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Modelling the Zloty–Euro Exchange Rate

1. Introduction

This paper continues the author's recent studies devoted to the problem of modelling the zloty–euro exchange rate[1]. The exchange rate is that type of macroeconomic category whose behaviour is often influenced by unexpected and significant breaks. Our past experiences *give the impression* that technical analysis used in that case seems to be a proverbial fortune telling from *dregs*. It happens that conclusions resulting from this analysis are extremely different and may wrongly suggest a feeling of helplessness in the possibility of finding the satisfactory recognition of studying some economic rules.

It seems that econometric modelling provides us with research tools, which allow us to manage this problem more efficiently. In particular, the application of dynamic models with the inclusion of integration and cointegration analysis may be very useful in such a case.

In macroeconomic studies we usually have at our disposal nonstationary series, often integrated of order one. Standard statistical tests used in univariate or multivariate analyses are inappropriate and results obtained, e.g. relating to the influence of explanatory variables, is false – such variables are not in any causal-outcome relation with a dependent variable and produce only a spurious dependency. Cointegration analysis applied in such cases protects us from building spurious regressions[2], and, in addition, creates the chance of finding a long-run relationship (LRR), i.e. relation in which included variables tend to the long-run equilibrium.

[1] Similar analysis based on the weekly data is done, for example, in Krauze (2003; 2004).

[2] See: Newbold and Davies (1978).

The next sections present various zloty–euro model specifications carried out in order to find the long-run tendency for this rate – and preliminary forecasts of the zloty–euro exchange rate. The main conclusions from our research end the paper.

2. Zloty–euro exchange rate models, results of estimation and testing, preliminary forecasts

The question arises whether we can formulate a univariate process which correctly and precisely enough represents the behaviour of the zloty–euro exchange rate or, if not, whether there exists any multivariate process in which variables are cointegrated. We are also interested if these solutions can be useful in the prediction process of the zloty–euro exchange rate. In empirical analysis we use the following monthly logarithmic series rates: Polish zloty–euro (pe_t), Swiss franc–euro (se_t), US dollar–euro (ue_t) and British pound–euro (be_t) from the period January 1999 – May 2003[3]. Graphs of pe_t , ue_t , be_t and se_t with inclusion of their trend functions (PET, UET, BET[4] and SET) are presented in Fig. 1–4, respectively.

Let us consider the following x_t relation testing for a unit root with structural shifts in mean and in trend:

$$x_t = \mu + \beta t + \delta D(T_b)_t + \theta DU_t + \gamma DT_t^* + (1+\rho)x_{t-1} + \sum_{k=1}^{K} c_k \Delta x_{t-k} + e_t, \quad (1)$$

where μ is intercept, β – trend coefficient, δ – impulse break effect, θ – change in the intercept, γ – change in the trend, $(1+\rho)$ – autoregressive parameter, c_k – k-th parameter of the augmentation term, e_t – disturbance term, and where:

[3] The data for pe_t, ue_t, be_t and se_t are obtained from: a) *Prices in the national economy in 2001. Information and statistical papers* (2002) and *Prices in the national economy in 2002. Information and statistical papers* (2003) - for the period 1999-2001, b) www.stat.gov.pl and as a result of own calculations based on the daily quotations of the Polish National Bank (NBP) (pe_t) and the Frankfurt Stock Exchange (ue_t, be_t, se_t) – for the further data. It was also considered (but unsuccessfully) inclusion of real consumer price index.

[4] Graph of BET includes two additional breaks in the deterministic trend of be_t, occurred in September 2000 and December 2002 in order to observe easier breaks in mean of this variable in December 2000. These endeavours in equation (1c) are not necessary as a consequence of inclusion variable be_{t-1}.

