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THE INFLATION IMPACT OF SELECTED EUROPEAN UNION MEMBERS ON POLISH INFLATION

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Abstract

The article aims at determining the inflation influence between Poland and selected EU member states. Although for some time the general inflation level in those countries was definitely controllable, the problem seems to be returning. That is why in this article, using the model of Vector AutoRegression (VAR) and Granger causality test, we are attempting to determine inflation influences on Poland.
The study confirmed the impact of the selected countries on Polish inflation, expressed the general HICP index. However, in the case of Germany, the method has not proved the existence of such interactions. For this reason, it is made an attempt to clarify the reasons for non-compliance findings with data showing Germany as a Polish main trading partner for more than two decades. The authors try to show that lack of influence can be seen in the excessive generality of the main HICP index and predict that the chosen method confirm the effect of foreign trade indices in the HICP.

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Introduction

The aim of this article is to attempt at determining inflation influence between Poland and selected European Union member states. Although for some time the general inflation level was definitely controllable, the problem seems to be returning again. Therefore, this article, using the vector autoregression model (VAR) and Granger causality test, we will try to determine inflation influences on Poland.
We also try to explain the reasons for possible discrepancies between examination results and adopted assumptions concerning expected effects of calculations, resulting from the influence exerted through other channel of price transmission than inflation.
The most fundamental question posed by the authors is whether selected countries influence inflation in Poland or not.
To verify this question we will use the VAR models, which were created in answer to criticism, as an expression of doubts concerning structural modeling, resulting from the lack of precise theoretical foundations to indicate interdependence of the processes, which led to varied specifications obtained as a result of observing the principle of identification of equations.
Theoretical constraints imposed on multi-equation structural models were treated as 'unreliable' in a ground-breaking study by Sims (1980), in which the basis of new methodology of

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modeling vector auto-regression was established. The main differences between the methodology used by Sims and the methodology of structural approach adopted by Cowles Commission are presented below:

1. there is no *a priori* division into endogenous and exogenous variables,
2. no zero conditions are imposed,
3. there is no precise (prior to modeling) economic theory which would constitute the basis of the model (Charemza and Deadman, 1992).

The choice of variables for the model, general shape of the model and determination of initial maximum delays

The aim of the analysis is to identify causal relations between inflations in presented countries in the period of time under examination. To do so, we will apply a tool in form of vector autoregression.

To build the VAR model, we selected 4 variables *X, Y, Z, W*:

X_t – monthly inflation for particular goods in Poland in *t* period of time

Y_t – monthly inflation for particular goods in Slovakia in a *t* period of time

Z_t – monthly inflation for particular goods in Great Britain in a *t* period of time

W_t – monthly inflation for particular goods in Germany in a *t* period of time.

Our sample covers 81 observations made from January 2004 to October 2010 for each presented time series. Due to the fact that data is presented in a three-month configuration, we assume 12 as maximum delay. Obviously, in the further part of the paper, using an appropriate test, we will verify the hypothesis concerning the maximum series of delays.

The basic form of the VAR model for 4 variables and 12 delays can be expressed by the formula below:

$$\begin{aligned}
x_t &= \alpha_{1,0} + \sum_{j=1}^{12} \alpha_{1,j} x_{t-j} + \sum_{j=1}^{12} \beta_{1,j} y_{t-j} + \sum_{j=1}^{12} \gamma_{1,j} z_{t-j} + \sum_{j=1}^{12} \delta_{1,j} w_{t-j} + \varepsilon_{1,t} \\
y_t &= \alpha_{2,0} + \sum_{j=1}^{12} \alpha_{2,j} x_{t-j} + \sum_{j=1}^{12} \beta_{2,j} y_{t-j} + \sum_{j=1}^{12} \gamma_{2,j} z_{t-j} + \sum_{j=1}^{12} \delta_{2,j} w_{t-j} + \varepsilon_{2,t} \\
z_t &= \alpha_{3,0} + \sum_{j=1}^{12} \alpha_{3,j} x_{t-j} + \sum_{j=1}^{12} \beta_{3,j} y_{t-j} + \sum_{j=1}^{12} \gamma_{3,j} z_{t-j} + \sum_{j=1}^{12} \delta_{3,j} w_{t-j} + \varepsilon_{3,t} \\
w_t &= \alpha_{4,0} + \sum_{j=1}^{12} \alpha_{4,j} x_{t-j} + \sum_{j=1}^{12} \beta_{4,j} y_{t-j} + \sum_{j=1}^{12} \gamma_{4,j} z_{t-j} + \sum_{j=1}^{12} \delta_{4,j} w_{t-j} + \varepsilon_{4,t}
\end{aligned} \tag{1}$$

where: x_t, y_t, z_t, w_t – observations on endogenous variables of the model for periods $t = 1, 2, \ldots, n$

$x_{t-j}, y_{t-j}, z_{t-j}, w_{t-j}$ – observations on endogenous variables delayed in time by period *j*.

