (+WHA, DST)	-0.84	-0.38	-0.53	2.8	-	-2.4	-1.1	-1.5
BS [RER] (-WHA)	-0.58	-0.38	-0.62	3.5	-	-2.0	-1.3	-2.2
BS [RER] (-WHA, DST)	-0.63	-0.42	-0.60	2.8	-	-1.8	-1.2	-1.7

[a)] In percentage points.
[b)] Inflation is artificial value-added deflator, composed solely of the NACE sectors C through O.
* The average of the FMOLS and DOLS estimate.
BB stands for Baumol-Bowen, BS for Balassa-Samuelson.
+WHA – wage homogeneity assumption
-WHA – without wage homogeneity assumption
DST – different share of tradables

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 two 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

Table 9: Estimates of Baumol-Bowen and Balassa-Samuelson effects (1999–2008)

Model specification	Estimates			Average of the independent variable in the sample	Share of the non-tradable sector	Effects estimates contribution of productivity differential to inflation[b)] or RER appreciation[a)]		
	LR^*	SR^{ECM*}	SR	$d\left(\bar{l^{diff}}\right)$		LR^*	SR^{ECM*}	SR
BB (+WHA)	0.63	0.20	0.17	5.4	0.8	2.7	0.9	0.7
BB (-WHA)	0.60	0.21	0.14	5.4	0.8	2.6	0.9	0.6
BS [inflation] (+WHA)	0.49	0.21	0.21	3.7	-	1.8	0.8	0.8
BS [inflation] (-WHA)	0.36	0.24	0.18	3.7	-	1.3	0.9	0.7
BS [RER] (+WHA)	-0.81	-0.29	-0.52	2.8	-	-2.3	-0.8	-1.5
BS [RER] (+WHA, DST)	-0.84	-0.38	-0.53	2.2	-	-1.8	-0.8	-1.2
BS [RER] (-WHA)	-0.58	-0.38	-0.62	2.8	-	-1.6	-1.1	-1.7
BS [RER] (-WHA, DST)	-0.63	-0.42	-0.60	2.2	-	-1.4	-0.9	-1.3

[a)] In percentage points.
[b)] Inflation is artificial value-added deflator, composed solely of the NACE sectors C through O.
* The average of the FMOLS and DOLS estimate.
BB stands for Baumol-Bowen, BS for Balassa-Samuelson.
+WHA –wage homogeneity assumption
-WHA – without wage homogeneity assumption
DST – different share of tradables

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 1.0-2.2 additional percentage points 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 1.8-2.8 (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 estimation period 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.

These results explain more of the respective cross-region and cross-sector price level

Table 10: 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
$d\left(p_{pl}\right)^{*}$	6.0	2.0	0.7 – 1.0	0.6 – 0.9	2.8 – 3.0	2.6 – 2.7
$d\left(p_{pl_ea}^{diff}\right)$	4.9	3.8	0.8 – 1.0	0.7 – 0.9	1.5 – 2.1	1.3 – 1.8
$d\left(Q_{GDP}\right)$	-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 to O.

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 (Table 10). 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

Table 11: The results of panel unit root tests (equation variables)

Variable	p-value		test statistics
	Im-Pesaran-Shin	Fisher ADF	Pesaran CIPS[1]
p_{pl}^{diff}	0.60	0.16	-2.22
$d\left(p_{pl}^{diff}\right)$	0.00	0.00	-3.09***
l_{pl}^{diff}	0.97	0.98	-1.81
$d\left(l_{pl}^{diff}\right)$	0.00	0.00	-2.59**
$p_{pl_ea}^{diff}$	0.13	0.11	-2.14
$d\left(p_{pl_ea}^{diff}\right)$	0.00	0.00	-3.22***
$l_{pl_ea}^{diff}$	0.76	0.67	-2.17
$d\left(l_{pl_ea}^{diff}\right)$	0.00	0.00	-2.77**
Q_{GDP}	0.00	0.00	-1.90
$d\left(Q_{GDP}\right)$	0.00	0.00	-4.78***
$l_{pl_members}^{diff}$	0.81	0.76	-1.39
$d\left(l_{pl_members}^{diff}\right)$	0.00	0.00	-2.55**
$l_{pl_members}^{*diff}$	0.33	0.19	-1.42
$d\left(l_{pl_members}^{*diff}\right)$	0.05	0.00	-2.59**

[1]The critical values for CIPS test are tabulated in Pesaran (2003).
One, two and three asterisks indicate statistical significance at the level of 10%, 5% and 1%, respectively.

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.

Table 12: 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)
	t most 2	-	-	30.62 (0.10)	25.58 (0.38)
		Westerlund Panel Cointegation Tests			
G_τ		−2.34 (0.10)*(0.02)	−1.47 (0.25)*(0.07)	−1.42 (0.09)*(0.08)	−1.97 (0.66)*(0.57)
P_τ		−7.17 (0.04)*(0.02)	−3.33 (0.01)*(0.00)	−1.44 (0.08)*(0.07)	−6.14 (0.62)*(0.59)

*Bootstraped critical values.
Note: Westerlund statistics were computed via the *xtwest* procedure implemented in STATA (see Persyn and Westerlund (2008) for reference)

The results are only slightly lower when we consider 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), and (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 models, taking into account the difference in sizes of non-tradable sectors between economies), we could narrow the ranges of the Balassa-Samuelson effect both with respect to inflation and real exchange rate to 0.8-0.9 in the short-run and 1.3 -1.8 in the long-run.

Table 13: 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)

*Bootstraped critical values.
Note: Due to small size of the panel Westerlund statistics could not be computed.

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 channeled fully through the domestic price growth, an annual appreciation of around 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.

Acknowledgements

The views expressed are those of the authors and do not necessarily reflect those of the institutions they are affiliated with. The authors are grateful to Aleksander Welfe, Marek Rozkrut, Joanna Bęza-Bojanowska, and the participants of 11th Ph.D. Workshop in Econometrics and Statistics in Skorzęcin in June 2010 as well as an anonymous Referee for very useful remarks.

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