NATIONAL BANK OF THE REPUBLIC OF BELARUS

BANK RESEARCH PAPERS
1/2004

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This paper is published in accordance with Resolution No. 73 of the Board of Directors of the National Bank of the Republic of Belarus dated 29.03.2004

The publication was prepared by the Monetary Policy and Economic Analysis Directorate of the National Bank of the Republic of Belarus.

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Valery Chernookiy*

SUMMARY

This paper discusses the econometric model of inflation processes in the Republic of Belarus which makes it possible to explain major factors determining the dynamics of the GDP deflator, consumer price index and producer price index during the period 1994 - 2003. For estimation of the model the author used the statistical tools of non-stationary time series econometrics: cointegration analysis and error-correction models. The model has good statistical properties, it demonstrates stability of the coefficients and enables one to conduct analysis of the various choices in the field of monetary and foreign exchange policies, as well as in the area of labour remuneration, prices and tariffs.

JEL classification: E31, C32, P24

Key words: inflation, GDP deflator, consumer price index, producer price index, the Republic of Belarus, cointegration

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“The Institute of the Economy of the National Academy of Sciences of Belarus”. This paper was worked out within the framework of the project aimed at studying inflation processes in the Republic of Belarus (Agreement No. 307/Д).

Econometric modelling of inflation processes in the Republic of Belarus

Under the conditions of macroeconomic instability that is currently typical of the Republic of Belarus, economic entities and individuals periodically face the issue of forecasting future inflation rates. The results of their decisions and, accordingly, related real losses and profits depend upon the accuracy of future inflation assessments. Forecasting inflation is also of critical importance for the monetary authorities. Taking into account the monetary nature of inflation, the central bank reconciles its short-term monetary and foreign exchange policies focused on stabilizing the output and unemployment rate with achieving a low level of inflation in the long run. Since these two objectives often contradict each other, it is necessary to estimate quantitatively the effects of the main monetary policy instruments on the inflation rates.

Forecasting inflation is a non-trivial problem and, taking into account the complexity of the inflation processes mechanisms, is practically impossible without applying special economic-mathematical models. Such models must reflect the main transmission channels through which monetary policy impacts prices and take into account a range of effects of the non-monetary factors; be dynamic in nature which makes it possible to trace various temporal effects of the inflation processes development; be sufficiently accurate in describing the actual situation, and, at the same time, be understandable as much as possible and simple in use.

The problem of inflation processes modelling in the Republic of Belarus is broadly presented in research papers of Belarusian economists who have been developing inflation models in two key directions: on the basis of econometric modelling methodology and on the basis of inter-industry balance models. The first direction is represented by the econometric models of inflation developed by I.V. Pelipas, A.O. Tikhonov, M.V. Pranovich, V.I. Malyugin, et al., that emphasize the monetary factors of inflation. As for the second direction models, they reflect the cost structure of individual industries and explicitly take into account inter-industry interrelationships of price indexes enabling thereby inflation processes analysis to be made based on cost determinants. The latter direction is represented by the works of V.N. Komkov, V.V. Pinigin and others.

This research paper is aimed at presenting a specific econometric model of inflation processes in the Republic of Belarus. The model has a quarterly periodicity and is based on the statistical data supplied by the Ministry of Statistics and Analysis (MSA) and the National Bank of the Republic of Belarus (NB RB) for the period 1994–2003. As distinct from the existing inflation models, it possesses a number of specific features.

First, the model *is dynamic in nature* that makes it possible to study the development of inflation processes in the Republic of Belarus both in the long and short run.

Second, *the model reflects both the demand-push and cost-push inflation mechanisms*. To do so, two key relationships were introduced into the model. The first of them represents the long-term relationship between inflation and the money supply thus making quantitative evaluation of inflation monetary factors possible. The second one describes the long-term interrelation between the price level dynamics and that of production costs. Thus, the latter relationship makes it possible to study the effect of cost-push inflation factors, in particular, the devaluation of the national currency or growth of imported fuel prices.