$\alpha_{k,j}, \beta_{k,j}, \gamma_{kj}, \delta_{kj}$ – structural parameters of the model accompanying relevant delayed variables in particular equations $k = 1, 2, 3, 4$. For example parameter $\alpha_{2,7}$ is a parameter accompanying variable *x* in the second equation, which is delayed by 7 periods.

$\alpha_{k,0}$ – absolute term in consecutive equations $k = 1, 2, 3, 4$

$\varepsilon_{k,t}$ – random element in period *t* in consecutive equations $k = 1, 2, 3, 4$.

Random elements from particular equations have independent normal distribution with zero expected value and variance Σ_{ε}. Moreover, there is a possibility to tie random elements between particular equations of the model, which means that simultaneous co-variances may be different from zero: $\operatorname{cov}(\varepsilon_{kt}, \varepsilon_{lt}) \neq 0$. The model parameters (1) may be estimated by means of a commonly known Least Squares Method, applying it to every equation separately. LSM estimators are consistent and asymptotically unbiased.

Examining stationariness of time series

The modeling of economic phenomena with VAR models consists of several stages.
One of the basic stages in constructing the VAR model is the examination of stationariness of processes generating variables included in the model. Stationariness of time series is a special state of 'their static balance' (Box and Jenkins, 1970). Lectures on stationariness begin with an introduction on stochastic processes, by which we understand a set of random variables ordered in line with indexes belonging to a certain set *T* (definite or indefinite). Obviously, for the purpose of this article we will focus our attention on the statement that a particular realization of a stochastic process are time series. Due to limited space of our paper, we omit the definition of stationariness of time series. More information on this can be found in the work of Box and Jenkins (1970). In the context of VAR models, the most important thing is the methods of discovering non-stationariness in time series and the ways of eliminating them. Initial identification may be done on the basis of a glance evaluation of time series graph, but then these results need confirmation of one of formal or informal methods. Among formal methods, the most popular ones are Dickey-Fuller test, Perron test and LM statistic-based test. The most important informal methods are the analysis of auto-correlation function and spectral density function, more on the former can be found in the work of Piłatowska (1997).
Because in the analyzed model we will check stationariness of presented time series using the Dickey-Fuller test only, it will be described in detail below.
The Dickey-Fuller test, in short the DF test, was first presented in 1979. It is also known as the unit root test. It verifies the hypothesis on the existence of a unit root, that is the hypothesis that $\theta = 1$ in equation:

$$x_t = \theta x_{t-1} + \varsigma_t, \tag{2}$$

where ζ_t is white noise process with expected value assumed to be zero, fixed variance and zero co-variance between observations. Therefore it is a stationary process (Box and Jenkins, 1970). In practice, we test the parameter with delayed variable basing on the equation:

$$\Delta x_t = \rho x_{t-1} + \varsigma_t \tag{3}$$

Which may also adopt the following form:

$$x_t = (1+\rho) x_{t-1} + \varsigma_t, \tag{4}$$

where $\theta = (1+\rho)$. If in formula (3) parameter ρ is negative, this is because in formula (2) parameter θ is lower than one and series x_t is stationary. In order to prove it, we use the Least Squares Method to estimate the parameter of equation (3). The null hypothesis: H_0: $\rho = 0$ assumes the existence of a unit root and confirms the non-stationariness of the analyzed series, whereas the alternative form: H_1: $\rho < 1$, allows us to confirm that the series generating variable x_t is stationary. The statistics which help us verify the above hypotheses has the following form:

$$DF = \frac{\hat{\rho}}{S(\hat{\rho})} \tag{5}$$

where: $\hat{\rho}$ - estimator of parameter ρ, $S(\hat{\rho})$ - standard error in parameter evaluation measured with standard deviation. At a first glance DF statistics is identical to t-Student test statistics for significance of the model parameters. Unfortunately, it has a totally different distribution. If the calculated value of DF test statistics is lower than critical value from relevant tables for a given size of sample n, we must make a decision to reject the null hypothesis on the existence of a unit root and accept the alternative hypothesis which claims stationariness of series x_t.

The tables with critical values for this test can be found in (2). Most computer packages which allow us to make calculations for VAR models automatically give critical values for this test. In this paper, we will analyze stationariness of examined series and many other calculations using the GRETL program. The results of DF test for 4 analyzed time series (processes) at work are included in Table 1.

Table 1: Results of DF test for series X_t, Y_t, Z_t i W_t

Series (process)	DF statistics value	*p - value*	Value of estimator ρ
X_t (POL)	- 2.439	0.131	-0.067
Y_t (SLO)	-2.286	0.176	-0.057
Z_t (UK)	-2.528	0.109	-0.099
W_t (GER)	-1.852	0.353	-0.083

Source: Own elaboration

On the basis of DF test results for analyzed series we can conclude that there are no grounds for rejecting H_0 about the existence of a unit root, which means that all series are non-stationary, thus they are not integrated series of zero grade I (0). In order to remove non-stationariness, we transform variables into their increases. Such transformed data are subjected again to DF test, the results of which are presented in Table 2.