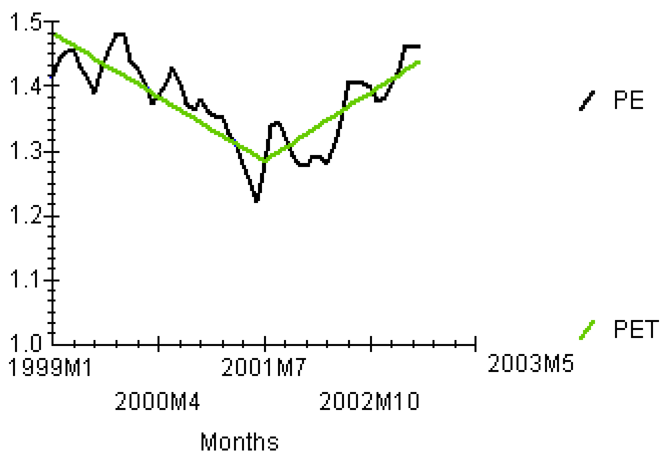

Fig. 1. Monthly series of PE and (pe_t) and PET

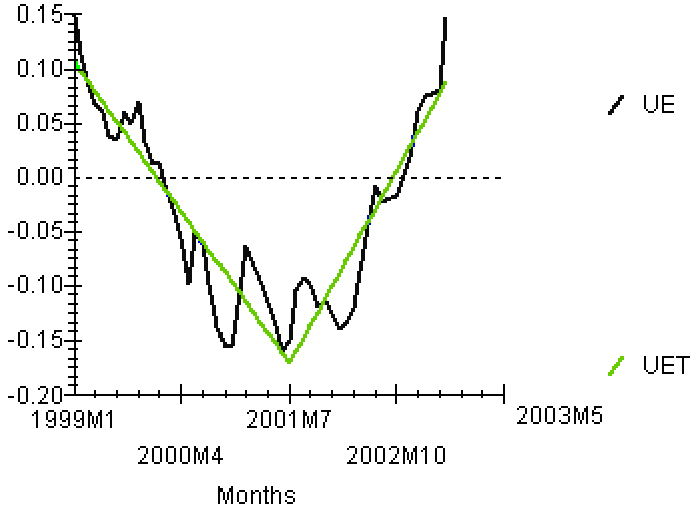

Fig. 2. Monthly series of UE (ue_t) and UET

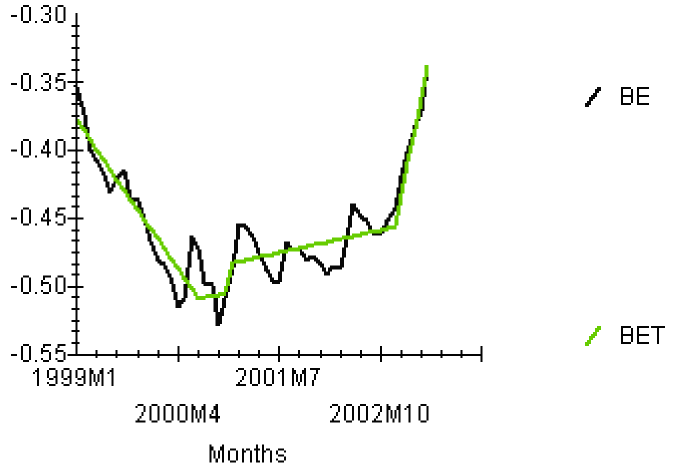

Fig. 3. Monthly series of BE (be_t) and BET

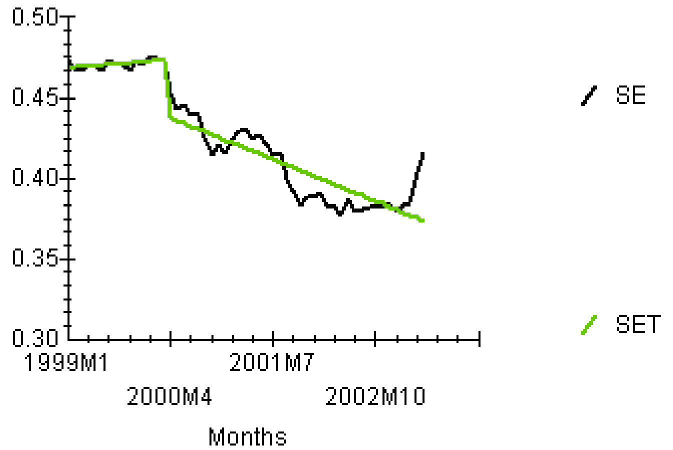

Fig. 4. Monthly series of SE (se_t) and SET

$$D(T_b)_t = \begin{cases} 1 \text{ dla } t = T_b + 1, \\ 0 \text{ dla } t \neq T_b + 1, \end{cases} \qquad DU_t = \begin{cases} 1 \text{ dla } t > T_b, \\ 0 \text{ dla } t \leq T_b, \end{cases} \qquad DT_t^* = \begin{cases} t - T_b \text{ dla } t > T_b, \\ 0 \quad\;\; \text{dla } t \leq T_b, \end{cases}$$

are appropriate dummy variables, while λ_b denotes the known timing break fraction $\lambda_b = T_b / T$, ($\lambda_b =$ 0.1, 0.2, 0.3, ..., 0.9; or an unknown timing break fraction ($0 < \lambda_b < 1$).