Third, *the model reflects dynamics not only of the general price level, but of the price structure as well*. Indeed, high inflation serves as a favourable background for significant changes in the relative prices of goods and services in the economy. In particular, consumer prices growth rates can substantially differ from growth rates of the producer goods prices. Dynamics of consumer goods prices and tariffs on services rendered to individuals, export goods prices and prices of domestic consumption goods can also vary. Taking this into account, inflation forecast based on the consumer price index may differ from the one based on other price indexes such as the producer price index or GDP deflator.

Forth, *the model makes it possible to take into account the fact that in the Republic of Belarus some prices, first of all, tariffs on housing and public utilities, passenger transportation and communication services, are subject to state regulation.* These prices, as a rule, are materially understated as compared with the prime cost of the services. A gradual increase in such prices is attributable to the necessity to eliminate the direct and cross subsidies. Nevertheless, the state authorities control their dynamics which allows us to regard such prices as exogenously determined. An index of tariffs on housing *and public utilities* was used in the model as an indicator of regulated prices. In 2003 the share of such services in the consumer basket amounted to 8.7 per cent.

Modern techniques of time series econometrics, such as cointegration analysis and error-correction models, were used for empirical estimation of the parameters of this model. This is due to the fact that a vast majority of the macroeconomic indicators of the Belarusian economy are non-stationary variables, which makes it impossible to apply the traditional methods, since one might be exposed to the risk of a "spurious regression" phenomenon.

The list of model variables, their designations and data source are given in Table 1. The price variables and indicators of real output are given in the form of indices calculated as a progressive total since December 1993 or Q4 of 1993. All variables in the model are given in the logarithmical form due to the necessity to linearise their dynamics.

The following three variables were used as inflation indicators: Δcpi is a logarithm of the consumer price index, Δipp is a logarithm of the producer price index and Δpy is a logarithm of the GDP deflator.

Table 1

Model Variables

Code	Name of variable	Data source
CPI	Consumer price index – CPI, calculated as a progressive total since December 1993	MSA RB[1], Statistical bulletin for 1994–2002
IPP	Producer price index – PPI, calculated as a progressive total since December 1993	MSA RB, Statistical bulletin for 1994–2002
PY	GDP deflator, calculated as a progressive total since Q4 1993	MSA RB, Statistical bulletin for 1994–2002
CPI^{G}	Consumer goods price index, calculated as a progressive total since December 1993	MSA RB, Statistical bulletin for 1994–2002
CPI^{S}	Consumer services price and tariff index, calculated as a progressive total since December 1993	MSA RB, Statistical bulletin for 1994–2002
IPP^{F}	Price index in fuel industry, calculated as a progressive total since December 1993	MSA RB, Statistical bulletin for 1994–2002
IPP^{USA}	Producer price index in the USA, calculated as a progressive total since December 1993	Bureau of Labour Statistics, the USA, www.freelunch.com
$M1$	Monetary aggregate M1, mln rubles, average per quarter	NB RB[2], Money review for 1994–2002
Y	Real GDP, bn rubles, in constant prices of 2000	MSA RB, Statistical bulletin for 1994–2002
Y^{IND}	Real industrial production growth index, calculated as a progressive total since December 1993	MSA RB, Statistical bulletin for 1994–2002
E^{MAR}	BYR/USD market nominal exchange rate, average per quarter	NB RB
E^{OF}	BYR/USD official nominal exchange rate, average per quarter	NB RB
I	Nominal refinancing rate of NB, % per annum, average per quarter	NB RB
W	Average monthly wage in the national economy, `000 rubles, average per period	MSA RB, Statistical bulletin for 1994–2002
W^{IND}	Average monthly wage in industry, `000 rubles, average per period	MSA RB, Statistical bulletin for 1994–2002

Considering inappropriateness of applying standard methods of model estimation for non-stationary time series and before proceeding to the description of its structure, it is necessary to test preliminarily the stationarity of the variables and to determine the order of their integration. For this purpose the augmented Dickey-Fuller test (ADF) was used. Its results (Table 2) show that all of the model variables are non-stationary and integrated of order one – I(1).

[1] MSA RB – Ministry of Statistics and Analysis of the Republic of Belarus.
[2] NB RB – National Bank of the Republic of Belarus.