Table 2: Results of DF test for first differences of series X_t, Y_t, Z_t and W_t

Series (process)	DF statistics value	*p - value*	Value of estimator ρ
X_t - X_{t-1} (POL)	- 4.83	0.00014	-0.483
Y_t - Y_{t-1} (SLO)	-6.826	0.0000002	-0.753
Z_t - Z_{t-1} (UK)	-6.313	0.0000008	-0.681
W_t - W_{t-1} (GER)	-4.957	0.000025	-0.832

Source: Own elaboration

After conducting the DF test for transformed data it turns out that in each of the analyzed cases we reject H_0 about the existence of a unit root and accept H_1, which says that the series of first differences are stationary series. This means that the analyzed series are integrated series of first grade I(1) and for further analysis we will use first differences of series due to their stationariness.

Delay length tests

Another essential stage in the vector autoregression analysis is to determine the maximum length of delays k for variables on the right side of the model equation (1). Due to the fact that in each of the VAR model equations, on the right we have all variables appearing on the left side and in other equations in a number equal to the maximum length of delays k, this causes that the number of these variables is significant. In the analyzed model (1) consisting of 4 equations we intuitively assumed the maximum number of delays $k = 12$ due to the fact that the analyzed data is analyzed month by month. Such number of equations and adopted value k cause that on the right side of each equation from model (1) there are 12*4 = 48 regressors plus absolute term. If the sample is not sufficiently numerous the estimation of such a model may be impossible due to too few degrees of freedom. The value of maximum length of delays k must be very rigorously determined. If the value k is too small, we often see auto-correlation of random elements of particular equations, and in a situation when we adopt too high k, we can observe the unacceptable limitation of degrees of freedom.

In order to determine the optimum length of delay, we use the likelihood-ratio test (LR). The procedure of establishing the optimum value k begins with establishing the longest probable delay (for example for monthly data this could be the multiple of 12) or the biggest possible value of delay lengths due to the number of degrees of freedom (the so-called sample size). In the next step of the procedure we limit the model with r delays and estimate its parameters for delays $k - r$. The examination of the length of delays consists in verification of hypothesis H_0, that the length of delays in the model equals $k - r$ against the alternative hypothesis H_1, stating that it amounts to k. Test statistics for the LR test can be expressed by the following formula:

$$LR = T\left(\log\left|\Sigma_{k-r}\right| - \log\left|\Sigma_k\right|\right), \qquad (6)$$

where: Σ_k, Σ_{k-r} – estimators of the variance co-variance matrix of the remainders of the models with the number of delays respectively k and $k - r$.

T – number of available observations.

The LR statistics has asymptotic distribution of chi-square with a number of degrees of freedom equal to the number of limitations imposed in the whole model. The values of test statistics bigger than critical value allow us to reject the null hypothesis, which says that the length of delay is $k - r$ and accept the alternative hypothesis, which assumes the model with a k number of delays. In a situation when the calculated value of test statistics is lower than critical value of distribution read from the table, there are no grounds for rejecting the H_0. In this case we should next test the hypothesis that the appropriate model is the one with an even smaller number of delays that $k - r$ (Enders, 1995). The second criterion which helps to determine the maximum length of delays in the VAR model is the AIC criterion (Akaike Information Criterion). Due to spatial constraints of this article, this criterion will not be analyzed here. Its detailed description can be found in the works of M. Pesaran and H. B. Pesaran (1997).

In order to determine the maximum number of delays for the model (1) we carried out LR and AIC tests for the model (1). The results obtained in the GRETL program are presented in Table 3.

Table 3: The results of the LR and AIC tests for the model (1)

Number of delays (k)	LR statistics	*p–value LR*	AIC criterion
1	-92.98407		3.27490
2	-60.31332	0.00000	2.791690
3	-46.23049	0.03021	2.847261
4	-27.04320	0.00134	2.754875
5	-7.22435	0.00088	2.644184
6	17.29047	0.00003	2.397378
7	43.25958	0.00001	2.108418
8	71.92473	0.00000	1.741312
9	102.65613	0.00000	1.314315
10	126.90319	0.00004	1.075270
11	163.94939	0.00000	0.465235
12	201.98452	0.00000	-0.173464

Source: Own elaboration

On the basis of results from Table 3.1 we can see that both LR test and AIC criterion confirmed that the optimum value of delay for the model is $k = 12$. It is also worth noticing that the value of delay $k = 12$ was adopted as the maximum one for the model (1) at the beginning of the analysis due to the monthly type of data.

