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The Exchange Rate and Two Price Inflations in Poland in the Period 1999–2009. Do Globalization and Balassa–Samuelson Effect Matter?

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Abstract

The abrupt depreciation of the zloty during the subprime crisis and fast-rising prices are serious problems, because Poland, having to fulfil five Maastricht criteria, makes the dependence of her domestic inflation on price increases in the EU countries the central point of the discussion about the optimal monetary and fiscal policy rules for the next few years. The primary objective of the paper is to test out some hypotheses about the main sources of the volatility of the Polish zloty / euro exchange rate and inflation in Poland. Because several competing theoretical models describing inflationary processes are widely used, special attention is paid to their empirical verification. The working-hypotheses allowing for the country-specific features of the consumer and producer price inflation are formulated and verified in the paper.

Keywords: cointegration, exchange rate, Balassa-Samuelson effect, price-wage loop

JEL Classification: E31, F31, C51, C32

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1 Preliminaries

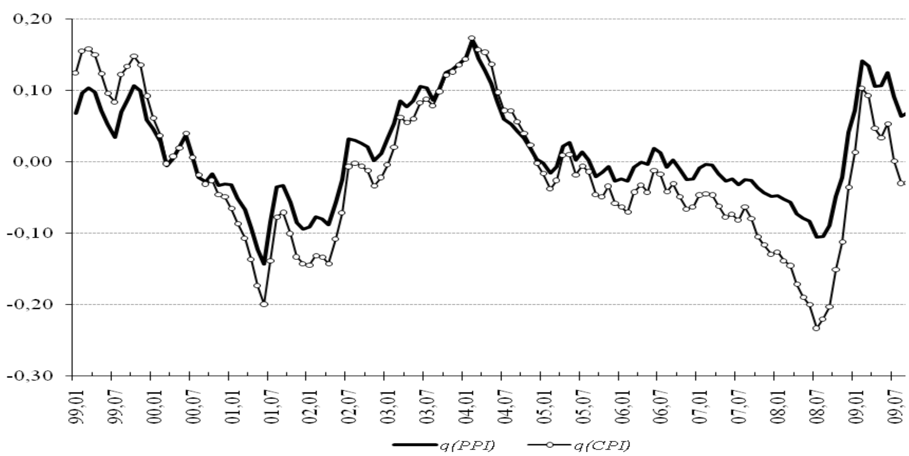
The usual account of globalization treats liberalization of international trade, the disappearance of the borders impeding foreign investments and the opening international labour market as the main factors driving global rates of GDP growth. This perspective is not changed by the conclusions offered by the analysis of the tumbling down *subprime* markets in the USA. As regards Poland, the immediate effect of the crisis was weaker transmission of the demand and supply-side shocks than in most European countries. The gently decelerating economic activity in Poland in 2009 may therefore symptomize a limited transmission of external shocks and be attributed to Poland's underdeveloped financial markets, as well as, the relative isolation of the country's economy. While the above characteristics should not be ignored, the cause of the deep depreciation of the nominal and real zloty exchange rates against major world currencies remains the key aspect in the analysis of economic processes in Poland.

The abrupt depreciation of the zloty and fast-rising prices are serious problems, because Poland, having to fulfil five Maastricht criteria, makes the dependence of her domestic inflation on price increases in the EU countries a central point of the discussion about optimal monetary and fiscal policy rules for the next few years. When the impacts of the 2008-2009 crisis are analysed from this angle, then the examination of globalization effect on Polish economy can be reduced to the import of foreign inflation (and to the possible violation of the inflationary criterion) in the first approximation. This is why the primary aim of this paper is to test out some hypotheses about the main sources of inflation in Poland after the country joined the European Union. This perspective prompts several fields of research. Firstly, as several competing theoretical models describing inflationary processes are available, a special attention should be paid to their empirical verification and to the formulation of the working-hypotheses, taking account of the country-specific characteristics of price or wage inflation. Secondly, the major determinants of the Polish zloty/euro exchange rate (hereafter PLN/EUR) need to be precisely identified.

An overview of the literature clearly indicates that in the analyses of inflation processes in the catching-up economies and of their exchange rates, the significance of the Balassa-Samuelson effect (hereinafter B-S) is strongly emphasised. The work by Halpern and Wyplosz (1997) presenting the results of their empirical investigations into the relationships between exchange rates in the Central-East European transition economies and into structural changes in these countries is widely acknowledged as a turning point in the research on the exchange rates of CEECs' currencies. As a result, the rising interest in the Balassa-Samuelson effect has brought numerous studies, most of which attempted to verify empirically whether, or not, the B-S effect was present, and to determine its magnitude (recently: Konopczak and Torój (2011), see also Égert, Halpern, MacDonald 2006 for a review of the literature and references). However, the conclusions that may be drawn from the existing studies dealing with the B-S effect are ambiguous in our opinion. In particular, verification of the relationship between relative productivities and the appreciation of CEECs' currencies turns out to be sensitive to the empirical models' specifications and to the level of inconsistency between the B-S model's assumptions and the real processes in the catching-up economies.

An affined problem, which is equally important but definitely underestimated in our opinion, is the somewhat automatic tendency to extend the CEEC exchange rate models to include the B-S effect. There are many reasons for which this approach deserves criticism, but two seem essential. Firstly, the B-S model is in fact a model of the overall price indices and its use in the exchange rate analyses can be reduced to correcting only the estimates of the real equilibrium exchange rates. Secondly, extending the analysis of the exchange rate determinants to allow for the supply-side factors is equivalent to making many restrictive assumptions, among which at least one, i.e. the purchasing power parity hypothesis (PPP) for tradable prices, is not fulfilled.

Figure 1: PPI- and CPI-based real exchange rate PLN/EUR (natural logarithms, 2000=0).

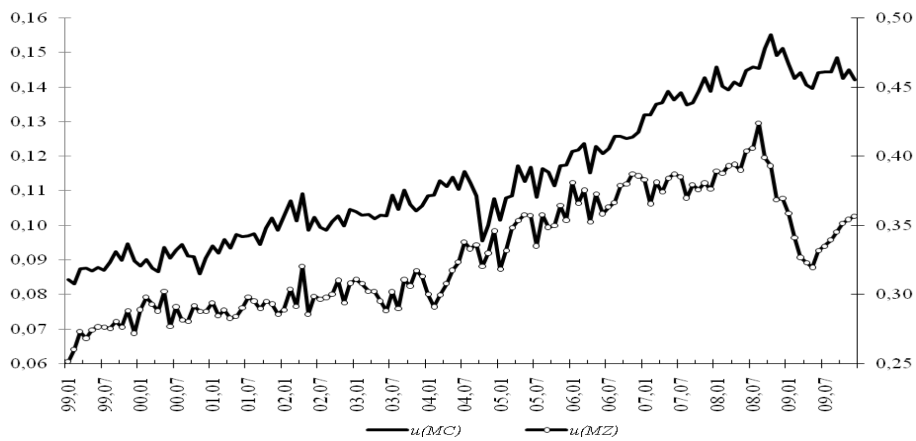


There are several hypotheses explaining why real exchange rates defined for the tradable prices in the European transition countries can be non-stationary (Figure 1). The natural appreciation hypothesis assumes that the PPI-based exchange rate can be appreciating because the CEECs' currencies were strongly undervalued in the early transition period (Halpern and Wyplosz 1997, Krajnyák and Zettelmeyer 1998). Égert and Lommatsch (2003) accentuate that appreciation may occur, because the improving quality of domestic goods and the changing structure of domestic demand make tradable prices grow faster. The basic drawback of the above explanations is that they emphasise the importance of adjustments observed in the early transition period, which are empirically undistinguishable, at least in the Polish case, from the effects of economic policy using the exchange rate as an anti-inflationary anchor. Further, Bęza-Bojanowska and MacDonald (2009) and Bęza-Bojanowska (2009) stress that the main reason for the Polish zloty/euro PPI-based real exchange rate to appreciate is the indirect impact of the B-S effect transmitted through the non-tradable component of the tradable prices. However, their results pointing to the Balassa-Samuelson effect having strong influence on the PPI-based real exchange rate may be due to the time span analysed (sample 1999:03-2007:12). In particular, the conclusion about a steady appreciation trend may result from a nominal 'anomaly in appreciation' that took place in the period 2007:01-2008:07, just before the subprime crisis occurred (see Figure 1). Further, the depreciation of the zloty observed between August and September 2008 does not necessarily have to mean that the financial crisis deflected then the exchange rate from the equilibrium appreciation path, because depreciation can also be viewed as a reaction to fluctuations in other fundamental variables, particularly to increasing risk premium. In the latter case, the zloty's depreciation noted in the second half of 2008 can be reflecting an adjustment process induced by risk premium which rose after a period of steady appreciation of the currency.

Kelm (2010), (2011) as well as Kębłowski and Welfe (2011) who support the latter perspective stress the importance of risk as a determinant of the Polish zloty exchange rate. According to the authors, the exchange rate's volatility after 1995 discourages the analyses of its non-stationarity with models predicting systematic appreciation of the zloty in that period. A more appropriate approach would be to use systems capable of identifying the sources of the transient yet strong appreciations in the years 2001-2002 and 2007-2008, of the abrupt depreciation in the 2003 and of the persisting effects of the *subprime* crisis.

The criticism of the attempts to analyse the exchange rate determinants and to quantify the Balassa-Samuelson effect within its simplified reduced-form comes down to challenging the fundamental assumption of the B-S model that in a small and open economy tradable prices are determined by the global markets. The implications of the PPP hypothesis for tradable prices are significant: an attempt at constructing the standard reduced forms of the B-S model containing overall price indices and an exchange rate may result in specification errors. The errors may prompt a false conclusion that the B-S effect is responsible for this part of exchange rate's volatility (and of the overall price indices) that in fact arises from the inadequacy of the PPP hypothesis. In this case a model including tradable prices and the real exchange rate should be constructed and tested in the first step. The model can be subsequently extended by adding medium-run imbalances on the balance of payments capital account and risk premium approximations, which ultimately leads to the capital-enhanced or behavioural equilibrium exchange rate models (CHEER or BEER; for a more detailed discussion see Clark and MacDonald 1999, Égert et al. 2006, MacDonald 2007). Controlling for the tradable price determinants brings down the problem of modelling the overall price index to including the B-S effect into the specification of the exchange rate model in the second step. Slightly simplifying the case, this means that variables cointegrated with the difference between CPI- and PPI-based real exchange rates are being sought. The last thing to do is to determine the direct effect of consumer import prices on domestic prices. The growing openness of Polish economy invalidates a frequently adopted assumption that expanding production and consumption generate approximately proportional increases in the demand for imports (see Figure 2). Moreover, the intensifying impact of the external production gap accompanied by a diminishing role of domestic unit labour costs as the determinants of domestic prices should be considered. The

Figure 2: Consumer imports / consumption ratio (u_M^C, left hand scale) and intermediate imports / industrial output sales ratio (u_M^Z, right hand scale).



paper has been structured to respond to the above issues and questions. Its first section discusses the most important problems that have to be solved to model the exchange rate, as well as presenting the specification and estimation results for a PLN/EUR exchange rate model. The final result is a vector error correction model that combines purchasing power parity with uncovered interest rate parity extended with an empirically selected risk premium proxy. This part of the paper greatly draws on Kelm's earlier results (Kelm 2010b, 2011, see also Kelm and Będza-Bojanowska 2005) and basically presents the updated estimates of the long-term parameters. The description of the mechanisms determining the PLN/EUR exchange rate is followed by the presentation of two complementary models that account for inflationary tensions. In the second section, the purchasing power parity for tradable prices is combined with the Balassa-Samuelson effect. In the resulting model, consumer prices depend on prices in the tradables sector, consumer import prices and the B-S effect proxy. Section three discusses a wage-price loop model comprised of producer prices, unit labour costs, import prices and labour market disequilibria. The paper closes with conclusions.