Table 2

Results of testing variables for stationarity:
the augmented Dickey-Fuller test (ADF)

Variable	t_{ADF}	Test	Variable	t_{ADF}	Test
cpi	-1.178	ADF(1) with seasonal dummies	Δcpi	-3.832**	ADF(2) with seasonal dummies
ipp	-0.720	ADF(1) with seasonal dummies	Δipp	-3.737**	ADF(1) with seasonal dummies
py	-1.055	ADF(1) with seasonal dummies	Δpy	-2.112**	DF with seasonal dummies
cpi^G	-1.196	ADF(1) with seasonal dummies	Δcpi^G	-2.284**	DF with seasonal dummies
cpi^S	0.628	ADF(1)	Δcpi^S	-1.691*	DF
ipp^F	1.353	ADF(1)	Δipp^F	-3.262**	DF
ipp^{USA}	-1.764	ADF(1) with seasonal dummies	Δipp^{USA}	-3.666**	DF with seasonal dummies
$m1$	-1.153	ADF(1) with seasonal dummies	$\Delta m1$	-3.921**	ADF(2) with seasonal dummies
y	-2.378	DF with trend and seasonal dummies	Δy	-5.448**	DF with seasonal dummies
y^{IND}	-2.638	DF with trend and seasonal dummies	Δy^{IND}	-6.575**	DF with seasonal dummies
e^{MAR}	0.887	ADF(1)	Δe^{MAR}	-2.558**	DF
e^{OF}	0.979	ADF(1)	Δe^{OF}	-2.420**	DF
i	-1.184	ADF(2)	Δi	-3.776**	DF
w	0.807	ADF(1)	Δw	-2.504**	DF
w^{IND}	0.739	ADF(1)	Δw^{IND}	-2.544**	DF

The following approach was used for developing this model. In the first stage, the long-run relationships, reflecting theoretically justified interrelations of selected inflation indexes and their underlying factors from the demand side or from the cost side, were estimated on the basis of the Johansen cointegration analysis. Then, in the second stage, the short-term error-correction models for each inflation index were developed which allow us to describe dynamically the impact of all inflation factors under consideration. Structurally these models contain the mechanism for adjusting deviation of price indexes from their long-term levels.

** and * mean significance at 5% and 10% level, respectively.

Econometric model of the GDP deflator

Within the model of the GDP deflator the following two key long-term relationships were developed and estimated.

The first relationship that was empirically tested with the actual data available for the Republic of Belarus reflects the widely recognized fact that in the long run the evolution of inflation processes in the economy is determined exclusively by the monetary factors. This relationship describes the long-term equilibrium on the money market, i.e. such a condition when households` and firms` money demand is equal to the supply of money which is controlled by the monetary authorities.

Taking into consideration the fact that the supply of money can be treated as an exogenous variable determined by the monetary policy of the central bank, the analysis was focused on the estimation of the long-run money demand function. As is well known, households` and firms` demand for money in an economy is determined by the fact that they seek to hold a certain part of their incomes (or wealth) in liquid money for covering current expenditures (transaction motive). Since incomes are irregular, whereas expenditures are incurred on a regular basis, economic entities and individuals are bound to hold part of their incomes in cash form, so that, at any given time, they would have a cover for settling such expenses. However, money assets have a zero or very low nominal rate of return, and therefore these holdings entail costs relating to foregone incomes from alternative and more profitable but less liquid real or financial assets. On the other hand, such alternative assets cannot be directly used to effect payments for goods and services, whereas their conversion into money assets also entails certain costs (payment of commission fees, time expenditures, etc.). Thus, households and firms create demand for money assets as a result of minimization of the total costs of holding money and alternative financial assets. The amount of the real money demand is determined in such a case by the size of the real income received by economic entities during the period under review as well by the nominal rate of return on the cash and alternative assets.

In this model the following specification of the money demand function was used:

$$m1 = \alpha_0 + \alpha_1 py + \alpha_2 y + \alpha_3 i + \varepsilon\,,\ \alpha_1 > 0\,,\ \alpha_2 > 0\,,\ \alpha_3 < 0, \tag{1}$$

where $m1$ is a logarithm of the average quarterly volume of M1;

py is a logarithm of the GDP deflator;

y is a logarithm of GDP at prices of 2000 (indicator of the real income level);

i is a logarithm of the average quarterly annual nominal refinancing rate (indicator of the nominal return rate of the alternative assets such as term deposits in rubles and government bills);

ε is a random error defining temporal deviation from the long-term equilibrium in the money market.