The following ordinary least squares (OLS) equation of pe_t is selected:

$$\hat{pe}_t = \underset{(4.35)}{0.542} - \underset{(-3.55)}{0.002}\, t + \underset{(3.82)}{0.005}\, DT_t^* + \underset{(-4.40)}{0.637}\, pe_{t-1} + \underset{(4.37)}{0.509}\, \Delta pe_{t-1}, \tag{1a}$$
$$R^2 = 0.916,\ DW = 1.87,\ h = 0.55\ [0.58],\ T_b = 31,$$

where *t*-values are in parentheses (at pe_{t-1} *t*-value for $\hat{\rho}$ is given – this convention is used below in all the unit root equations), R^2 is the determination coefficient, DW is the Durbin-Watson statistic, *h* is (normally distributed) the Durbin statistic, while in square parentheses the p-value is given (i.e. the significance level relating to *h*). An unknown timing (endogenous) break in the trend in July 2001 was estimated and growing tendency in the zloty–euro exchange rate appeared at a level of about 0.3 (=0.5–0.2)% per month. Zivot-Andrews' test critical value (*ZA*) for the case of an unknown timing break in trend, at 5% significance level[5] equals to –4.42 and it is slightly lower than the empirical *t*-value at pe_{t-1}, i.e., –4.40. Consequently, there is no reason to reject the null hypothesis of the unit root in pe_{t-1} (the alternative hypothesis that this variable is stationary with a shift in the trend could be accepted at a 10% significance level because in that case the critical value equals –4.11). The first differences of pe_t are stationary and, thus, this variable is integrated of order 1.

The above result motivates us to carry out a multivariate analysis. We should determine the orders of integration of variables, which can potentially be included in the cointegrating relationship (in which the linear combination of these variables is stationary). Consequently, the following testing equations of ue_t, be_t and se_t are selected:

$$\begin{aligned} \hat{ue}_t = {} & \underset{(1.23)}{0.018} - \underset{(-2.13)}{0.002}\, t + \underset{(3.02)}{0.059}\, D(T_{00.06})_t + \underset{(2.45)}{0.048}\, D(T_{00.12})_t \\ & + \underset{(2.76)}{0.052}\, D(T_b)_t + \underset{(2.96)}{0.006}\, DT_t^* + \underset{(-2.29)}{0.788}\, ue_{t-1} + \underset{(2.86)}{0.351}\, \Delta ue_{t-k}, \\ & R^2 = 0.957, \quad DW = 2.04, \quad h = -0.19\ [0.85], \quad T_b = 31, \end{aligned} \tag{1b}$$

[5] See: Zivot and Andrews (1992). The critical values of the ZA test used below come from this cited paper.

$$
\begin{aligned}
\hat{be}_t = &-\underset{(-1.25)}{0.022}+\underset{(5.27)}{0.001}t+\underset{(4.03)}{0.045}D(T_{00.06})_t-\underset{(-2.55)}{0.028}D(T_{00.08})_t-\underset{(-3.18)}{0.035}D(T_{00.10})_t \\
&+\underset{(3.09)}{0.034}D(T_{01.08})_t+\underset{(4.02)}{0.045}D(T_b)_t-\underset{(-3.22)}{0.021}DU_t+\underset{(-0.18)}{0.993}be_{t-1}, \\
&R^2=0.942, \quad DW=1.62, \quad h=1.42\ [0.15], \quad T_b=24,
\end{aligned}
\tag{1c}
$$

$$
\begin{aligned}
\hat{se}_t = &\underset{(1.25)}{0.007}+\underset{(1.25)}{0.0002}t-\underset{(-1.83)}{0.013}D(T_b)_t-\underset{(-2.38)}{0.008}DU_t+\underset{(-2.38)}{0.981}se_{t-1}, \\
&R^2=0.972, \quad DW=2.01, \quad h=-0.04\ [0.97], \quad T_b=15,
\end{aligned}
\tag{1d}
$$