2 Polish zloty / euro exchange rate model

As well as being conceptually the simplest, the exchange rate model that takes advantage of the purchasing power parity produces results that raise most doubts as to their reliability (the lower case letters denote natural logarithms in this paper):

$$q = b - p + p^* \tag{1}$$

where: b - a nominal exchange rate defined as the price of a foreign exchange unit in the domestic currency, p, p^* – indices of domestic and foreign prices.
The PPP hypothesis is very controversial and the arguments against it are well known. There are serious reservations about the very concept of generalizing the law of one price to aggregated price indices and about the assumption that arbitrage is effective in the tradable market. Building the analysis on the PPP model is equivalent to making a series of usually overly restrictive assumptions, such as homogenous tradables, international tradability of all goods used to define the aggregated price indices, an insignificant role of transportation, information gathering and processing costs, non-existence of customs barriers and protectionism. Ignored are equally important problems of market monopolization, *pricing-to-market* and short-term nominal rigidities that decelerate price adjustments (see reviews: Officer 1976, Froot and Rogoff 1995, Sarno and Taylor 2002, more recently: MacDonald 2007).

2.1 Theoretical framework

An overview of the early empirical studies reveals that a very long sample span or large panel data are required for the PPP hypothesis to be accepted. The real exchange rates are mainly characterised by a very slow mean-reverting process, which implies that the PPP model is incapable of explaining longer periods during which the nominal exchange rate deviates from the PPP-determined trajectory. However, the more recent econometric studies clearly indicate that the last conclusion is not so obvious in the case of models allowing for the presence of transaction costs and of a real exchange rate fluctuation band without price arbitrage. The standard setting for analysing the potential non-linearity of the PPP model is the smooth transition autoregressive models (STAR), where the real exchange rate adjusts much faster outside the no-arbitrage band than the typical consensus half-life estimates for 3-5 years might suggest (e.g. Froot and Rogoff 1995). However, this approach also attracts critique. In particular, MacDonald (2007) stresses that the assumption about transaction costs exerting insignificant impacts (or proportional to the relative prices) is only one of a number of assumptions underlying the PPP model. Therefore, trying to connect the non-linear mean-reversion of the real exchange rate only with

the arbitrage costs may be perceived as an overly simplified approach.
In line with the above critique, the exchange rate modelling problems are frequently solved by assuming that the analysis of exchange rates is typically medium-term and by making more detailed analysis of the balance of payments disequilibria (e.g. Juselius 1995, Clark and MacDonald 1999, Juselius 2006). This approach allows examining the interdependencies between the real exchange rate and the capital account imbalance persisting in the medium term can be considered. Why the imbalance appears can be explained with the twin-deficit model where the imbalance of the external sector is attributed to the fiscal sector's deficit or with the internal-external balance models emphasising the relationships between the balance of payments, unbalanced domestic savings and investments. The natural solution would therefore be to apply the uncovered interest rate parity (UIP):

$$b_t = E_t\left(b_{t+k}\right) - k\left(i_{t,k} - i^*_{t,k}\right) + \lambda, \tag{2}$$

where: i, i^* denote the domestic and foreign nominal interest rates for assets with maturities of k months, λ – the risk premium, E – expectations operator.
The simplest approach adopted within the empirical applications of the UIP equation (2) assumes rational exchange rate expectations and constant (or stationary) risk premium. Although the rational expectations hypothesis (REH) continues to be one of the key approaches used in macroeconomic analysis, alternative models of expectations offering better description of the data generation process in the case of shorter samples are becoming increasingly useful. In particular, effective knowledge accumulation and adaptive learning are the cornerstones of the new IKE models (*imperfect knowledge economics*, Frydman i Goldberg 2007, also: Frydman and Goldberg 2008, Frydman, Goldberg, Johansen, Juselius 2008, Juselius 2010), in which economic agents are assumed to have incomplete knowledge about the economic system they function within. The main differences between models assuming imperfect knowledge and those founded on rational expectations lie in the fact that the first type rejects the possibility that the mechanism shaping expectations is predeterminable. Economic agents in the IKE models have incomplete knowledge of the mechanisms determining the variables for which they formulate their expectations on and constantly correct their forecasting models.
The consequences of replacing the REH with the IKE assumptions are difficult to overestimate. Juselius (2010) considers a UIP model where the IKE assumptions extend equation (2) to time-varying risk aversion a^{ν}:

$$E_t\left(b_{t+1}\right) - b_t = i_{t,1} - i^*_{t,1} - a^{\nu}_t. \tag{3}$$

Because a^{ν} is a non-observable variable, its approximation poses an empirical problem. Frydman and Goldberg (2007) and Juselius (2010) propose a solution where the loss averseness premium is proportional to nominal exchange rate deviations from the path set by relative prices:

$$\Delta b_{t+1} = \left(i_{t,1} - i^*_{t,1}\right) - \omega_1\left(b_t - p_t + p^*_t\right). \tag{4}$$

The concept where the fluctuations in the nominal exchange rate are combined with the medium-term deviations from the PPP path is not new and the arguments behind it are similar to those underpinning the capital-enhanced equilibrium exchange rate models that Katarina Juselius proposed at the beginning of the 21st c. (e.g. Juselius 1995, Juselius and MacDonald 2000, 2004, 2006; for Polish zloty see Welfe, Karp, Kębłowski 2006, Stążka 2008, Kębłowski and Welfe 2010, Kelm 2011). It is argued that the exchange rate analysis should simultaneously cover processes taking place in (i) the goods markets that are in equilibrium when the PPP hypothesis holds and in (ii) the capital markets that remain in balance owing to the mechanisms described by the UIP model. Juselius (1995) uses the simplest UIP model comprised of a nominal exchange rate, domestic and foreign prices and domestic and foreign nominal interest rates to consider and test the hypothesis about exchange rate expectations being influenced by relative prices and nominal interest rates disparities:

$$E_t\left(b_{t+1}\right)=\omega_1\left(p_t-p_t^*\right)+\omega_2\left(i_{t,1}-i_{t,1}^*\right). \tag{5}$$

In the more complex framework that allows for joint testing of (i) the purchasing power parity, (ii) the uncovered interest rate parity, as well as (iii) the term structure of the interest rates and (iv) the real interest rate parity Juselius and MacDonald (2000), (2004), (2006) connect exchange rate expectations with real exchange rate fluctuations and inflation and nominal short-term interest rate differentials:

$$E_t\left(\Delta b_{t+1}\right)=-\omega_1 q_t+\omega_2\left(\Delta p_{t+1}-\Delta p_{t+1}^*\right)+\omega_3\left(i_{t,1}^S-i_{t,1}^{*S}\right). \tag{6}$$

where: $E_t\left(\Delta b_{t+1}\right)=E_t\left(b_{t+1}\right)-b_t$.

2.2 Empirical results

In the empirical analysis being presented, the standard vector error correction model (VEC, e.g. Juselius (2006)) was employed:

$$\Delta y_t=\alpha\beta' y_{t-1}+\sum_{s=1}^{S-1}\Gamma_s\Delta y_{t-s}+\mu+\Psi d_t+u_t, \tag{7}$$

where: y – the vector of endogenous variables, d – the vector of deterministic variables, α – the matrix of adjustment parameters, β – the matrix of V cointegrating vectors, Γ – the matrix of the short-term parameters, Ψ – the matrix of the deterministic variables' parameters, u – error terms. Under the long-run weak exogeneity of y^x, the VEC model takes the following form:

$$\Delta y_t^E=\alpha^E\beta' y_{t-1}+\sum_{s=1}^{S-1}\Gamma_s^E\Delta y_{t-s}+\Theta\Delta y_t^x+\mu^E+\Psi^E d_t+u_t^E, \tag{8}$$

where y_t^E – $(M-H)\times 1$ – the vector of the endogenous variables and y^x – $H\times 1$ – the vector of the weakly exogenous variables.

The starting point was a vector equilibrium correction representation of the CHEER model with time-varying risk premium:

$$y_t = [b_t, p_t, p_t^*, i_t, i_t^*, \lambda_t]. \tag{9}$$

Two problems had to be solved before the model could be constructed.
Firstly, a serious interpretational problem is likely to arise when model (9) is estimated with standard I(1) cointegration procedures disregarding the fact that some variables are generated by I(2) processes (see Juselius 2006). This problem is quite important, as there are strong arguments supporting the statement that domestic prices, and thus the nominal EUR/PLN exchange rate, are 'almost-I(2)' variables (e.g. Kelm and Majsterek 2006, 2007). There are also grounds for concluding that the price-generating processes in the euro area countries where price inflation is relatively low and stable may also contain double unit roots (e.g. Juselius 1995, 2006). If so, a medium-term cointegration analysis is necessary to estimate the PLN/EUR exchange rate. There are two alternative options: (i) the VEC model can be transformed to make it explicitly allow for the presence of variables integrated of order 2, or (ii) the variables can be transformed within a standard VEC model (7), so that the elements of vector are integrated of order one. The second approach (a nominal-to-real transformation or an I(2)-in-I(1) analysis) is empirically justified, if the processes generating the I(2)-variables have common double unit root (see Juselius 2006). In the case of the CHEER model (9), the I(2)-in-I(1) cointegration analysis involves the acceptance of the long-term homogeneity restriction linking domestic and foreign prices and the nominal exchange rate. Then the transformed model (9) takes the following form:

$$y_t = [q_t, \Delta p_t, \Delta p_t^*, i_t, i_t^*, \lambda_t]. \tag{10}$$

The second problem is the choice of risk proxies. Theoretical analyses frequently put forward an approach where the general equilibrium model is linked with the consumption-based capital asset pricing model (C-CAPM, e.g. Smith and Wickens 2002, Groen and Balakrishnan 2006, Kočenda and Poghosyan 2009). This procedure starts with the maximization of the representative consumer utility U that can most simply be approximated by means of the exponential function:

$$U(C_t) = C_t^{1-\sigma}(1-\sigma)^{-1}, \tag{11}$$

where: C – real consumption, σ – the risk aversion parameter. For a standard budget constraint, the solution of the utility maximisation problem is a function that connects the risk premium λ with the covariance of (i) the growth rate of consumption Δc and (ii) the deviation of the growth rate of the nominal exchange rate from the nominal interest rate disparity e^R:

$$\lambda_t = \sigma \text{cov}\left(\Delta c_{t+1}, e_{t+1}^R\right), \tag{12}$$

where $e_{t+1}^R = \Delta b_{t+1} - (i_t - i_t^*)$.
The main problem that comes up when the C-CAPM models are used to analyse a

risk premium is that the model based on the power utility function (11) is usually oversimplified. It is enough, though, to remove the assumption about time separability of consumer utility to have models where the risk premium is a function of the unobservable variables (Groen and Balakrishnan 2006). In the most general case, the utility maximization problem can be solved with a factor model:

$$\lambda_t = \sum_{i=1}^{L} \beta_l \mathrm{cov}\left(z_{l,t+1}, e_{t+1}^{R}\right), \tag{13}$$

where z_l stands for factor l.
Because no precise criteria for choosing the z_k factors are available, the specification of model (10) poses an empirical problem. The recommendations available in the literature usually concentrate on the analysis of the country's fiscal position and accentuate the role of the government debt (Juselius 1995, Giorgianni 1997, Clark and MacDonald 1999). Ultimately, the risk premium's effect on the real PLN/EUR exchange rate was analysed here using variables recommended by the aforementioned studies and those variables whose effect on the PLN/EUR exchange rate has been confirmed by the earlier studies of the determinants of the Polish zloty exchange rate (Kelm and Bęza-Bojanowska 2005, Kelm 2010b, 2011). The analysis started with the identification of the relationships between the short-term governmental debt (actually the ratio between the domestic and foreign shares of the short-term government debt in GDP, U^{DST}) and the budget's deficit (the share of the budget's domestic deficit in GDP, U^{SD}), on one hand, and the internal (λ^{INT}) and external (λ^{EXT}) determinants of the exchange rate risk, on the other. Assuming that main factors expanding the short-term debt are (i) the fiscal sector disequilibrium that can be described by means of the budget deficit function and (ii) the demand for assets denominated in the Polish zlotys that fluctuates following global risk changes, the short-term debt can be written as follows:

$$U_t^{DST} = U_{t-1}^{DST} + \rho_1\left(U_t^{SD}\right) + \rho_2\left(\lambda_t^{EXT}\right), \tag{14}$$

where: $U^{DST} = \frac{DST/Y}{DST^*/Y^*}$; DST, DST^* – short-term government debt in Poland and the euro zone (current prices), Y, Y^* – nominal GDPs in Poland and euro zone.
Because the first two components are determined by domestic variables, model (14) can be equivalently written as:

$$U_t^{DST} = \widetilde{\rho}_1\left(\lambda_t^{INT}\right) + \rho_2\left(\lambda_t^{EXT}\right) = \rho\left(\lambda_t\right). \tag{15}$$

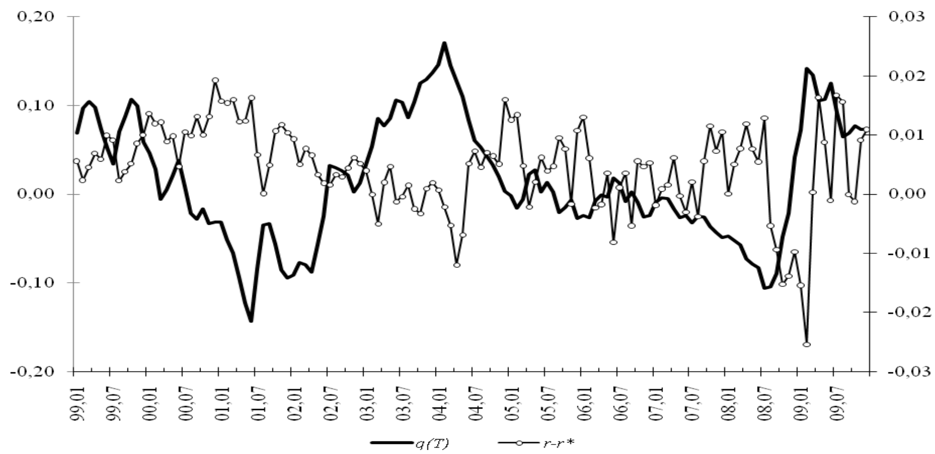
Debt U^{DST} expanding because of larger T-bill issues means that either the government has bigger problems with financing its current expenditures or that the investors are becoming less trustful of securities with longer maturity. In the first case, the reasons may be an extremely expansionary variant of fiscal policy, that the government pursues to fund its excessive expenditures by increasing the short-term debt or, in less controversial scenario, the government' reaction to suddenly falling

production. Whatever the motivation, a deep budget deficit will appear, the short-term debt will grow larger and risk premium will increase. The transmission of global risks may also cause U^{DST} to fluctuate. Because selling the long-term securities is a safer way of funding government expenditure, larger T-bill issues can be expected when the demand for bonds meets the barrier of interest rates.
To sum up, the following VEC model was verified empirically:

$$y_t = \left[q_t^T, i_t, i_t^*, \Delta p_t^T, \Delta p_t^{*T}, U_t^{DST}\right], \tag{16}$$

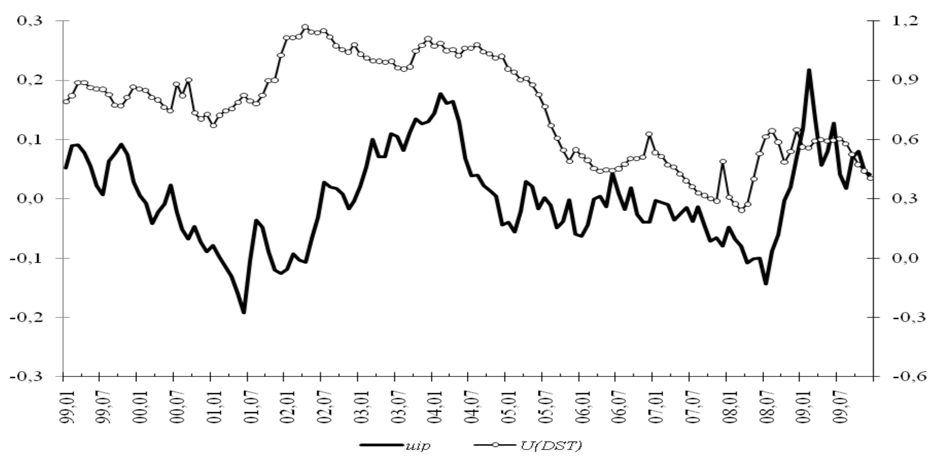
where: $q^T = b - p^T + q^{*T}$; the domestic and foreign prices (p^T, p^{*T}) in the open sectors were approximated with prices in manufacturing in Poland and the Euro Area; the nominal interest rates i and i^* were represented by three-month interest rates on loans in the inter-bank markets WIBOR 3M and EURIBOR 3M.
The empirical investigation covered the period 1999:01-2009:12. The data were derived from various sources. The domestic data were extracted from the official publications by the Polish Central Statistical Office and the National Bank of Poland. The variables that are not observable at monthly frequency (GDP) were estimated using procedures proposed in Kelm (2008). The information about the euro area was found in the OECD, EUROSTAT, ECB and Bundesbank databases. When the monthly data were not available, the quarterly data were interpolated. Graphical analysis of (i) the real exchange rate q^T, (ii) the disparity in real interest rates ($r - r^* = (i - \Delta p^T) - (i^* - \Delta p^{*T})$),(iii) the real exchange rate's deviation from the UIP path defined for three-month expectations ($uip = q^T - 3(r - r^*)$), and (iv) the risk premium proxy U^{DST} is sufficient to formulate hypotheses about common stochastic trends being present in the processes generating VEC model (16) variables (see Figures 3-4). The oscillations of the real exchange rate q^T follow those in the real interest rates disparity $r - r^*$; besides, a considerably larger amplitude of q^T oscillations can be noticed, meaning that the estimate of parameter k may exceed a value of 3 implied by equation (2). However, when the specification of UIP model (16) additionally includes the risk proxy U^{DST}, the ultimate result of the statistical verification of the model, i.e. of the hypothesis that $k = 3$, becomes uncertain because even a perfunctory comparison of the oscillations in U^{DST} and the real exchange rate deviations uip may reveal the presence of a common stochastic trend. Before the VEC model (16) parameters were estimated, two tests were performed: (i) to verify the long-term homogeneity of the nominal exchange rate e and of the deflators p^T and p^{*T} (based on the VEC model (9)) and (ii) to test risk premium for weak exogeneity (in the VEC model (16)). The long-term homogeneity was accepted. Depending on the significance levels assumed, the cointegration rank tests suggest that either 2 or 1 cointegrating vectors are present. For $V = 1$, the probability value in the LR test for long-term homogeneity restriction is 0.727, whereas in the model spanned by two cointegrating vectors the value is lower, 0.091. The latter is a borderline result, yet it still justifies accepting the homogeneity restriction and conducting the analysis with the transformed VEC model (16). Juselius (1995) states that the super-super-

Figure 3: PPI-based real exchange rate PLN/EUR (q^T, left hand scale) and differential of the real interest rates ($r - r^*$, right hand scale).



Note: An increase of the real exchange rate denotes zloty's depreciation.

Figure 4: UIP deviations (uip, left hand scale) and risk premium proxy (U^{DST}, right hand scale).



Note: An increase of the real exchange rate denotes zloty's depreciation.

consistency of the long-term parameter's estimators in the standard I(1) VEC model with the I(2) variables may make it extremely difficult to confirm the hypotheses about the long-term structure of the model. From this perspective, the long-term homogeneity restriction in model (9) seems fully acceptable.

The results of the weak exogeneity tests for risk proxy are unequivocal and the prerequisite for conditioning the VEC system (16) on U^{DST} is very strong regardless of the number of the cointegrating vectors; for $V = 1$ and $V = 2$ the p-values in the weak exogeneity test sufficiently exceed the standard significance levels (0.972 and 0.416, respectively) and for $V = 3$ and $V = 4$ they are still above the borderline (0.169 and 0.237).