Dynamics of the variables of model (1) is given in Figure 1. As one can see, the development of inflation processes in the Republic of Belarus during 1994–2002 was determined mostly by the growth of the money supply in excess of the growth of money demand.

The choice of M1 is justified by the fact that it includes the most liquid part of the money supply: cash in circulation and demand deposits, i.e. money that is directly involved in servicing the turnover of goods and services and to a greater extent than other less liquid money assets reflects transaction motive of individuals and economic entities. Undoubtedly, one may object that ruble term deposits constitute the resource base of the commercial banks which use these funds to provide lending to the economy. Thus, time deposits indirectly play the function of money as a means of circulation. However, the usage of borrowings by individuals or enterprises for settling their merchandises and services bills is reflected in an increased amount of cash in circulation or demand deposits, i.e. in the growth of M1.

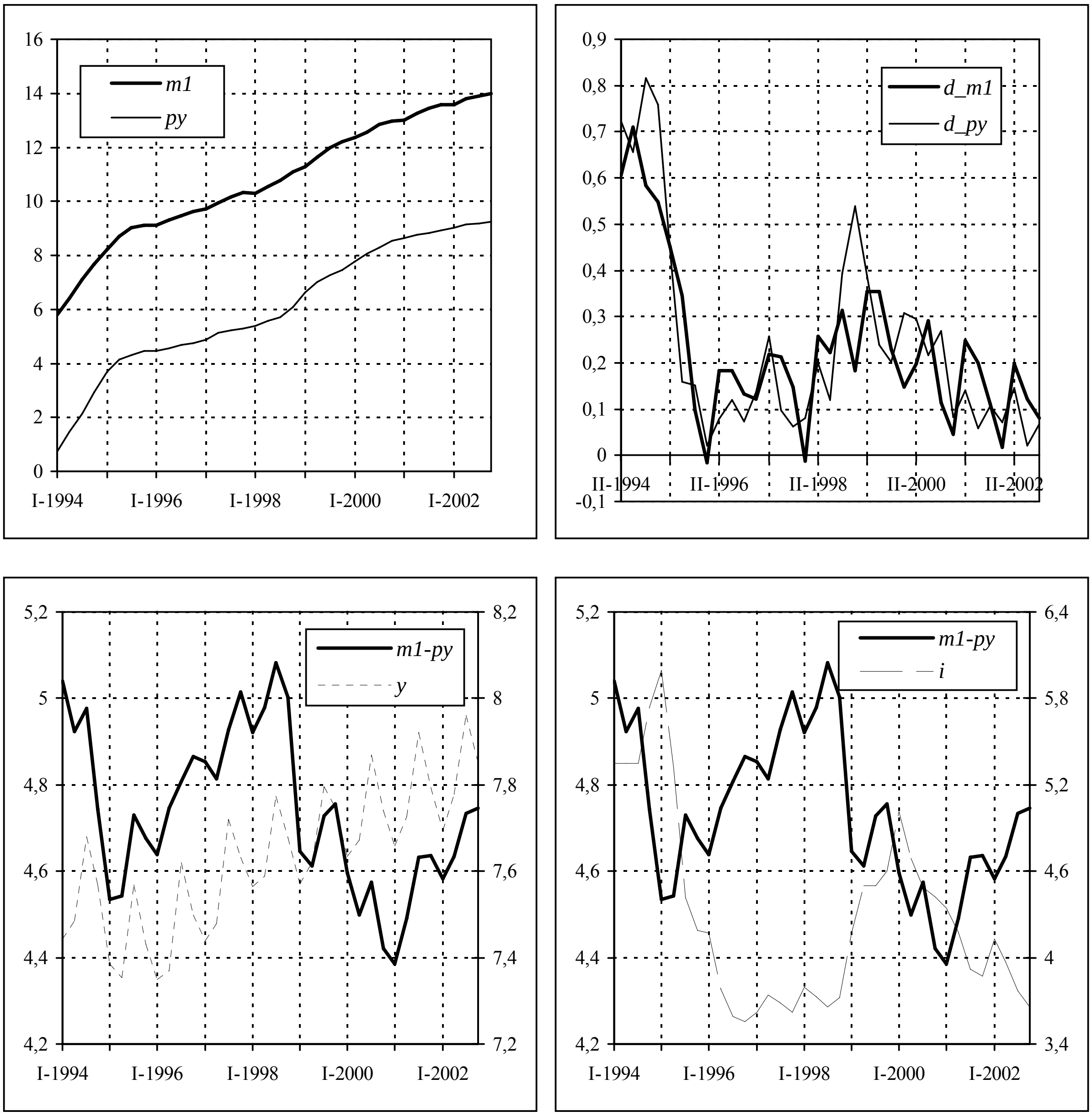


Figure 1. Dynamics of model (1) variables

In this paper the Johansen cointegration analysis[3] is used for estimation of the money demand function in the Republic of Belarus (1). To this end, the vector autoregression model (VAR) consisting of the variables $m1$, py, y, i and including seasonal dummies was estimated. Analysis of the VAR-model lag structure conducted on the basis of series of sequential Lagrange multiplier tests (LMF) and the use of the Akaike information criterion (AIC) and the Schwarz information

[3] Analysis was carried out in the Ox Professional 3.1. package.

criterion (SIC) revealed that there is no autocorrelation in residuals for the VAR(3)-model.

Results of the Johansen $\lambda_{\max}$ and λ_{trace} tests for cointegration (Table 3) confirm the existence of two cointegration vectors; however, only the first cointegration relationship has theoretically correct parameters. The money demand function is as follows:

$$m1 = -5{,}188 + 0{,}833py + 1{,}534y - 0{,}177i + error \, . \tag{2}$$

However, the assumption of the money demand homogeneity with respect to the price level does not hold true for equation (2) because the elasticity coefficient α_1 is equal to 0.833, which is much less than 1. Thus, price level growth by 1% is connected in the long run with that of the monetary aggregate M1 by 0.83 %, i.e. the inflation rate is growing faster than liquidity.