where subscript b at additional impulse variables is substituted by the date of break (two last figures of year and month). Estimated in (1b)–(1d) trend coefficients are –0.2%, 0.1% and 0.2%. An unknown timing shift in the trend of ue_t (+0.6%), shifts in the mean of be_t (–2.1%) and se_t (–0.8%) appeared in July 2001, December and March 2000, respectively. A few impulse variables in (1b) and (1c) allow to estimate one-time effects of strong disturbances (of maximum order +5,2% and +4,5%, respectively). They have remarkable influence already at a 1% significance level. It may suggest the existence of exogenous breaks. At last we treat them in unit root testing as endogenous breaks strengthening our valuation criterion. Empirical *t*-values at variables ue_t and be_t, i.e. –2.29 and –0.18 are distinctly higher than the corresponding critical values of the *ZA* test (at 5% significance level, i.e. –4.42 and –4.80), and thus, all these variables are nonstationary. Their first differences are already stationary, and therefore, they are integrated of order 1. Furthermore, we re-estimate (using OLS) equations (1b)–(1d) including variable pe_{t-1}. This new regressor in each case appears not to be causal in Granger sense[6] at a 5% significance level.

The above results give a chance to find a cointegrating relationship which includes the considered above exogenous variables. Such a relation for variable y_t can be written as:

$$
y_t = \mu_0 + \beta_0 t + \delta_0 D(T_c)0_t + \theta_0 DU0_t + \gamma_0 DT0_t^* + x_t d + \xi_t, \tag{2}
$$

where μ_0, β_0, δ_0, θ_0, γ_0, d, ξ_t are: intercept, trend parameter, impulse break effect, shift in mean, shift in trend, vector of parameters of strongly exogenous variables (x_t) and disturbance term, while dummy variables are given as:

[6] It means, that determining some forecasts, based on these relations, we shouldn't use information relating to the past values of an endogenous variable; see: Granger (1969).

$$D(T_c)0_t = \begin{cases} 1 \text{ dla } t = T_c + 1, \\ 0 \text{ dla } t \neq T_c + 1, \end{cases} \quad DU0_t = \begin{cases} 1 \text{ dla } t > T_c, \\ 0 \text{ dla } t \leq T_c, \end{cases} \quad DT0_t^* = \begin{cases} t - T_c \text{ dla } t > T_c, \\ 0 \quad\ \text{ dla } t \leq T_c, \end{cases}$$

where $\lambda_c (= T_c / T)$ is known ($\lambda_c = 0{,}1,\ 0{,}2,\ 0{,}3,\ ...,\ 0{,}9$) or an unknown ($0 < \lambda_c < 1$) break fraction.

Using appropriate statistical criteria the following OLS distributed lag model (DL) is selected:

$$\begin{aligned} \hat{pe}_t = {} & \underset{(12.51)}{1.066} - \underset{(-4.22)}{0.001}t + \underset{(2.52)}{0.056}D(T_{99.10})0_t + \underset{(3.54)}{0.080}D(T_{00.10})0_t + \underset{(1.93)}{0.045}D(T_{00.11})0_t \\ & - \underset{(-2.34)}{0.053}D(T_{01.06})0_t + \underset{(1.85)}{0.041}D(T_{01.09})0_t + \underset{(1.93)}{0.043}D(T_{02.07})0_t \\ & + \underset{(2.31)}{0.052}D(T_{02.09})0_t + \underset{(2.73)}{0.013}DT0_t^* + \underset{(11.49)}{0.928}\hat{ue}_t - \underset{(-4.27)}{0.805}\hat{be}_{t-1}, \\ & R^2 = 0.918, \quad DW = 1.20, \quad T_b = 48, \end{aligned} \tag{2a}$$

where fitted values of ue_t and be_t are obtained in estimation of (1b) and (1c), treated as marginal equations. We find that the estimate of endogenous shift of +1.3% in trend of pe_t occurred in December 2002 (at a decreasing trend of order 0.1% monthly, to this break point). Now, this shift is stronger and occurs later than in equation (1a). The re-estimated (OLS) version of equation (2a) including residual regressors from (1b)–(1d) is as follows:

$$\begin{aligned} \hat{pe}_t = {} & \underset{(12.05)}{1.024} - \underset{(-4.80)}{0.001}t + \underset{(2.19)}{0.051}D(T_{99.10})0_t + \underset{(3.45)}{0.075}D(T_{00.10})0_t \\ & + \underset{(1.84)}{0.041}D(T_{00.11})0_t - \underset{(-2.57)}{0.056}D(T_{01.06})0_t + \underset{(2.06)}{0.044}D(T_{01.09})0_t \\ & + \underset{(2.71)}{0.062}D(T_{02.07})0_t + \underset{(2.20)}{0.048}D(T_{02.09})0_t + \underset{(3.28)}{0.016}DT0_t^* \\ & + \underset{(11.73)}{0.949}\hat{ue}_t - \underset{(-4.79)}{0.904}\hat{be}_{t-1} - \underset{(-5.16)}{1.015}(ue_t - \hat{ue}_t) + \underset{(0.35)}{0.126}(be_{t-1} - \hat{be}_{t-1}), \\ & R^2 = 0.929, \quad DW = 1.26, \quad T_b = 48, \end{aligned} \tag{2a'}$$

where regressor $(ue_t - \hat{ue}_t)$ has significant influence on pe_t ($t = -5.16$), therefore, ue_t is not weakly exogenous, and thus not strongly exogenous. It inclines us to undertake further attempts to find a cointegrating relationship.

Then, we will check whether it might be achieved by using the following (OLS) autoregressive distributed lag model (ADL) of pe_t:

$$
\begin{aligned}
\hat{pe}_t = {} & \underset{(2.73)}{0.279} - \underset{(-1.50)}{0.0003}\,t + \underset{(2.36)}{0.035}\,D(T_{99.08})0_t - \underset{(-2.03)}{0.029}\,D(T_{00.03})0_t \\
& + \underset{(3.12)}{0.046}\,D(T_{00.10})0_t + \underset{(3.93)}{0.060}\,D(T_{01.07})0_t + \underset{(2.67)}{0.039}\,D(T_{02.07})0_t \qquad \text{(2b)} \\
& + \underset{(1.79)}{0.005}\,DT0_t^* + \underset{(9.68)}{1.046}\,pe_{t-1} + \underset{(-4.95)}{0.414}\,pe_{t-2} + \underset{(5.87)}{0.446}\,\hat{ue}_t - \underset{(-4.76)}{0.541}\,\hat{be}_{t-1}, \\
& R^2 = 0.966, \quad DW = 1.24, \quad T_b = 48,
\end{aligned}
$$

where the date of endogenous shifts in trend is the same as in (2a). Now, estimates of the trend coefficient (0,03%) and shift in the trend (0.5%,) are corrected due to the inclusion of pe_{t-1} and pe_{t-2}. Equation (2b) is re-estimated after adding residuals of ue_t and be_t from (1b) and (1c) as regressors. Estimation results are as follows:

$$
\begin{aligned}
\hat{pe}_t = {} & \underset{(3,25)}{0,313} - \underset{(-2,26)}{0,0005}\,t + \underset{(2,59)}{0,038}\,D(T_{99.08})0_t - \underset{(-2,35)}{0,031}\,D(T_{00.03})0_t + \underset{(3,58)}{0,049}\,D(T_{00.10})0_t \\
& + \underset{(3,64)}{0,052}\,D(T_{01.07})0_t + \underset{(3,21)}{0,046}\,D(T_{02.07})0_t + \underset{(2,26)}{0,007}\,DT0_t^* + \underset{(10,30)}{1,036}\,pe_{t-1} - \underset{(-5,97)}{0,469} \\
& \cdot pe_{t-2} + \underset{(6,98)}{0,525}\,ue_t - \underset{(-5,91)}{0,676}\,be_{t-1} - \underset{(-1,96)}{0,266}(ue_t - \hat{ue}_t) + \underset{(0,10)}{0,023}(be_{t-1} - \hat{be}_{t-1}), \\
& R^2 = 0,973, \quad DW = 2,04, \quad T_b = 48, \qquad \text{(2b')}
\end{aligned}
$$