With the above results in mind, the exchange rate model was derived from the VEC model with a weakly exogenous relative short-term debt:

$$y_t^E = \left[q_t^T, i_t, i_t^*, \Delta p_t^T, \Delta p_t^{*T}\right], \quad y_t^x = \left[U_t^{DST}\right] \tag{17}$$

The trace cointegration tests (with Bartlett correction) suggest that two or three cointegration vectors should be considered (see table 1). However, the four characteristic roots of the companion matrix lying very close to the unit circle suggest that only one cointegrating vector is present. As the results are slightly different, a compromise variant of the VEC model (17) was finally adopted: the presence of two cointegrating vectors meets the assumption requiring that the specification of the exchange rate model should contain the mechanisms described by both the PPP and UIP equations.

The estimates of the adjustment parameters show that the producer price inflation Δp^T oscillates along an equilibrium trajectory defined by the first cointegrating vector; the second cointegrating vector defines an equilibrium path to which the real exchange rate q^T adjusts. The structuralizing restrictions on the long-run parameters are

Table 1: PPP-UIP model (17) - cointegration tests and companion matrix roots.

V	Trace	Trace(B)	p-value	p-value (B)
0	166.4	146.7	0.000	0.000
1	86.38	74.43	0.001	0.023
2	48.22	41.70	0.050	0.184*
3	15.81	13.87	0.744*	0.857
4	0.90	0.71	0.999	0.999

Roots	Real	Imaginary	Modulus
1	0.975	0.031	0.976*
2	0.975	-0.021	0.976*
3	0.864	0.119	0.872*
4	0.864	-0.119	0.872*
5	0.362	0.510	0.625
6	0.362	-0.510	0.625

Note: (B) indicates *Trace* test with Bartlett correction.

fully consistent with the model specification hypotheses, i.e. with the UIP model (containing a risk premium) and the PPP equation (see table 2). A normalizing restriction and three structuralizing restrictions are imposed on the first cointegrating vector. According to the restrictions, the real exchange rate is defined by (i) the

Table 2: PPP-UIP model (17) – estimates and diagnostics.

	q	Δp	i	Δp^*	i^*	U^{DST}	t
β_1	-0.0232 (4.0)	**1**	0	0	0	–	0.00003 (3.0)
β_2	1	-5.111 (3.6)	5.111 (3.6)	5.111 (3.6)	-5.111 (3.6)	-0.144 (5.1)	0
α_1 (2.2)	-0.957 (6.9)	-0.759 (5.5)	0.027 (2.4)	-0.200	0.004 (2.1)	–	–
α_2	-0.131 (4.2)		0.002 (5.8)	0.013 (2.2)	–	–	–
LR = 0.353							
AR(1) = 0.125 AR(2) = 0.086				DH = 0.065			
AR(3) = 0.177 AR(4) = 0.424				ARCH(1) = 0.817 ARCH(2) = 0.887			

Note: t-ratios are reported in parentheses. Dots stand for the parameters with t-ratios smaller than 2. LR is a likelihood ratio test for over-identifying restrictions, DH – Doornik-Hansen normality test, AR(s) – autocorrelation LM test, ARCH(s) – ARCH effect test (for details see: Juselius 2006); p-values are reported for LR, AR, DH and ARCH tests.

difference between the Polish real interest rates and those in the euro zone and by (ii) the risk premium. The second cointegrating vector is normalized against price inflation in manufacturing in Poland. Summing up, the estimates of the long-run parameters are as follows:

$$\Delta p_t^T = \underset{(4.0)}{0.0232} q_t^T - \underset{(3.0)}{0.00003} t, \tag{18}$$

$$q_t^T = -\underset{(3.6)}{5.111}\left(\left(i_t - \Delta p_t^T\right) - \left(i_t^* - \Delta p_t^{*T}\right)\right) + \underset{(5.1)}{0.144} U_t^{DS}. \tag{19}$$

The LT test' results allow accepting the above model structure for standard significance levels. The results show that the residuals are Gaussian, without major autocorrelation and heteroskedasticity symptoms like the variants of model (16) discussed above.

The presence of a deterministic trend makes it difficult to interpret the results obtained for the first cointegrating vector. Although a more in-depth analysis of the determinants of price inflation in the Polish open sector is provided in section three, two important conclusions should be presented already now. Firstly, price dynamics in the open sector follows the growth in the nominal exchange rate and in foreign prices. At the same time, the domestic price inflation declines when domestic prices lie above the path determined by the PPP model. Secondly, the structure of the cointegrating vector (18) proves that there is a mechanism that balances prices and the exchange rate in line with the PPP model defined for price indices in the tradable sectors; in the medium run the combination $e - p^T + p^{*T}$ is integrated of order 1, which means that a CI(2,1) cointegrating relationship exists between the nominal exchange rate and the prices.

The interpretation of the long-run parameters and the adjustment parameter in the real exchange rate equation q^T is quite straightforward:

(i) the impacts of the domestic and foreign interest rates are symmetric,

(ii) the estimate of the parameter at the real interest disparity suggests that the time horizon of exchange rate expectations only slightly exceeds five months,

(iii) the parameter at the risk premium proxy is significantly different from zero, thus confirming that the UIP model must be extended to allow for risk fluctuations,

(iv) the increased relative short-term debt leads to zloty depreciation as assumed,

(v) the error correction term in equation Δq^T shows that the real exchange rate adjusts to the identified equilibrium path relatively fast – the deviations from the path decrease at more than 13% a month, so the half-life of the real exchange rate's deviations from the trajectory (19) is below six months.

Finally, the estimates of the long-run parameters were tested for stability. The recursive estimation results obtained for the period 2004:07-2009:12 did not reveal any major oscillations in the price parameter estimates (see Figure 5). This finding basically supports the use of model (18)-(19) as a starting point for further discussion.

3 Exchange rates, consumer prices and the Balassa-Samuelson effect

The empirical analysis of the Balassa-Samuelson effect frequently involves the modelling of the exchange rate (for detailed review see Égert, Halpern, MacDonald 2006 and the references therein). In the standard procedure, the real exchange rate is defined for the overall price indices and then its trends driven by prices that grow faster in the more expanding economies are appropriately approximated. From the empirical perspective, two modelling approaches seem particularly attractive. Firstly, assuming that the general level of domestic and foreign prices is a weighted average of the tradable and non-tradable prices, the real exchange rate for the overall price indices q can be defined using the real exchange rate for tradable prices q^T and the relative rates of price inflation in domestic and foreign sheltered sectors:

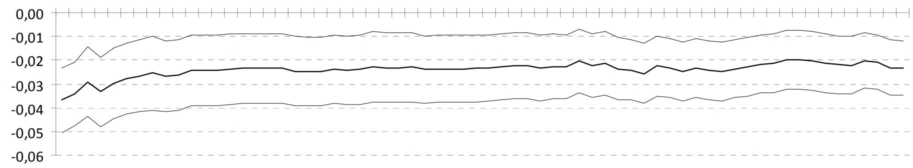
$$\Delta q_t = \Delta q_t^T - (1-\tau)\left(\Delta p_t^{NT} - \Delta p_t^T\right) + (1-\tau^*)\left(\Delta p_t^{*NT} - \Delta p_t^{*T}\right), \qquad (20)$$

where: $p = \tau p^T + (1-\tau)p^{NT}$, $p^* = \tau^* p^{*T} + (1-\tau^*)p^{*NT}$, p^{NT}, p^{*NT} – the indices of non-tradable prices, τ, τ^* – tradables' shares in the overall price index at home and abroad.

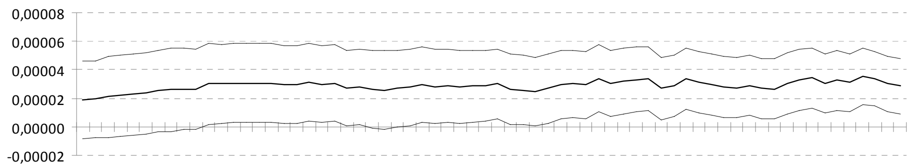
Figure 5: Recursive estimation of the long-term parameters

a) model (18) – producer price inflation, 2004:07-2009:12.

$\beta_1(q^T)$

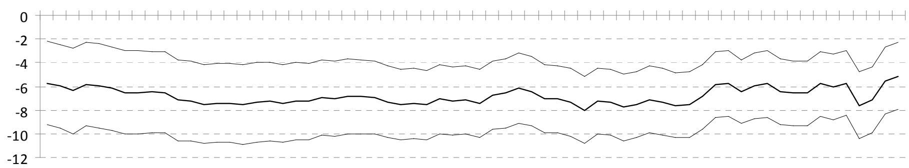


$\beta_1(t)$

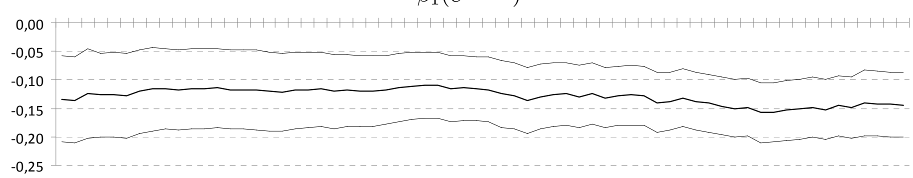


b) model (19) - real exchange rate, 2004:07-2009:12.

$-\beta_2(i) = \beta_2(\Delta p^T) = \beta_2(i^*) = -\beta_2(\Delta p^{*T})$



$\beta_1(U^{DST})$



In an alternative approach, the Balassa-Samuelson model is derived from a two-factor Cobb-Douglas production function with constant returns to scale:

$$\Delta q_t = \Delta q_t^T - (1-\tau)\left(\gamma\beta^{-1}\Delta a_t^T - \Delta a_t^{NT}\right) + (1-\tau^*)\left(\gamma^*\beta^{*-1}\Delta a_t^{*T} - \Delta a_t^{*NT}\right), \quad (21)$$

where: a^T, a^{NT}, a^{*T}, a^{*NT} – the logarithms of TFP in the tradable and non-tradable sectors at home and abroad, γ, β, γ^*, β^* – the parameters of the production function for tradable ($Y = AK^{\beta}L^{1-\beta}$) and non-tradable ($Y = AK^{\gamma}L^{1-\gamma}$) sectors at home and abroad.