The lack of homogeneity can be explained by the elimination of multiple foreign exchange rates in 2000 that gave rise to a sharp devaluation of the official exchange rate, increase in critical import prices and, as a consequence, the GDP deflator is growing faster than the monetary aggregate M1. To support this assumption, cointegration analysis of modified relation (1) was made:

$$m1 = \beta_0 + \beta_1 py + \beta_2 y + \beta_3 i + \beta_4 e^{OF} + \varepsilon, \;\; \beta_4 < 0 \, , \tag{3}$$

where e^{OF} is a logarithm of the average quarterly official nominal exchange rate of the Belarusian ruble against the US dollar.

Analysis of the VAR-model lag structure for (3) showed that there is no autocorrelation in residuals of the VAR(3)-model. The results of the cointegration tests conducted for this model are shown in Table 4.

The results of the Johansen test confirm the existence of two cointegration vectors. At the same time, only the first cointegration relationship has economically reasonable parameters:

$$m1 = -6{,}205 + 1{,}021 py + 1{,}620 y - 0{,}187 i - 0{,}168 e^{OF} + error. \quad (4)$$

Table 3

Results of the cointegration analysis of model (2)

(A) Determining the cointegration rank				
Eigenvalue	0.720	0.456	0.304	0.072
Null hypothesis	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{max}	42.062**	20.074*	11.961	2.479
λ_{trace}	76.577**	34.514**	14.440*	2.479
(B) Standardized cointegrating vectors β'				
$m1$	py	y	i	const
1	-0.833	-1.534	0.177	5.188
-0,734	1	-3.705	-0.661	32.880
(C) Standardized adjustment coefficients α				
$m1$	-0.498	-0.129		
py	0.139	-0.213		
y	-0.013	0.037		
i	0.202	0.042		

As expected, the assumption of homogeneity with respect to the price level is satisfied for the function of form (4). Indeed, the coefficient β_1 proved to be equal to 1.021, practically coinciding with 1. The elasticity of money demand with respect to real GDP y is equal to 1.62. Consequently, an increase in the real output by 1% in the Republic of Belarus leads to an increase in money demand by 1.62%. At the same time, the estimate of the long-run elasticity with respect to the refinancing rate i is equal to -0.187. Thus, an increase in the refinancing rate induces growth of the nominal rate of return on alternative financial assets, such as time deposits and government short-term obligations, and decreases incentives for holding liquid cash funds, i.e. reduces the money demand on the part of economic entities and individuals.