Testing the significance of parameters at delayed variables (causality in the Granger sense)

This part of the paper is devoted to causality and the method of testing it. Causality is quite a significant issue in a widely understood empirical research. Economists are interested, for example, whether the price of oil in the world will cause the fall in car sales and whether this relation works in the opposite direction. In order to test such hypotheses, we should establish the definition of causality which will allow its verification. In econometrics the most widely applied operational definition of causality was developed by Granger. According to it, variable X is a cause (in Ganger sense) of variable Y, which is written as $X \to Y$, if the current values y may be obtained on the basis of forecasts with greater precision using future values x, than in any other way, applying the principle of *ceteris paribus* (Charemza and Deadman, 1992).
Looking through the prism of the above definition, the problem of verifying the hypothesis whether $X \to Y$, boils down to establishing whether variable X may be removed from the equation of the VAR model which describes variable Y. Specialist literature offers numerous tests helping us examine causal relations. The comparison of various tests concerning causality can be found in the work of Geweke, Meese and Dent (1983). Below, we will describe the F test for Granger causality on the example of model (1). The choice of this test was determined by the fact that it can be carried out using the GRETL software, in which all calculations for this work were made.
The F test consists in checking whether delayed values (of, for example variable Y_t) have considerable significance in projecting another variable (for example X_t) in model (1). Therefore,

for the illustrating purposes we will analyze the limited part of model (1), consisting of only two equations and two variables:

$$
\begin{aligned}
x_t &= \alpha_{1,0} + \sum_{j=1}^{12} \alpha_{1,j} x_{t-j} + \sum_{j=1}^{12} \beta_{1,j} y_{t-j} + \varepsilon_{1,t} \\
y_t &= \alpha_{2,0} + \sum_{j=1}^{12} \alpha_{2,j} x_{t-j} + \sum_{j=1}^{12} \beta_{2,j} y_{t-j} + \varepsilon_{2,t}
\end{aligned} \tag{7}
$$

The null hypothesis stating that variable Y_t is not the cause in Granger sense of variable X_t has in F test the following form:

$$H_0 : \beta_{1,1} = \beta_{1,1} = \ldots = \beta_{1,12} = 0 \,,$$

will contrasted with the alternative hypothesis stating that Y_t is the cause in Granger sense of variable X_t. The test of the hypothesis will be F statistics expressed with the following formula:

$$F = \frac{\left(RSS^{'}_R - RSS_{UR}\right)/M}{RSS_{UR}/\left(N-K-1\right)} \sim F\left(M, N-K-1\right) \tag{8}$$

where: M – number of parameters in a limited model, K – number of parameters in an unlimited model, RSS_R – the residual sum of squares for the limited model, RSS_{UR} – the residual sum of squares for the unlimited model. F statistics has F distribution with the number of degrees of freedom $r_1 = K$ and $r_2 = N\text{-}K\text{-}1$. In the description of formula (8) elements there is a term of limited and unlimited model which calls for more precision. Testing the hypothesis H_0, that variable Y_t is not the cause in Granger sense of variable X_t, the unlimited model will be the first equation from model (7), while the limited model will be expressed by the following equation:

$$x_t = \alpha_{1,0} + \sum_{j=1}^{12} \alpha_{1,j} x_{t-j} + \varepsilon_{1,t} \,. \tag{9}$$

For the limited and unlimited models we make calculations used to determine the value of F statistics. The hypothesis H_0 is rejected and thus it is stated that variable Y_t is the cause in Granger sense of variable X_t if the calculated value of F statistics is bigger than critical value found in F distribution tables for a particular number of degrees of freedom. In the opposite case, there are no grounds for rejecting H_0. Obviously, analogically we can test whether the causality relation between variables X_t and Y_t appears in the opposite direction.

Table 4: Results of the F test for Granger causality for variables X_t, Y_t, Z_t and W_t

	X_t	Y_t	Z_t	W_t
X_t	---------------	$X_t \rightarrow Y_t$ $F=1.61$ $p_v = 0.125$	$X_t \rightarrow Z_t$ $F=1.014$ $p_v =0.453$	$X_t \rightarrow W_t$ $F=1.126$ $p_v =0.365$
Y_t	$Y_t \rightarrow X_t$ $F=1.246$ $p_v = 0.285$	-----------------	$Y_t \rightarrow Z_t$ $F=0.34$ $p_v =0.976$	$Y_t \rightarrow W_t$ $F=0.737$ $p_v =0.708$
Z_t	$Z_t \rightarrow X_t$ $F=2.172$ $p_v = 0.031$	$Z_t \rightarrow Y_t$ $F=0.784$ $p_v =0.663$	----------------	$Z_t \rightarrow W_t$ $F=2.561$ $p_v =0.012$
W_t	$W_t \rightarrow X_t$ $F=1.15$ $p_v =0.3483$	$W_t \rightarrow Y_t$ $F=1.896$ $p_v =0.062$	$W_t \rightarrow Z_t$ $F=2.868$ $p_v =0.0055$	-----------------

Source: Own elaboration

In the next section of this chapter we will perform the above-described F test for the existence of the following causal relations for the model expressed by the following equation (1):

$$X_t \rightarrow Y_t,\ Y_t \rightarrow X_t,\ X_t \rightarrow Z_t,\ Z_t \rightarrow X_t,\ X_t \rightarrow W_t,\ W_t \rightarrow X_t$$
$$Y_t \rightarrow Z_t,\ \ Z_t \rightarrow Y_t,\ \ Y_t \rightarrow W_t,\ \ W_t \rightarrow Y_t,\ \ Z_t \rightarrow W_t,\ \ W_t \rightarrow Z_t .$$

Limited and unlimited models for testing the above causal relations have the analogical form to the first equation of model (7) and model (9) for the causal relation $Y_t \rightarrow X$ and will not be presented here due to spatial constraints of this article. The results of the F test for all causal relations obtained in the GRETL program are included in Table 4.