Seeking the correct definitions the prices or TFP in the tradable and non-tradable sectors is as an important stage in investigations using specifications (20)-(21) as problematic, because it is extremely difficult to draw a clear line between the open and sheltered sectors. The consequences of misspecification can be serious:

because the hypotheses about parameters τ and τ^* may be impossible to confirm, the satisfactory accuracy of their estimates and the estimates' being as expected are frequently recognised as sufficiently supporting the thesis that the B-S effect is present. Other shortcomings of the models (20) and (21) stem from the fact that the models describe the sources of oscillations in the real exchange rate q, while the mechanisms determining the overall price indices remain hidden. Therefore, if the analysis sets out to test the B-S effect for the overall price index, then the analytical framework has to be redefined. The simplest solution consists in combining models (20) and (21) together, which allows defining an internal transmission mechanism between productivity differential and the relative prices of the non-tradable goods:

$$\Delta p_t^{NT} - \Delta p_t^T = \gamma\beta^{-1}\Delta a_t^T - \Delta a_t^{NT}. \tag{22}$$

Because the overall price index is comprised of the tradable and non-tradable prices, it is easy to derive an identity that defines the overall price index as a function of tradable goods' prices and TFP differential:

$$\Delta p_t = \Delta p_t^T + (1-\tau)(\Delta a_t^T - \Delta a_t^{NT}). \tag{23}$$

where $\beta = \gamma$ is assumed.

For model (23) to be useful for analysing the consumer price index three additional problems have to be tackled. Firstly, that CPI variability partly arises from changing tax rates must be taken into consideration. Secondly, the impact of the imported consumer goods and services on CPI has to be explicitly accounted for. Thirdly, an approximation of relative productivity in the tradable and non-tradable sectors has to be determined.

The first two problems are straightforward to solve (Wallis 2004, also: Kębłowski, Majsterek, Welfe 2008, Majsterek and Welfe 2010):

$$\begin{aligned} p_t^C &= \left(1-u_{M,t}^C\right)\left(p_t^T + r_t^{VAT} + r_t^{EXC}\right) + \\ &\quad + u_{M,t}^C\left(p_t^M + r_t^{VAT} + r_t^{EXC} + r_t^{DUT}\right) + \theta h_{\Sigma,t}^{BS} = \\ &= p_t^T + u_{M,t}^C\left(p_t^M - p_t^T\right) + \theta h_{\Sigma,t}^{BS} + r_t^{TAX} \end{aligned}, \tag{24}$$

where: p^C – CPI, p^M – import prices, u_M^C – the consumer imports / consumption ratio, r^{VAT} – the effective VAT rate, r^{EXC} – the effective rate of excise tax, r^{DUT} – the effective rate of customs duty, $r^{TAX} = r^{VAT} + r^{EXC} + u_M^C r^{DUT}$, h_Σ^{BS} – the proxy of the cumulated B-S effect, i.e. $h_\Sigma^{BS} = a^T - a^{NT}$.

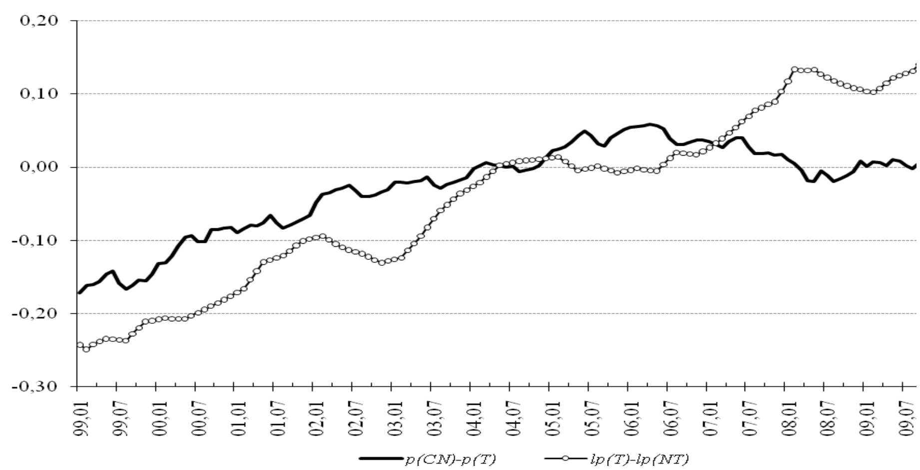
Assuming that the long-run prices of imports are a function of the nominal exchange rate and of the prices in the foreign tradable sector, i.e. $p^M = b + p^{*T} = q^T + p^T$, the CPI equation is:

$$p_t^{CN} = p_t^C - r_t^{TAX} = p_t^T + g_t^{CPI} + \theta h_{\Sigma,t}^{BS}. \tag{25}$$

where $g^{CPI} = u_M^C q^T$ is a 'supplementary tax' imposed by high import prices (Wallis 2004) or, interpreted alternatively, a growing effect of globalisation on the consumer

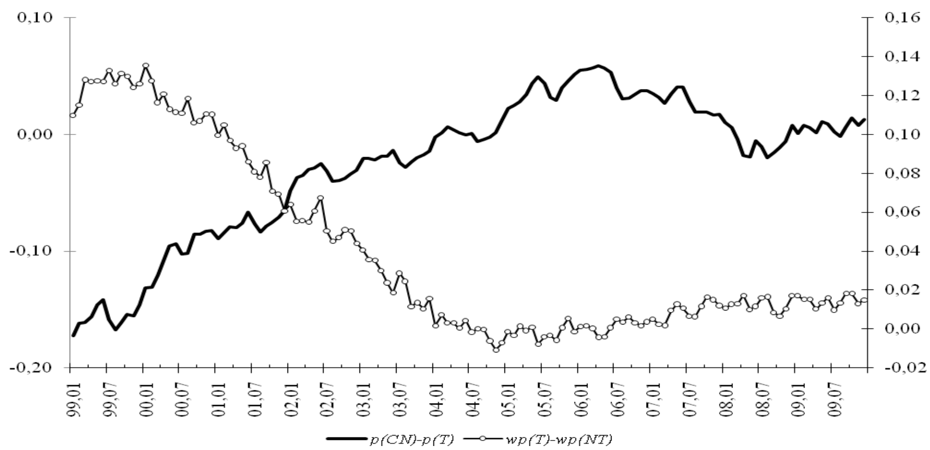
prices that comes from imported inflation and the widening openness of the domestic economy.
A direct application of the ratio between consumer imports and consumption makes the empirical analysis more complex, because a non-linear relationship is added to the VEC model then (see Figure 2). However, the u_M^C values show, that this approach is necessary, because they grow from 8-9% in 1999 to more than 14-15% at the end of 2009. This upward trend confirms that gaining importance as a price-forming factor globalisation increases the risk of exogenous supply shocks.
The criteria for choosing which variables should represent the Balassa-Samuelson effect h_Σ^{BS} are ambiguous. The standard approaches replace TFP changes with labour productivity or GDP growth rates, but this may cause specification errors and misleading results, if the growth rates of the capital-labour ratios in the tradable and non-tradable sectors are considerably different from each other in the period in question. Even a perfunctory analysis of the time series points out that the thesis about consumer prices and prices in the open sector and the B-S effect approximated with the relative labour productivity being interrelated is at least doubtful (see Figure 6). The $p^{CN} - p^T$ price wedge kept widening until the end of 2005, though at decelerating rate, and then stabilised, while the relative labour productivity $l_p^T - l_p^{NT}$ oscillated around a rising trend throughout the sample period. Relative labour productivity used to approximate the B-S effect is not enough to explain the 2005-2009 price wedge. Different conclusions may be derived when the choice of the h^{BS}

Figure 6: Price wedge ($p^{CN} - p^T$) and relative labor productivity ($l_p^T - l_p^{NT}$).



proxy is contrasted with the mainstay assumption of the Balassa-Samuelson model that adjustment processes lead to wage equalization in the tradable and non-tradable sectors. Figure 7 illustrates the changes in the price wedge $p^{CN} - p^T$ and the ratio between nominal wages in the tradable and non-tradable sectors $w_p^T - w_p^{NT}$. It is easy to see that while between 1999 and 2004 relative wages were falling and the price wedge was growing, in the period 2005-2009 the wage levels in the open and sheltered sectors were approximately equal and the CPI-to-PPI ratio was almost stable then. The above analysis supports the thesis that the wage equalization process must be

Figure 7: wedge ($p^{CN} - p^T$, left hand scale) and relative nominal wages ($w_p^T - w_p^{NT}$, right hand scale).



included if one intends to build a model quantifying the Balassa-Samuelson effect for the consumer prices. The analysis also shows that both characteristics of the B-S effect, i.e. relative wages $h_{WP}^{BS} = w_p^{NT} - w_p^T$ and relative labour productivities $h_{LP}^{BS} = l_p^T - l_p^{NT}$, have to be included in the VEC system:

$$y_t = \left[p_t^{CN}, p_t^T, g_t^{CPI}, h_{WP,t}^{BS}, h_{LP,t}^{BS}\right]. \tag{26}$$

Because model (26) embodies also tradable prices, the PPP equation (18) can be tested in a setting where the effects of imported inflation and of non-tradable processing on tradable prices can be estimated. Hence, the B-S effect was analysed with a model extended to include the real exchange rate q^T:

$$y_t = \left[p_t^{CN}, p_t^T, g_t^{CPI}, h_{WP,t}^{BS}, h_{LP,t}^{BS}, q^T\right]. \tag{27}$$

The estimation procedure was the same as that used to construct the exchange rate and producer price model in manufacturing. Model (27) where at least two variables (CPI and PPI) had the I(2) properties was considered first. The cointegration tests and the analysis of the characteristic roots of the companion matrix provided grounds for taking account of one or two cointegrating vectors. Then the following procedures were performed for $V = 1$ and $V = 2$: (i) tests for the homogeneity of CPI, PPI in manufacturing and for the globalisation effect g^{CPI} and (ii) tests for CPI and PPI homogeneity. The first of the restrictions tested defines a price wedge containing the effects of imported inflation:

$$p^{WDG}_{1,t} = p^{CN}_t - p^T_t - g^{CPI}_t; \tag{28}$$

when there are no grounds for rejecting (28), then the VEC model should be considered:

$$y_t = \left[p^{WDG}_{1,t}, \Delta p^{CN}_t, \Delta p^T_t, h^{BS}_{WP,t}, h^{BS}_{LP,t}, q^T_t\right]. \tag{29}$$

In the second case, the price wedge is simply the difference between CPI and manufacturing PPI:

$$p^{WDG}_{2,t} = p^{CN}_t - p^T_t, \tag{30}$$

and the VEC model is:

$$y_t = \left[p^{WDG}_{2,t}, \Delta p^{CN}_t, g^{CPI}_t, h^{BS}_{WP,t}, h^{BS}_{LP,t}, q^T_t\right]. \tag{31}$$