Taking into account the price level homogeneity of the money demand function, a specific version of model (3) with restriction $\beta_1 = 1$ was constructed, having the following form:

$$m1 - py = \beta_0 + \beta_2 y + \beta_3 i + \beta_4 e^{OF} + \varepsilon. \quad (5)$$

Table 4

Results of the cointegration analysis of model (3)

(A) Determining the cointegration rank					
Eigenvalue	0.898	0.576	0.425	0.205	0.129
Null hypothesis	$r = 0$	$r \le 1$	$r \le 2$	$r \le 3$	$r \le 4$
λ_{max}	75.173**	28.327**	18.271	7.582	4.539
λ_{trace}	133.890**	58.719**	30.392**	12.121	4.539
(B) Standardized cointegration vectors β'					
$m1$	py	y	i	e^{OF}	const
1	-1.021	-1.620	0.187	0.168	6.205
-0,399	1	-2.823	-0.721	-0.319	24.201
(C) Standardized adjustment coefficients α					
$m1$	-0.672	-0.156			
py	0.640	-0.189			
y	-0.055	0.029			
i	0.391	0.077			
e^{OF}	1.165	-0.138			

Performing the Johansen cointegration analysis in this case (Table 5) confirmed the existence of a single cointegration vector, with parameters estimates being close to the values of model (4) parameters.

Thus, to reflect the monetary nature of inflation in the long term the present model uses the following relationship for the real money demand:

$$m1 - py = -6{,}596 + 1{,}671 y - 0{,}169 i - 0{,}156 e^{OF} + error. \quad (6)$$

The second important long-run relationship for the GDP deflator characterizes the inflation factors from the cost side and considers the price level as a function of costs per unit of the real output, taking into account the markup which reflects production profitability. Increased production costs in the economy as a whole necessitate increases in prices for some goods and services and, therefore, increases in the general price level. Indeed, enterprises are often forced to raise prices to maintain the rate of profitability necessary for their simple reproduction. However, increases in the price level can be restrained by both monetary policy and prices for the products of foreign competitors. In this case the markup may be reduced. By contrast, a declining general level of costs under steady conditions of monetary policy and foreign trade would be reflected in the increased rate of production profitability. Thus, the markup is fluctuating around its certain long-run level.

Table 5

Results of the cointegration analysis of model (5)

(A) Determining the cointegration rank				
Eigenvalue	0.723	0.396	0.269	0.020
Null hypothesis	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{max}	42.389**	16.657	10.348	0.668
λ_{trace}	70.063**	27.674*	11.016	0.668
(B) Standardized cointegration vectors β'				
$m1 - py$	y	i	e^{OF}	const
1	-1.671	0.169	0.156	6.596
(C) Standardized adjustment coefficients α				
$m1 - py$	-1.139			
y	-0.143			
i	0.016			
e^{OF}	0.474			

In the Republic of Belarus, the following factors have a major impact on the general cost level.

First, nominal wage growth in excess of productivity increase. This factor played a decisive role in the dramatic drop of production profitability in the Belarusian economy in 2001 when, against the background of a tightening monetary policy, a significant administrative increase in wages and salaries was observed and their real equivalent increased by 29.5 %.

Second, devaluation of the national currency resulting in sharp hikes in prices for intermediate imports and, consequently, in the total level of costs. All other factors being equal, increases in the price level abroad have a similar effect.

Third, instability of prices for energy inputs.