Analyzing the results from Table 4.1 we can notice that the F test confirmed the existence of only the following causal relations: $Z_t \rightarrow X_t$, $Z_t \rightarrow W_t$ and $W_t \rightarrow Z_t$. In each case, adopting the significance level of 0.05 we reject the hypothesis H_0 stating that there is no causal relation between analyzed variables.

On the basis of obtained results we can state that:

1. the monthly inflation for particular goods in Great Britain is the cause in Granger sense of the monthly inflation for particular goods in Poland,
2. the monthly inflation for particular goods in Great Britain is the cause in Granger sense of the monthly inflation for particular goods in Germany,
3. the monthly inflation for particular goods in Germany is the cause in Granger sense of the monthly inflation for particular goods in Great Britain.

Harmonized Indices of Consumer Prices

The HICP inflation ratio was defined in the Commission Regulation (CE) No 2495 from 23rd October 1995 on Harmonized Indices of Consumer Prices (Official Journal L 257 from 27.10.1995) as final money expenses on households. Article 9 indicates that this ratio should be the Laspeyres

type of index (price changes month to month are measured as the average of price indices reflecting weights of expenses in accordance with the structure of consumption and prices and the index based on weighted period in the area of member states beginning with the ratio for January 1997). HICP is not the measure of changes of minimum costs of achieving the same living standards in 2 compared periods where only costs are measured (Wynne, 2002, p. 12).
According to article 3 of the above regulation, the scope of HICP covers goods and services which are offered for sale in order to satisfy direct needs of consumers in a particular economic area of a Member State. Price indices of consumer goods are observed not only in the procedure aiming at taking into account goods and services which have become significant recently, not only for inflation, and whose weights are updated annually, but also those whose weights are updated less frequently (*Commission Regulation,* 2001, p. 217).
Price indices for basic HICP aggregates are made using one of two formulas included in section 1 of attachment II of the above Regulation or using a comparable formula which leads to the index not differing regularly by more than a tenth of percentage point compared to previous year from the index calculated on the basis of one of two formulas (Wynne, 2002, p. 12).
HICP as the Laspayres index:

$$HICP_{t,t} = \sum_{i} \omega_{i,j,b} \left(\frac{p_{i,j,t}}{p_{i,j,r}} \right) \tag{10}$$

– $p_{i,j,t}$ – the price of i-th product in country j in comparison to month t,
– $p_{i,j,r}$ – the price in the reference month r,
– $\omega_{i,j,b}$ – the weight assigned to i-th product in the share of expenses in base expenses or in base period b.

As we already mentioned, weights used to aggregate basic aggregates can be based on expenditure patterns from up to seven past years for which HICP is constructed, however, some countries update their weights annually. The reference period for HICP sub-indices is 1996=100, and for new sub-indices introduced in January 2000 it is December 1999=100 (Wynne, 2002, p. 13). $p_{i,j,t}$ or relative prices $(p_{i,j,t}/p_{i,j,b})$ are calculated as arithmetic or geometric mean, dpending on conditions. Relative prices should be calculated as $(p_{i,j,t}/p_{i,j,b}) = ((\frac{1}{n})\sum_k p_{i,j,t}/(\frac{1}{n})\sum_k p_{i,j,b})$ (lower index b or t) or $(p_{i,j,t}/p_{i,j,b}) = (\prod_k (p_{i,k,j,t} \,/\, p_{i,k,j,b}))^{\frac{1}{n}}$. The classification of goods and services in HICP is based on the classification of expenditures selected for the System of National Accounts from 1993.

Component price aggregates

According to Attachment II of the Commission regulation, for calculation of price indices of particular aggregates we should use arithmetic $\frac{\frac{1}{n}\sum p^t}{\frac{1}{n}\sum p^b}$ or geometric $\frac{[\prod p^t]^{\frac{1}{n}}}{[\prod p^b]^{\frac{1}{n}}}$ mean. In both fractions: p^t is the present price, p^b is the reference price (*Commission Regulation,* 2001, p. 217).

Arithmetic mean of relative prices

The price index for elementary (component) aggregates may also be calculated as the chain index by means of two formulas. The first one is the relation of arithmetic means (*Commission Regulation,* 2001, p. 218):

$$I^{tb} = \frac{\sum_{i \in S_b} p_i^1}{\sum_{i \in S_b} p_i^b} \cdot \frac{\sum_{i \in S_1} p_i^2}{\sum_{i \in S_1} p_i^b} \cdots\cdots\cdots \frac{\sum_{i \in S_{b-1}} p_i^t}{\sum_{i \in S_{b-1}} p_i^{t-1}} \qquad (11)$$

p_i^t means i^{th} price for a given component aggregate in period t and S_t, means the sample of prices obtained for a component aggregate in period t. The sample is usually updated every month. This choice may, in practice, be monthly or, usually, if prices are not available, it may be updated after a longer period of time.