Both new regressors are insignificant at a 5% level. Therefore, variables ue_t and be_t are weakly exogenous. It suggests considering a single equation model, i.e. with the exclusion of marginal equations. Consequently, the following OLS equation pe_t is selected:

$$
\begin{aligned}
\hat{pe}_t = {} & \underset{(3,55)}{0,457} - \underset{(-2,44)}{0,001}\,t + \underset{(2,41)}{0,035}\,D(T_{99.08})0_t + \underset{(2,78)}{0,040}\,D(T_{00.10})0_t \\
& + \underset{(3,57)}{0,052}\,D(T_{01.07})0_t + \underset{(2,94)}{0,041}\,D(T_{02.07})0_t + \underset{(2,54)}{0,009}\,DT0_t^* + \underset{(11,55)}{1,103}\,pe_{t-1} \qquad \text{(2c)} \\
& + \underset{(-5,80)}{0,468}\,pe_{t-2} + \underset{(6,73)}{0,485}\,ue_t - \underset{(-5,44)}{0,633}\,be_{t-1} - \underset{(-2,00)}{0,456}\,se_t, \\
& R^2 = 0,970, \quad DW = 2,03, \quad T_b = 48,
\end{aligned}
$$

where the date of the endogenous shift in the trend is the same as in (2a) and (2b). The estimates of this shift (+0.9%) and trend coefficient (monthly decries 0.1%) are different. ADL models in (2b) and (2c) have, e.g., distinctly better fitting than the DL model in (2a). It should be noted that due to the fact that variables ue_t, be_t and se_t are weakly exogenous, and pe_{t-1} is not their cause, these three variables are strongly exogenous. As a result, the suggested OLS long-run relationships (LRR) based on the ADL models is the following:

$$\begin{aligned} \hat{pe}_t &= 0.758 - 0.001t + 0.096D(T_{99.08})0_t - 0.080D(T_{00.03})0_t \\ &+ 0.126D(T_{00.10})0_t + 0.162D(T_{01.07})0_t + 0.106D(T_{02.07})0_t \\ &+ 0.015DT0_t^* + 1.213\hat{ue}_t - 1.471\hat{be}_t, \end{aligned} \tag{2*}$$

$$\begin{aligned} \hat{pe}_t &= 1.250 - 0.004t + 0.095D(T_{99.08})0_t + 0.110D(T_{00.10})0_t \\ &+ 0.142D(T_{01.07})0_t + 0.112D(T_{02.07})0_t + 0.025DT0_t^* \\ &+ 1.327ue_t - 1.248be_t - 1.731se_t. \end{aligned} \tag{2**}$$

We consider using single equation tests for cointegration. They include residual-based tests and error correction model (*ECM*) tests. The tests from the first group were discussed by Engle and Granger (1987). The latter ones are described in Kremers, Ericsson and Dolado (1992). Below we apply extensions of both groups of testing procedures proposed by Krauze (2002). The residual-based version of cointegration test postulates using the OLS residuals ($\hat{\xi}_t = pe_t - \hat{pe}_t$, where $\hat{pe}_t$ are fitted values from LRR) in the equation of the Dickey-Fuller type (*CDF*), i.e.:

$$\Delta\hat{\xi}_t = b\hat{\xi}_{t-1} + \zeta_t, \tag{3a}$$

where b ($-2 \le b \le 0$) is the autoregressive parameter, and ζ_t is the disturbance term. Next, the *ECM* type of the test uses the following testing equation:

$$\Delta pe_t = a\Delta\hat{pe}_t + b\hat{\xi}_{t-1} + \varepsilon_t, \tag{3b}$$

where b is the error correction coefficient, a is the short-run parameter ($0 \le a \le 1$), ε_t is the disturbance term, while ζ_t and ε_t are independent from ξ_t. It is assumed that variables in a cointegrating relationship are integrated of order 1. We test $b = 0$ (variables are not cointegrated) against the alternative $-2 < b < 0$ (variables are cointegrated). Estimated (OLS) equations (3a) and (3b), corresponding to (2*) and (2**), are given as follows:

$$\Delta\hat{\xi}_t = \underset{(-5.79)}{-0.797}\hat{\xi}_{t-1} + \hat{\zeta}_t, \quad DW = 1.95, \tag{3a*}$$

$$\Delta pe_t = \underset{(7.71)}{0.338}\Delta\,\hat{pe}_t - \underset{(-7.79)}{0.481}\hat{\xi}_{t-1} + \hat{\varepsilon}_t, \quad DW = 1.20, \tag{3b*}$$

$$\Delta\hat{\xi}_t = \underset{(-5.90)}{-0.820}\hat{\xi}_{t-1} + \hat{\zeta}_t, \quad DW = 1.95, \tag{3a**}$$

$$\Delta pe_t = \underset{(7.17)}{0.311}\Delta \hat{pe}_t - \underset{(-7.69)}{0.468}\hat{\hat{\xi}}_{t-1} + \hat{\varepsilon}_t\,, \quad DW = 1.04. \qquad (3b^{**})$$