Thus, the following specification of the long-run price level equation on the cost side was used in the present model:

$$py = \alpha_0 + \alpha_1 w + \alpha_2 y + \alpha_3 ipp^{FUEL} + \alpha_4\left(e^{MAR} + ipp^{USA}\right) + \varepsilon\,,\ \alpha_1 > 0\,, \alpha_2 < 0\,, \alpha_3 > 0 \text{ and } \alpha_4 > 0\,, \quad (7)$$

where py is a logarithm of the GDP deflator;

w is a logarithm of the average quarterly wage in the national economy of the Republic of Belarus;

y is a logarithm of GDP at constant prices of 2000 (characterizing labour productivity in the national economy);

ipp^{FUEL} is a logarithm of the price level in the fuel industry;

e^{MAR} is a logarithm of the average quarterly market nominal exchange rate of the Belarusian ruble to the US dollar;

ipp^{USA} is a logarithm of the producer price level in the USA;

ε is a random error characterizing the temporal deviation of the markup from its long-run level.

Dynamics of the model (7) indicators shown in Figure 2 confirms that all of the above factors on the cost side affected the development of inflation processes in the Republic of Belarus in 1994-2002.

The Johansen cointegration approach was used for estimation of long-run relationship (7). Analysis of the VAR-model lag structure based on a number of consecutive Lagrange multiplier tests (LMF) and on the use of the Akaike information criterion (AIC) and the Schwarz information criterion (SIC) showed that there is no autocorrelation in residuals of the VAR(3)-model.

Results of the Johansen $\lambda_{\max}$ and λ_{trace} cointegration tests (Table 6) confirm the existence of two cointegration vectors. At the same time, only the second cointegration relationship has theoretically correct parameters after relevant normalization (the negative coefficient α_2). Thus, the long-run relationship reflecting the development of inflation processes on the costs side has the following form:

$$py = 9{,}077 + 0{,}745w - 1{,}433y + 0{,}195ipp^{FUEL} + 0{,}082\left(e^{MAR} + ipp^{USA}\right) + error. \qquad (8)$$