If between the base period b and period t there are no updates, then I^{tb} takes the form $I^{tb} = \frac{\sum_{i \in S_b} p_i^t}{\sum_{i \in S_b} p_i^b}$, which is the form of simple relation of arithmetic means. The relative price average cannot be applied if we use chain indices more than once a year (the above change in the applied formula also refers to geometric mean from point 1).

The analysis of calculations

In the course of calculations the adopted assumptions were fully confirmed, with the exception of the influence of German economy on Polish economy. This is quite surprising, as numerous data confirm great share of German economy in Polish trade and in the HICP index.

Table 6: Shares of country's inflation in the Euro-zone HICP in 1997-2011

	BE	**DE**	EE	IE	GR	ES	FR	IT	CY	LU	MT	NL	AT	PT	SI	SK	FI
1997	3.8	**34.5**	·	0.9	0.0	8.9	21.9	18.1	0.0	0.2	0.0	5.3	3.1	1.7	0.0	·	1.6
1998	3.8	**34.5**	·	0.9	0.0	8.9	21.9	18.2	0.0	0.2	0.0	5.3	3.0	1.7	0.0	·	1.6
1999	4.0	**34.5**	·	1.0	0.0	9.1	21.1	18.8	0.0	0.2	0.0	5.1	2.9	1.8	0.0	·	1.5
2000	4.0	**34.7**	·	1.0	0.0	9.1	20.9	18.3	0.0	0.2	0.0	5.7	2.9	1.8	0.0	·	1.5
2001	3.3	**30.9**	·	1.2	2.4	10.4	20.5	18.7	0.0	0.2	0.0	5.3	3.3	2.1	0.0	·	1.6
2002	3.4	**30.6**	·	1.2	2.5	10.3	20.4	19.3	0.0	0.3	0.0	5.2	3.2	2.0	0.0	·	1.6
2003	3.3	**29.9**	·	1.3	2.6	10.9	20.5	19.2	0.0	0.3	0.0	5.4	3.2	2.1	0.0	·	1.6
2004	3.3	**29.2**	·	1.3	2.7	11.1	20.7	19.3	0.0	0.3	0.0	5.3	3.1	2.1	0.0	·	1.6
2005	3.3	**29.0**	·	1.3	2.7	11.4	20.7	19.2	0.0	0.3	0.0	5.2	3.1	2.1	0.0	·	1.6
2006	3.4	**28.7**	·	1.3	2.9	12.0	20.3	19.1	0.0	0.3	0.0	5.2	3.1	2.2	0.0	·	1.6
2007	3.4	**28.2**	·	1.4	3.1	12.3	20.7	18.3	0.0	0.2	0.0	5.3	3.1	2.1	0.3	·	1.6
2008	3.4	**27.0**	·	1.5	3.4	12.7	20.5	18.6	0.2	0.3	0.1	5.0	3.1	2.2	0.3	·	1.6
2009	3.4	**26.1**	·	1.6	3.5	12.8	20.6	18.5	0.2	0.3	0.1	5.1	3.0	2.2	0.4	0.7	1.7
2010	3.2	**26.2**	·	1.5	3.6	12.6	20.8	18.2	0.3	0.3	0.1	5.1	3.0	2.2	0.4	0.7	1.7
2011	3.3	**25.9**	0.1	1.3	3.8	12.7	20.7	18.5	0.3	0.3	0.1	4.8	3.2	2.2	0.4	0.7	1.7

Source: ECB 2011, http://www.ecb.int/stats/prices/hicp/html/hicp_coicop_inw_000000.4.U2W.en.html

BE Belgium, DK Denmark, DE Germany, EE Estonia, IE Ireland, EL Greece, ES Spain, FR France, IT Italy, CY Cyprus, LU Luxemburg, MT Malta, NL Netherlands, AT Austria, PT Portugal, SI Slovenia, SK Slovakia, FI Finland.

The power of German economy influence can be seen in the table containing changing shares of country inflation in HICP in the Euro-zone in 1997-2011. In all this period Germany created at least ¼ of this index in Euroland (**Table 6**). Because of this we can assume that German inflation must exert strong influence on all European Union countries, including its direct neighbor, Poland.

The influence of German economy on Polish economy is also confirmed by data from **Table 7**. They show that this country has been our main trading partner for the last 20 years, influencing in this way various prices. As our calculations have brought the unconvincing lack of influence of German inflation, this calls for further analyses. Simultaneously **Table 7,** although it does not show any other analyzed country (Slovakia, Great Britain) as our main trading partner, for each of them our calculations have allowed us to prove their inflation influence on our economy.