Critical values of *CDF* and *ECM* tests determined in the case of an endogenous shift in the mean in two ($N = 2$) marginal equations and an endogenous shift in the trend in a conditional equation, at 5% (*CDF*) and 1% (*ECM*) significance level and for $T = 51$, are equal to: –5.47 (=–4.93–27.75/51) and –6.22 (=–5.31–46.61/51). If in a marginal process we have a shift in the trend instead of a shift in the mean, such critical values are equal to: –5.43 (=–4.94–25.00/51) and –5.98 (=–5.42–28.71/51).[7] Therefore, the alternative hypothesis, assuming that variables pe_t, ue_t and be_t are cointegrated, should be accepted basing on the *CDF* ($t = –5.79$) and on the *ECM* ($t = –7.79$) tests, at 5% and 1% significance level, respectively, in case of a shift in the mean as well as a shift in the trend in the marginal process.

Geogory and Hansen (1996a, b) considered testing for cointegration in some special cases of shifts in single-equation models. The asymptotic critical value of their (*CDF* type) test of case shift in mean[8] in the cointegrating relationship, including three exogenous variables, at 5% significance level equals to –5.57. Using this criterion[9] on the base *t*-value (=–5.90) in (3a**) the hypothesis assuming cointegration of pe_t, ue_t, be_t and se_t should be accepted. It seems that critical values of Krauze's (*CDF* and *ECM* types) tests determined for the finite sample size may be an interesting reference-point for the analysed above single-equation model. These values are distinctly lower than the corresponding asymptotic critical values of (*CDF* type) test proposed by Geogory and Hansen. Consequently, they strengthen the criterion of rejecting the null hypothesis, assuming that considered variables are not cointegrated. Critical values of Krauze's *ECM* and *CDF* tests in the case of an endogenous shift in the mean, for ($N =$) 3 marginal equations and endogenous shift in trend in the conditional equation, at 5% (*CDF*) and 1% (*ECM*) significance levels, for $T = 52$ are equal to: –5.86 (=–5.20–34.49/52) and –6.55 (=–5.56–51.37/52), and in the case of a shift in the trend in the marginal process are equal to: –5,81 (=–5.23–30–20/52) and –6.39 (=–5.62–39.93/52)[10]. Corresponding empirical values of the *CDF* and *ECM* statistics given in (3a**) and (3b**), i.e. –5.90 and –7.69, are lower, and thus, variables pe_t, ue_t, be_t, se_t are cointegrated.

Basing on equations (1b), (1c) and (1d), forecasts of ue_t, be_t and se_t for June 2003 were determined. Next, they were used in calculation of pe_t forecasts basing on equations (2b) and (2c), i.e. 1.4929 and 1.4753 or taking anti-

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[7] See: Krauze (2002), 150–151 and 156–157.

[8] Geogory and Hansen did not consider the case of shift in trend analysed in this paper.

[9] From the author's studies results that this criterion is more rigorous than in case of shift in trend in a cointegrating relationship.

[10] See: Krauze (2002), 150–151 and 156–157.

logarithm 4.4500 and 4.3723 (zl/euro), respectively. The real value of this ratio in June was 4.4373, therefore, its overestimation and underestimation reached 0.29% and 1.46%, respectively. Thus, the first forecast, based on equation (2b), is distinctly more precise. However, we should remember that one-period forecasting horizon does not allow for any generalization.

3. Conclusions

Shifts in trend of the monthly series of Polish zloty–euro, US dollar–euro, British pound–euro and Swiss franc–euro rates in July and again July 2001, December and March 2002, respectively, were detected. All these series were nonstationary and integrated of order one. Basing on the ADL model we determined the long-run relationship. We proved that prices of the US dollar, British pound and Swiss franc in euro were weakly and strongly exogenous, cointegrated with the Polish zloty–euro rate. Two long-run relationship specifications were proposed, i.e. conditional model (including marginal process) and single-equation model. Basing on these models ex-post forecasts of zloty–euro rate were calculated. Conditional model was more useful. However, a sufficiently longer period of empirical verification of forecasts is necessary for formulating some generalization.

Further studies aiming at more precise recognition of regularities characterizing behaviour of exogenous variables are advisable. It has key meaning for achieving success in the prediction process of the zloty–euro exchange rate. Some experiences from analysis of ex post predictive information would be also very useful.

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