The homogeneity tests produced borderline probability values. For the VEC system (27) with one cointegrating vector the values were 0.143 and 0.053 for restrictions (28) and (30), respectively; for $V = 2$ they were slightly smaller (0.105 and 0.048). These results show that globalization has a relatively weak effect on consumer prices despite growing consumer imports; a price wedge definition taking account of g^{CPI} makes it easier to accept the homogeneity restrictions, which means that g^{CPI} variations are limited compared with the fluctuations in CPI and PPI.
Both variants of the CPI model, i.e. VEC systems (29) and (31), were tested empirically.
The estimates obtained from the VEC model (29) prove that the model has two cointegrating vectors that attract both consumer price inflation and producer price inflation, but all attempts to structuralize the long-term relationships failed and the parameter estimates appeared to be extremely unstable. It is easy to notice that the undesirable properties of model (29) come from the nominal-to-real transformation and its 'overspecification'. If globalization effects on CPI values were limited, the VEC model containing p^{WDG}_1, Δp^{CN} and Δp^T would be estimated using a time series that would approximately 'duplicate' the same information. The resulting parameter estimates would be extremely unstable, because of the near co-linearity of the variables. This problem can be easily solved by replacing in model (29) producer price inflation with the dynamics of the globalization effects Δg^{CPI}:

$$y_t = \left[p^{WDG}_{1,t}, \Delta p^{CN}_t, \Delta g^{CPI}_t, h^{BS}_{WP,t}, h^{BS}_{LP,t}, q^T_t\right]. \tag{32}$$

but even then the estimates stir doubts. Firstly, the producer price inflation removed, the rank of the cointegrating space becomes reduced to a single vector; it also becomes possible to impose the exclusion restriction on the real exchange rate, which is hardly surprising given the conclusions arising from the CHEER model (18)-(19). Secondly, the restrictions as implied by equation (25) cannot be imposed in the VEC system without the real exchange rate q^T – the probability value of the LR test for over-identifying restrictions in the VEC system spanned by the following cointegrating vector

$$\Delta p_t^{CN} = -\underset{(3.8)}{0.0567} p_{1,t}^{WDG} + \Delta g_t^{CPI} + \underset{(3.3)}{0.254} h_{WP,t}^{BS} + \underset{(3.4)}{0.0378} h_{LP,t}^{BS} - \underset{(2.1)}{0.00004} t \qquad (33)$$

is much smaller (p-value = 0.0082) than its standard values.

The estimates obtained from model (31) were much more promising. The

Table 3: CPI and B-S model (34) - cointegration tests and companion matrix roots.

V	Trace	Trace(B)	p-value	p-value (B)
0	111.9	99.25	0.000	0.003
1	60.16	53.49	0.074*	0.228*
2	29.71	21.36	0.481*	0.906
3	10.40	8.42	0.881	0.959
4	3.81	3.48	0.762	0.805

Roots	Real	Imaginary	Modulus
1	0.963	0.064	0.966*
2	0.963	-0.064	0.966*
3	0.943	0.000	0.943*
4	0.802	-0.027	0.803
5	0.802	0.027	0.803

Note: (B) indicates *Trace* test with Bartlett correction.

Table 4: CPI nad B-S model (34) – estimates and diagnostics.

	p_2^{WDG}	Δp^{CN}	g^{CPI}	h_{WP}^{BS}	h_{LP}^{BS}	t
β_1	0.0635 (4.3)	**1**	-0.0635 (4.3)	-0.262 (5.1)	-0.0407 (3.9)	0.00005 (2.3)
α_1	-0.283 (1.7)	**-0.568** (5.4)	–	0.593 (4.7)	–	–
LR = 0.616						
AR(1) = 0.084 AR(2) = 0.365			DH = 0.744			
AR(3) = 0.379 AR(4) = 0.326			ARCH(1) = 0.279 ARCH(2) = 0.386			

Note: t-ratios are reported in parentheses. Dots stand for the parameters with t-ratios smaller than 2. LR is a likelihood ratio test for over-identifying restrictions, DH – Doornik-Hansen normality test, AR(s) – autocorrelation LM test, ARCH(s) – ARCH effect test (for details see: Juselius (2006); p-values are reported for LR, AR, DH and ARCH tests.

cointegration tests confirmed that only one equilibrium condition was present, to which producer price inflation Δp^{CN} adjusts (table 3); as in model (29), there were enough arguments for removing the real exchange rate q^T from the cointegrating space. The number of the cointegrating vectors in model

$$y_t = \left[p_{2,t}^{WDG}, \Delta p_t^{CN}, g_t^{CPI}, h_{WP,t}^{BS}, h_{LP,t}^{BS}\right]. \tag{34}$$

did not change. There were also grounds for imposing the weak exogeneity restriction on the proxy of the Balassa-Samuelson effect h_{LP}^{BS} and on imported inflation g^{CPI}. In the last step, the long-run homogeneity restriction of the price wedge p_2^{WDG} and globalization effects g^{CPI} was accepted (see table 4).
Several conclusions can be drawn from the estimates.
Firstly, the VEC model (34) points to the existence of a double equilibrium correction mechanism. The $\Delta^2 p^{CN}$ adjustments run along the cointegrating vector:

$$\Delta p_t^{CN} = \underset{(4.3)}{-0.0635}(p_{2,t}^{WDG} - g_t^{CPI}) + \underset{(5.1)}{0.262}h_{WP,t}^{BS} + \underset{(3.9)}{0.0407}h_{LP,t}^{BS} - \underset{(2.3)}{0.00002}t \tag{35}$$

Processes equilibrating consumer prices along levels determined by the price wedge can also be identified. The internal equilibrium correction mechanism increases consumer price inflation because of producer prices and consumer import prices are growing. On the other hand, the excessive CPI value (compared with PPI and imported inflation tax) induces negative consumer price dynamics.
Secondly, from equation (35) transformed into:

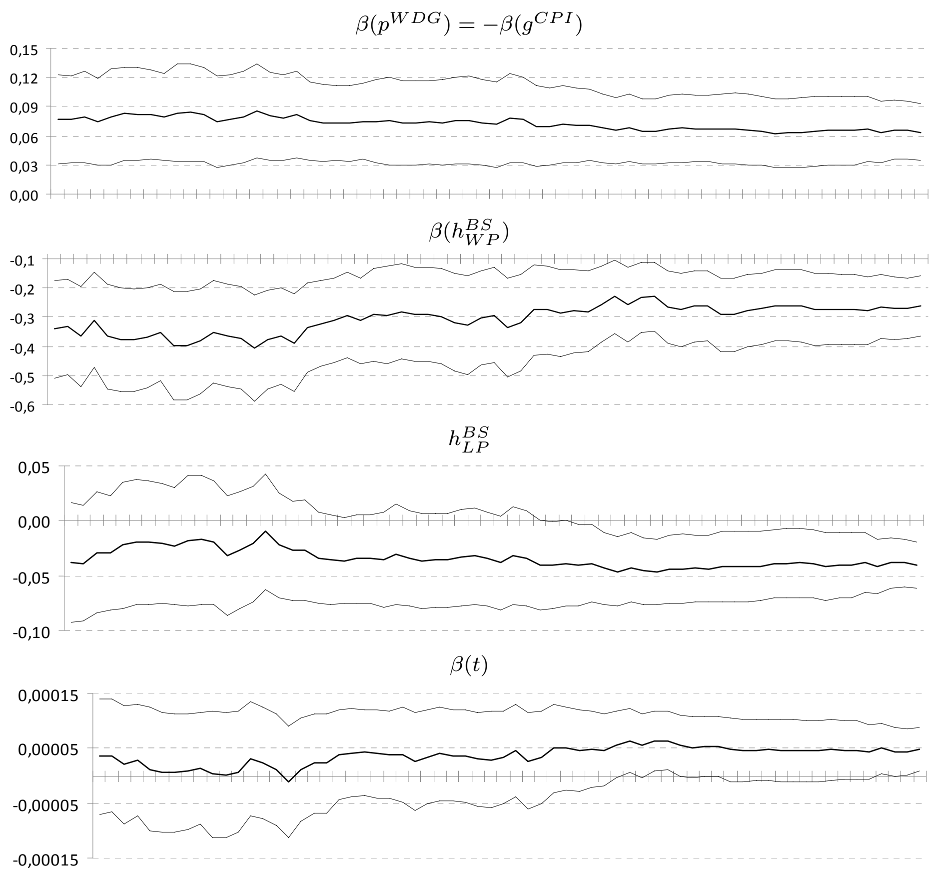
$$\Delta p_t^{CN} = \underset{(4.3)}{-0.0635}(p_{2,t}^{WDG} - g_t^{CPI} - \underset{(3.9)}{0.641}h_{LP,t}^{BS}) + \underset{(5.1)}{0.262}h_{WP,t}^{BS} + \underset{(2.3)}{-0.00002}t \tag{36}$$

it can be concluded that the Balassa-Samuelson mechanism has a stable and significant effect on the overall level of prices in Poland, as indicated by the theoretical models. The B-S effect transmitted via h_{LP}^{BS} is comparable in size with the share of non-tradable prices in the consumer basket.
Thirdly, for the above mechanisms to be identified in the VEC model (34), the mechanisms controlling for the wage equalization must be added. After h_{WP}^{BS} is added to the model, the new system has satisfactory stochastic properties (table 1), as well as stable estimates of its equilibrium parameters (see Figure 8).