Table 7: The top three countries with the biggest share of Polish imports and exports (1990-2009)

		I partner	share in %	II partner	share in %	III partner	share in %
1990	Import	Germany	20.1	USSR	19.8	Italy	7.5
	Export	Germany	25.1	USSR	15.3	Great Britain	7.1
1995	Import	Germany	26.6	Italy	8.5	Russia	6.7
	Export	Germany	38.3	Netherlands	5.6	Russia	5.6
2000	Import	Germany	23.9	Russia	9.4	Italy	8.3
	Export	Germany	34.9	Italy	6.3	France	5.2
2005	Import	Germany	24.7	Russia	8.9	Italy	7.1
	Export	Germany	28.2	France	6.2	Italy	6.1
2007	Import	Germany	24.1	Russia	8.7	China	7.1
	Export	Germany	25.9	Italy	6.6	France	6.1
2008	Import	Germany	23	Russia	9.7	China	8.1
	Export	Germany	25	France	6.2	Italy	6
2009	Import	Germany	22.4	China	9.3	Russia	8.5
	Export	Germany	26.2	France	6.9	Italy	6.9

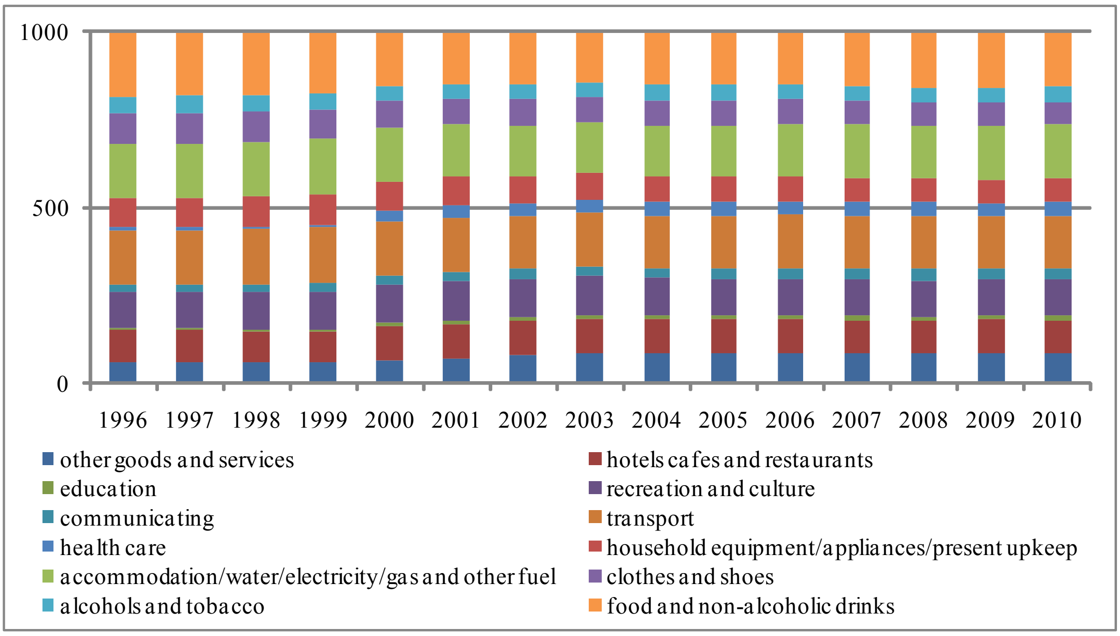
Source: GUS 2010, p. 37

Conclusions

The main reason why we did not manage to demonstrate the influence using our method is the fact that the general HICP index is too general, with particular goods having assigned appropriate weights as far as quantity and price are concerned. Therefore we can draw a conclusion that the applied method will allow us to confirm the foreign influences within sector HICP indices. Below we present HICP limitations (Diewert, 2002, p. 44-67), which seem to influence in such a way that the applied method does not confirm the influences resulting from trading bonds and from geographic proximity of Germany:

1. high speed of removing goods and services of decreasing inflation significance and introducing new ones accounts for considerable difficulties in using the Laspeyres price index with fixed base. This is due to the fact that from the perspective of the significance of the product or service, close and full replacement cannot last longer than a few months,
2. Eurostat must give a lot of guidelines concerning when a particular country should change its base year. However, we must remember about the existence of a period lasting up to 7 years of maintaining a particular base year by a given country. This situation is conducive to lowering harmonization of an essential element of HICP, hindering reliable comparisons and analyses of influences Another methodological constraint was the problem of selecting the index in its chain version and its fixed base, which is considered one of controversial issues related to improvement of harmonization of inflation analyses (*Commission Regulation* 2001, p. 186),
3. it is worth remembering that HICP is defined as weighted average of sub-indices (sector indices), with all the consequences of averaging. In order to use a particular price in calculating HICP (as the Laspeyres type of index), we must determine a particular representative sample for each group of basic goods according to prices collected monthly. In order to prepare data we must average long-term relative prices, which is considered another problem limiting the use of HICP,
4. HICP shows the tendency to lowering inflation, which could be manifested in the results of the applied vector autoregression method and in the Granger causality test, indicating lack of influence of German economy on Polish economy. In spite of intensifying the research on general inflation in EU, we still feel shortages in particular categories of HICP,
5. scientists often criticize inaccuracy of research on inflation in EU and shortages in its verification (Diewert, 2002, p. 6-7). They even point at too low measures of individual prices within general inflation tendencies, which accounts for the fact that some consumer expenditure is not well reflected in HICP. Among other constraints, specialists also mention lack of appropriate full description of data preparation and HICP calculation by statistical offices in particular countries, which may result from shortage of information on full implementation of reference documentation concerning HICP (Diewert, 2002, p. 14),
6. Since 1998, member states have been obliged to update weights in the HICP index annually. However, a serious problem is caused by the phenomenon of seasonality of goods, which means their consumption level changes in the year. Some of these goods are practically not bought in particular periods of time, aggravating the problem due to their share in shaping inflation phenomena up to 30% in some countries. Even if the seasonal good is available in all months of the year, we systematically encounter the problem of quantity changes in relation to average weights of expenses from the base year. Therefore, for at least a few months in a year, these annual averaged weights will not reflect the changes concerning seasonal goods, even due to excessive freedom in weight revision given to member states.

Figure 1: Weighted composition of product groups in the HICP index basket in the EU (1996-2010)



Source: European Commission> Economic and Financial Matters >Focus on > Inflacja w centrum uwagi 2011. Retrived from: http://ec.europa.eu/economy_finance/focuson/inflation/measuring_pl.htm

Table 8: Polish – German trade in 2009 (in USD thousand)

	Import	Export
TOTAL	**33418473**	**35679585**
Food and livestock	2388137	3135287
Beverages and tobacco	98480	120356
Inedible material excluding fuels	700323	977628
Mineral fuels, lubricants and derivatives	1375198	764717
Oils, animal and plant fat and wax	147570	83836
Chemical products and derivatives	**5937486**	1923434
Industrial goods by raw materials	**8026644**	**7355002**
Machines, equipment and transport vehicles	**11817336**	**15007479**
Miscellaneous industrial goods	2507837	**6294620**
Goods and transactions not classified by SITC	419461	17227
The share of the three main trade elements	77%	80%

Source: GUS 2010, p.132-133

In spite of the fact that in each country the main inflation factors are food, transport and costs of household upkeep (as demonstrated by the weights in Figure 1 in EU-25 in 1996-2010), the structure of import and export does not coincide with the HICP elements with the biggest share in Polish-German trade. For example the 3 main points of this trade (Table 8) in imports account for 77% (chemical products and derivatives, industrial goods by raw materials, machines, equipment and transport vehicles) and in exports – 80% (miscellaneous industrial products, industrial goods by raw materials, machines, equipment and transport vehicles). This situation indicates further research directions, concentrating on sector HICP inflation indices. The data included here indicate that at this level the applied calculation methods will allow us to obtain consistency with other data showing different channels of influence. This research will be reported in further publications, as we should note that in spite of two decades that have passed HICP is still a developing project.

References

Box, G. E. P., Jenkins, G. M. (1970). *Time Series Analysis: Forecasting and Control*. San Francisco: Holden- Day.

Charemza, W. W., Deadman, D. A. (1992). *New Direction in Econometric Practice.* Hants: Edward Elgar.

Commission Regulation, (2001). (EC) No 1749/96 of 9 September 1996 on initial implementing measures for Council Regulation (EC) No 2494/95 concerning harmonized indices of consumer prices in: Office for Official Publications of the European Communities, *Compendium of HICP reference documents (2/2001/B/5),* Luxemburg.

Diewert, E. (2002). Harmonized Indexes of Consumer Prices: Their Conceptual Foundations. Working Paper Series Working Paper 130. European Central Bank March. Retrieved from: http://www.ecb.int/stats/prices/hicp/html/hicp_coicop_inw_000000.4.U2W.en.html.

Enders, W. (1995). *Applied Econometric Time Series*. New York: John Wiley & Sons, Inc.

Geweke, J., Meese, R., Dent, W. (1983). Comparing Alternative Tests of Causality in Temporal System. *Journal of Econometrics,* 21, 161 -194.

Główny Urząd Statystyczny (2010). *Rocznik statystyczny handlu zagranicznego*. Komisja Europejska> Sprawy Gospodarcze i Finansowe> Focus on> Inflacja w centrum uwagi 2011. Retrieved from: http://ec.europa.eu/economy_finance/focuson/inflation/measuring_pl.htm.

Kufel, T. (2007). *Ekonometria. Rozwiązywanie problemów z wykorzystaniem programu GRETL*. Warszawa: Wydawnictwo Naukowe PWN.

Kusideł, E. (2001). Modele wektorowo-autoregresyjne VAR. Metodologia i zastosowania*.* In: B. Suchecki (ed.), *Dane panelowe i modelowanie wielowymiarowe w badaniach ekonomicznych*. Volume 3. Łódź: Wyd. Absolwent.

Maddala, G. S. (2006). *Ekonometria*. Warszawa: Wydawnictwo Naukowe PWN.

Pesaran, M., Pesaran, H. B. (1997). *Working with Microfit 4.0*. New York: Oxford University Press.

Piłatowska, M. (1997). *Alternatywne metody eliminacji trendu a interpretacja modelu ekonomicznego.* Materiały na IV Ogólnopolskie Seminarium Naukowe w Toruniu. Katedra Ekonometrii i Statystyki UMK.

Wynne, M. A., Rodriguez-Palenzuela, D. (2002). *Measurement Bias in the HICP: What Do We Know, and What Do We Need to Know?* Working Paper no. 131. Working Paper Series. European Central Bank.