4 The wage-price loop in the open sector

The last part of the analysis of the zloty exchange rate and prices in Poland is the determinants of tradables price. The discussion presented above indicates that at least three important problems should be solved. Firstly, it is necessary to explain the long-term disinflationary trend in model (18)-(19), which has been approximated so far by means of a deterministic trend. Secondly, the importance of imported inflation tax for producer price levels must be established. The conclusions derived from the analysis of the CPI model offer solid arguments in support of the statement that

Figure 8: Recursive estimation of the long-term parameters of the model (35) – consumer price inflation, 2004:07-2009:12.



globalization has weak direct effects on consumer prices, but formulating the same opinion for tradable prices could be premature. A perfunctory analysis is enough to find out that towards the end of the sample period the ratio between consumption imports and consumption u^C_M is almost three times smaller than the ratio between intermediate imports and global industrial production u^Z_M, the latter ratio growing by about 15 percentage points between 1999 and 2008 and then declining to about 35% in 2009 (see Figure 2). In this context, the basic question concerns the relationship between strong disinflationary trends in Poland and the growing significance of the import of low inflation. Finally, it must be remembered that the role of the unit labour costs (ULC) as the producer price determinants has been completely omitted from the discussion so far. This approach is fully justified when a strictly long-term perspective is adopted, in which case the assumption about unit labour costs fully

adjusting to the world prices is correct. However, when the time horizons are shorter, the question about ULC impacts needs to be answered.
The expectations-augmented Phillips curve is one of the usual starting points for price analysis (reviews in: Roberts 1995, Paloviita 2005):

$$\Delta p_t = \beta_1 \Delta p_t^e + \beta_2 G_t + \beta_3 D_t, \tag{37}$$

where: Δp, Δp^e – observed and expected price inflation, G – approximation of the short-run disequilibria, D – the variable representing the impact of exogenous shocks, β_1, β_2, β_3 – parameters.
To model producer prices using the above approach, the mechanism determining price expectations must be identified and the production or demand gap defined. In particular, one can think of a model where first price expectations are formulated following to the PPP equation (18) and then an attempt is made to add the relevant proxy of the short-run disequilibrium.
While the usefulness of the above solution for modelling producer prices cannot be rejected a priori, its simplifications should be borne in mind. For instance, model (37) implicitly assumes that all effects of changing labour productivity are fully transmitted onto wages. This problem can be illustrated with a wage Phillips curve extended to allow for inflationary expectations:

$$\Delta w_{P,t} = \gamma_1 \Delta p_t^e + \gamma_2 G_t + \gamma_3 \Delta l_{P,t} + \gamma_4 D_t, \tag{38}$$

where: Δw_P – wage inflation, Δl_P – labour productivity growth, γ_1, γ_2, γ_3, γ_4 – parameters.
Assuming that both Phillips curves are homogenous in the long term, i.e. $\beta_1 = \gamma_1 = 1$, and imposing a restriction requiring that the rate of wage increase be determined by the rate of inflation and by the rate of labour productivity:

$$\Delta p_t = \Delta w_{P,t} - \Delta l_{P,t} \tag{39}$$

it becomes possible to demonstrate that the resulting price Phillips curve embodies a relationship between the growth of inflation and of labour productivity:

$$\Delta p_t = \Delta p_t^e + \gamma_2 G_t - (1 - \gamma_3) \Delta l_{P,t} + \gamma_4 D_t. \tag{40}$$

The bottom line is that specification errors will not bias parameter estimates of model (37) only if $\gamma_3 = 1$, i.e. unless falling prices entirely offset the positive effects of productivity changes on wage increases.
The assumption about labour productivity growth being proportionally transmitted onto wages is not obvious when the time horizon is shorter, although not necessarily short. An interesting generalization of this issue can be found in the paper by Ball and Moffitt (2001), who indicate that the differences between labour productivity and wage aspirations are significant factors shaping wages (Akerloff and Yellen 1990). A

discussion on the Phillips curve for Poland leads to many other issues. Among the major ones, there is the criticism of the neo-Keynesian Phillips curve itself (Mankiw 2000), which has a number of properties that are hardly acceptable already at the theoretical research level, for instance (i) the possibility of predicting booms under disinflation, (ii) problems with generating growth paths with high rates of inflation, and (iii) system reactions to exogenous shocks that are inconsistent with theory-based expectations.

Because the possibility of constructing the reduced forms of the models using the expectations-augmented Phillips curve, is problematic, it is justified considering some alternative approaches that explicitly model the relationships linking prices, wages, labour productivities and production (or demand) gaps. The wage-price loop seems a good starting point for econometric analysis that utilizes a cost-push equation connecting nominal wages with producer prices and labour productivity and describes prices as a function of the unit labour costs and imports deflators.

The standard cost-push price formula:

$$p_t^T = \delta_1\left(w_{P,t}^T - l_{P,t}^T\right) + \delta_2 p_t^M + \left(1-\delta_1-\delta_2\right)k_t^O + \mu_1, \tag{41}$$

where: p^M – the import deflator, k^O – other costs, μ_1 – mark-up, δ_1, δ_2 – parameters, defines the conditions of a steady-state equilibrium. Price inflation is induced then by the supply-side or structural factors, such as the 'autonomous' changes in labour productivity and/or absorption of imports. Making allowance for the growing importance of the latter factor and assuming that the effects of the other non-wage costs are proportional we arrive at an alternative form of the cost-push price equation:

$$p_t^T = \left(1-u_{M,t}^Z\right)\left(w_{P,t}^T - l_{P,t}^T\right) + u_{M,t}^Z p_t^M + \mu_2 \tag{42}$$

where: u_M^Z – the ratio between intermediate imports and gross value added.

In the paper, the wage-price loop was analysed empirically using a VAR model containing producer prices p^T, nominal wages w_P^T and labour productivity in the open sector l_P^T, the PPI-based real exchange rate q^T and the weakly exogenous unemployment rate U:

$$y_t^E = \left[p_t^T, w_{P,t}^T, l_{P,t}^T, q_t^T\right], \quad y_t^x = \left[U_t\right]. \tag{43}$$

As in the previous models, the potential I(2)-ness of the nominal variables was taken into account (e.g. Kelm and Majsterek 2007). Assuming that prices p^T and unit labour costs $w_P^T - l_P^T$ have a common double unit root, the nominal-to-real transformation leads to the following model:

$$y_t^E = \left[\Delta p_t^T, \Delta w_{P,t}^T, p_t^{ULC}, g_t^{PPI}, q_t^T\right], \quad y_t^x = \left[U_t\right]. \tag{44}$$

where

$$p_t^{ULC} = p_t^T - \left(w_{P,t}^T - l_{P,t}^T\right) \tag{45}$$

is the producer price wedge. The globalization effect on PPI is given by:

$$g_t^{PPI} = p_t^M - u_{M,t}^Z \left(w_{P,t}^T - l_{P,t}^T \right). \tag{46}$$

Nominal wages, producer prices and labour productivity were tested for homogeneity using the VEC model (43) and assuming that 1, 2 and 3 cointegrating vectors were present. The conclusions were clear: because the probability values from the LR tests distinctly exceeded the standard significance levels (0.520, 0.384 and 0.238, respectively), the VEC model (44) could be used in further analyses.
Both the trace tests and the moduli of the characteristic roots suggest that two equilibrium conditions are present in the VEC model (44) (see table 5). The estimates of the error correction terms indicate that the first cointegrating vector is an attractor of wage inflation and the second cointegrating vector defines equilibrium conditions for producer-price inflation. With the above results, it is possible to structuralize the first cointegrating vector following the standard wage Phillips curve where the nominal wages are determined by prices, labour productivity and labour market disequilibrium approximated with the rate of unemployment. The second cointegrating vector is structuralized starting with equation (42) to account for a growing ratio between intermediate imports and industrial global production. The cointegrating vectors

Table 5: PPI-ULC model (44) - cointegration tests and companion matrix roots.

V	Trace	Trace(B)	p-value	p-value (B)
0	184.3	163.5	0.000	0.000
1	92.94	80.73	0.001	0.008
2	43.71	38.64	0.136*	0.311*
3	23.80	20.57	0.226	0.409
4	9.61	8.58	0.296	0.298

Roots	Real	Imaginary	Modulus
1	0.961	0.126	0.969*
2	0.961	-0.126	0.969*
3	0.813	0.000	0.813*
4	0.063	-0.713	0.716
5	0.063	0.713	0.716
6	0.639	-0.000	0.639

Note: (B) indicates *Trace* test with Bartlett correction.

normalized and structuralized, the following estimates of the long-run parameters were obtained (see table 6 for diagnostics):

$$\Delta w_{P,t}^T = \underset{(3.6)}{0.291} \Delta p_t^T + \underset{(3.7)}{0.0255} p_t^{ULC} - \underset{(7.0)}{0.0434} q_t^T - \underset{(7.6)}{0.00074} U_t - \underset{(5.1)}{0.00010} t, \tag{47}$$

$$\Delta p_t^T = -\underset{(4.1)}{0.0630} p_t^{ULC} + \underset{(3.9)}{0.174} g_t^{PPI}. \tag{48}$$

Table 6: PPI-ULC model (44) – estimates and diagnostics.

	Δp^T	Δw_P^T	p^{ULC}	g^{PPI}	q	U	t
β_1	-0.291 (3.6)	**1**	-0.0255 (3.7)	0	0.0434 (7.0)	0.00074 (7.6)	0.00010 (5.1)
β_2	**1**	0	0.0630 (4.1)	-0.174 (3.9)	0	0	0
α_1	–	**-1.307** (9.7)	1.684 (7.5)	0.498 (3.4)	–	–	–
α_2	**-0.735** (5.5)	-0.743 (6.3)	–	–	–	–	–
LR = 0.579							
AR(1) = 0.733 AR(2) = 0.036			DH = 0.295				
AR(3) = 0.035 AR(4) = 0.839			ARCH(1) = 0.033 ARCH(2) = 0.131				

Note: t-ratios are reported in parentheses. Dots stand for the parameters with t-ratios smaller than 2. P-values are reported for LR, AR, DH and ARCH tests.

Wage inflation positively responds to producer prices and labour productivity, but negatively to nominal wages' positive deviations from the path defined by the producer price wedge p^{ULC}. This mechanism can be interpreted in terms of internal equilibrium correction. The positive elasticity estimate at price inflation seems moderate, but the negative and very precise parameter estimates at the real exchange rate and unemployment rate indicate a trade-off between labour and the costs of the intermediate imports, on one hand, and the mechanisms predicted by the wage Phillips curve with a constant or very slowly changing non-accelerating wage unemployment rate (NAWRU), on the other. Besides, the parameter estimates show that internal equilibrium correction is present in the PPI equation too. The balancing mechanism decelerates PPI growth in manufacturing when its level is too high compared with the path defined by p^{ULC}. At the same time, a stable dependence of the domestic price inflation on growing globalization effects g^{PPI} is empirically confirmed.

There are two aspects suggesting that above results should be interpreted with caution. Firstly, model (42) indicates that the producer price wedge p^{ULC} and the globalisation effect g^{PPI} have equal parameters but this restriction cannot be imposed for the high accuracy of both the estimates. Secondly, wage equation (47) contains a trend suggesting that a steady mechanism reducing wage dynamics is present in the sample period, which calls for clarification.

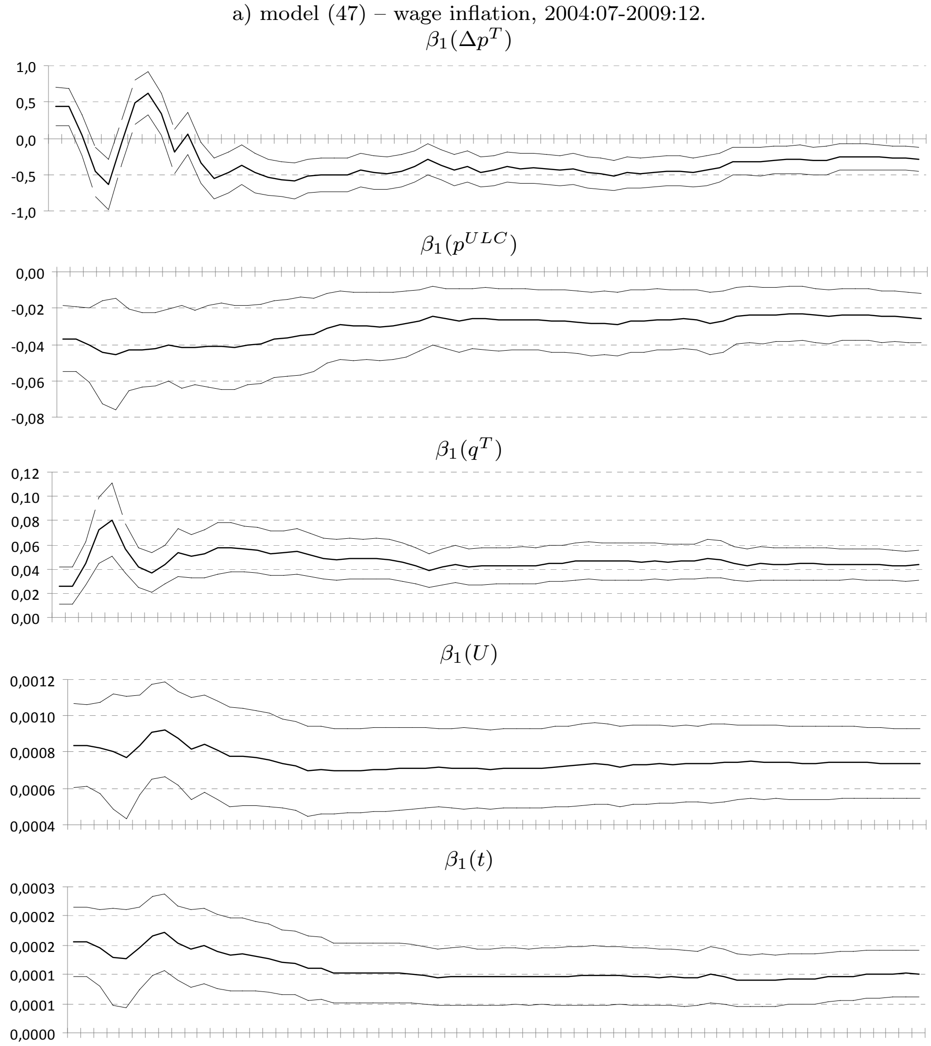
In order to solve the first problem, equation (48) has to be appropriately transformed. Having imposed a relatively non-controversial assumption about imports prices being determined by producer prices in the euro area and by the nominal zloty/euro exchange rate, i.e. $p^M = p^{*T} + b$, and having considered the average ratio between intermediate imports and industrial global production $\overline{u}_M^Z$, the price inflation equation

(48) takes the following form:

$$\Delta p_t^T = -\underset{(4.1)}{0.0630}\left(p_t^T - \left(w_{P,t}^T - l_{P,t}^T\right)\right) + \underset{(3.9)}{0.174}\overline{u}_M^Z \frac{u_{M,t}^Z}{\overline{u}_M^Z}\left(p_t^{*T} + b_t - \left(w_{P,t}^T - l_{P,t}^T\right)\right). \quad (49)$$

Figure 9: Recursive estimation of the long-term parameters

a) model (47) – wage inflation, 2004:07-2009:12.



Because in the analysed period $\overline{u}_M^Z$ was approximately 0.32 (hence $0.174\overline{u}_M^Z =$

= 0.057), model (49) can be simplified to the following equation:

$$\Delta p_t^T = -0.06 \left(p_t^T - \frac{u_{M,t}^Z}{\overline{u}_M^Z} \left(p_t^{*T} + b_t \right) - \left(1 - \frac{u_{M,t}^Z}{\overline{u}_M^Z} \right) \left(w_{P,t}^T - l_{P,t}^T \right) \right), \tag{50}$$

indicating that the effects of the internal equilibrium correction mechanism on producer price inflation result from the increasing openness of the tradables sector. The structure of equation (50) directly stems from price equation (42). The positive results of the cointegration analysis should be treated as an empirical proof that the thesis about non-linearity of the relationships linking producer prices, unit labour costs and import prices is correct. The recursive estimation results strongly support this conclusion (see Figure 9).

The stability of parameter estimates in wage equation (47) is generally acceptable, but explaining why a deterministic trend is present in wage equation (47) is somewhat problematic, so additional hypotheses have to be tested out. The investigation should primarily determine if the downward trend in the dynamics of the tradable sector's wages and the inflow of foreign direct investments (FDI) are interrelated. The working hypothesis might assume that FDI inflows change the utilization structure of particular production factors. In an economy saturated with modern technologies the demand for labour may be systematically declining, finally causing a downward pressure on wages.

Being only preliminary and requiring more in-depth studies, is clearly supported by the first estimates (table 7) generated by a VEC model with a FDI-to-GDP ratio (without a deterministic trend):

$$y_t^E = \left[\Delta p_t^T, \Delta w_{P,t}^T, p_t^{ULC}, g_t^{PPI}, q_t^T \right], \quad y_t^x = \left[U_t, FDI_t \right]. \tag{51}$$

b) model (51) – producer price inflation, 2004:07-2009:12.

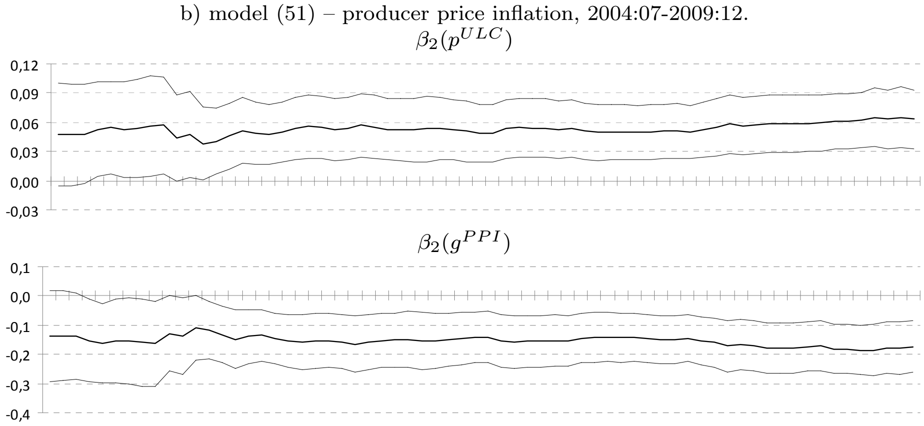


Table 7: PPI-ULC model (44) – estimates and diagnostics.

	Δp^T	Δw_P^T	p^{ULC}	g^{PPI}	q	U	FDI	c
β_1	-0.456 (5.8)	**1**	-0.0202 (2.9)	0	0.0438 (6.7)	0.00056 (6.8)	0.0457 (4.0)	-0.0219 (6.5)
β_2	**1**	0	0.0610 (4.3)	-0.167 (4.1)	0	0	–	-0.0035 (4.0)
α_1	–	**-1.255** (9.4)	1.612 (7.6)	0.424 (3.1)	-1.293 (2.4)	–	–	–
α_2	**-0.736** (5.8)	-0.822 (6.4)	–	–	-1.348 (2.6)	–	–	–
LR = 0.604								
AR(1) = 0.545 AR(2) = 0.036				DH = 0.393				
AR(3) = 0.015 AR(4) = 0.916				ARCH(1) = 0.329 ARCH(2) = 0.662				

Note: t-ratios are reported in parentheses. Dots stand for the parameters with t-ratios smaller than 2. P-values are reported for LR. AR. DH and ARCH tests.

5 Conclusions

The paper discusses estimation results obtained from three models:

(a) the model of tradables price inflation and of the PLN/EUR real exchange rate specified in accordance with the PPP hypothesis and the uncovered interest rate parity extended to include the risk premium (19):

$$q_t^T = -\underset{(3.6)}{5.111}\left(\left(i_t - \Delta p_t^T\right) - \left(i_t^* - \Delta p_t^{*T}\right)\right) + \underset{(5.1)}{0.144} U_t^{DS},$$

(b) the model of consumer price inflation extended to allow for the Balassa- Samuelson effect (36):

$$\begin{aligned} \Delta p_t^{CN} &= -\underset{(4.3)}{0.0635}\left(p_t^{CN} - p_t^T - g_t^{CPI} - \underset{(3.9)}{0.641}\left(l_{P,t}^T - l_{P,t}^{NT}\right)\right) + \\ &\qquad + \underset{(5.1)}{0.262}\left(w_{P,t}^{NT} - w_{P,t}^T\right) - \underset{(2.3)}{0.00002} t \end{aligned},$$

(c) the model of wage and price inflation in the tradables sector, whose structure follows the cost-push formula of producer prices and the wage Phillips curve (50):

$$\Delta p_t^T = -0.06\left(p_t^T - \frac{u_{M,t}^Z}{\bar{u}_M^Z}\left(p_t^{*T} + b_t\right) - \left(1 - \frac{u_{M,t}^Z}{\bar{u}_M^Z}\right)\left(w_{P,t}^T - l_{P,t}^T\right)\right).$$

The empirical analysis offers a range of conclusions.

The purchasing power parity model is not sufficient to fully describe the PLN/EUR exchange rate and tradables prices. In particular, the long-run deviations of the real exchange rate from the PPP path follow the fluctuations in the real interest rate differential and in an empirically selected currency risk.

If the tradables prices-based PPP hypothesis must be rejected, then one of the key assumptions underlying the Balassa-Samuelson model is rejected too. Moreover, the analysis of the determinants of consumer price inflation shows that an equally important assumption in the B-S model about the equilibration of wages in the tradable and non-tradable sectors is not satisfied, either. Therefore, one has to control for changes in the relative nominal wages, when intending to quantify the relationships between relative prices and relative productivities of labour. The empirical results point out that this approach to modelling the Polish CPI is correct, which raises serious doubts whether the Balassa-Samuelson model can be verified using its simplified reduced forms (20) and (21).

The dependence of producer prices on unit labour costs and import prices is non-linear - with the Polish economy becoming increasingly open the mechanisms determining the foreign demand gap gain in significance, while the unit labour costs lose their importance. This finding corresponds to the results obtained from the first equation of the CHEER model (see equation (18)), where the domestic price adjustments to the PPP path were non-linear too.

Summing up, the empirical analyses have pointed out that external price impulses are mainly transmitted via tradable prices, whose role increases with the economy becoming more and more open. There are two other mechanisms causing price fluctuations in Poland. One is the exchange rate following the changes in the foreign exchange risk. The other one transmits real shocks generated by FDI inflows and TFP changes.

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