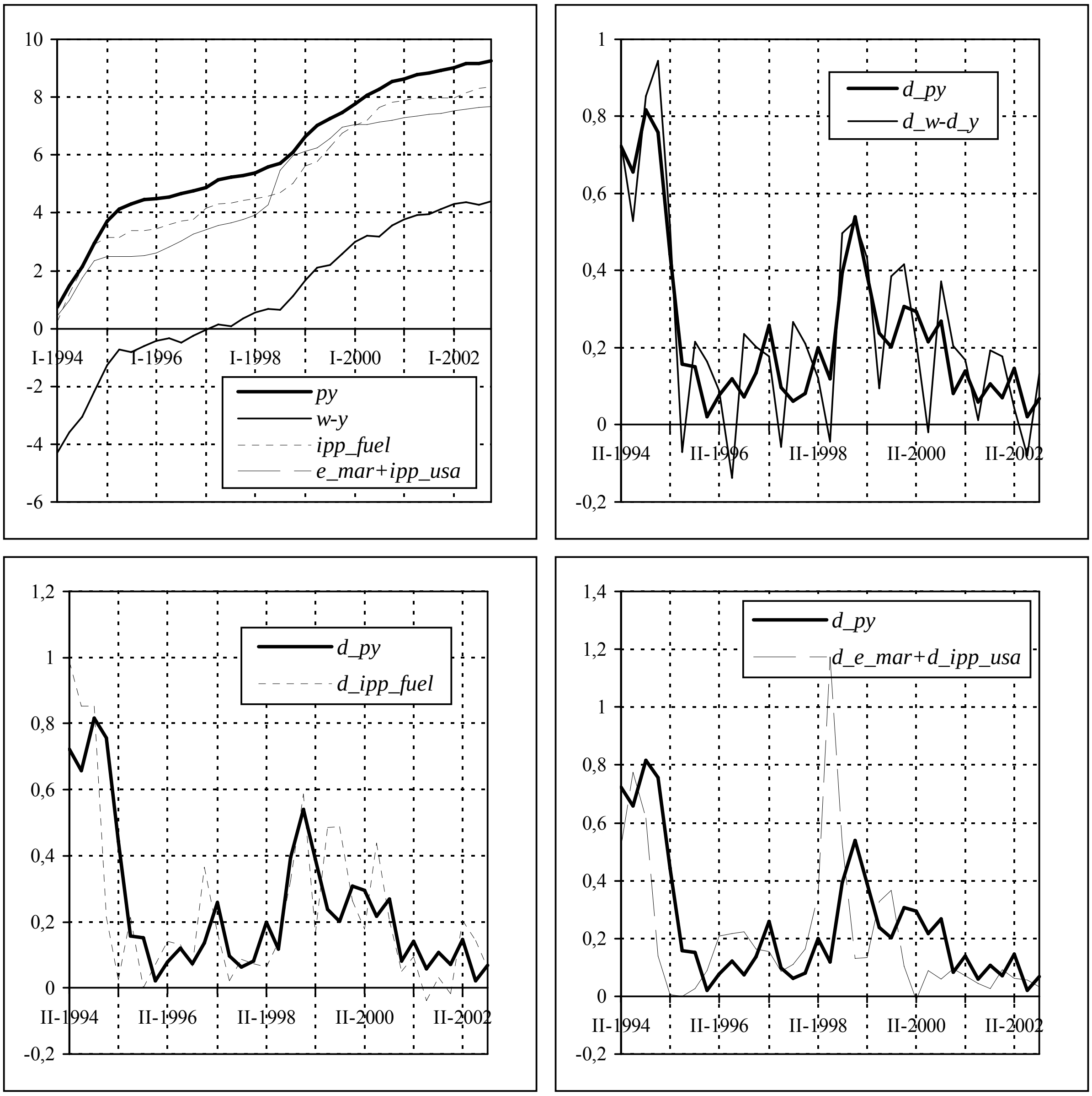


Figure 2. Dynamics of model (7) indicators

Proceeding from equation (8) we can make the following conclusions. Wage increases by 1% under constant labour productivity result in a 0.745% increase in the GDP deflator. On the other hand, labour productivity growth by 1% under wage rigidity is associated with a fall in the price level by 1.433%. A 1% increase in prices for fuel industry products which is attributable either to rising world prices for energy resources or to the devaluation of the official exchange rate of the Belarusian ruble would cause a 0.195 % increase in the price level. At the same time, devaluation of the market exchange rate by 1% or inflation in foreign countries would bring about a 0.082 % increase in the GDP deflator.

Thus, the signs of relationship (8) parameters coincide with theoretical assumptions. Besides, the homogeneity condition $\alpha_1 + \alpha_3 + \alpha_4 = 1$ is also practically satisfied, since the actual value of elasticity of the price level with respect to the major cost factors amounted to 1.022. However, the condition $\alpha_1 = -\alpha_2$ for given equation is not satisfied, that is, the assumption that the growth of wages caused by the growth of labour productivity should not influence the price level. Indeed, the elasticity with respect to wage (0.745) turned out to be practically two times less in absolute value than the elasticity with respect to real GDP (-1.433).

Table 6

Results of model (7) cointegration analysis

(A) Determining the cointegration rank					
Eigenvalue	0.729	0.609	0.393	0.173	0.030
Null hypothesis	$r = 0$	$r \le 1$	$r \le 2$	$r \le 3$	$r \le 4$
λ_{max}	43.033**	30.947**	16.481	6.249	0.989
λ_{trace}	97.698**	54.665**	23.718	7.238	0.989
(B) Standardized cointegration vectors β'					
py	w	y	ipp^{FUEL}	$e^{MAR} + ipp^{USA}$	const
1	-0.546	-0.929	-0.030	-0.265	7.206
-1,343	1	-1.925	0.262	0.110	12.190
(C) Standardized adjustment coefficients α					
py	-0.447	0.708			
w	0.188	0.053			
y	0.039	-0.250			
ipp^{FUEL}	-1.044	0.026			
$e^{MAR} + ipp^{USA}$	-1.387	-2.653			

Nevertheless, relationship (8) confirms, on the whole, the impact of the stated factors on the development of cost-push inflation in the Republic Belarus and allows us to estimate the long-run tendencies in the dynamics of production costs.

Estimated cointegration relationships (6) and (8) specifying the long-run factors of inflation development on the demand side and on the cost side, respectively, were used for developing a short-run model of the GDP deflator. To this end, the following error-correction model was considered:

$$\Delta py = \pi_0 + \sum_{s=1}^{3} \pi_s seas_s + \alpha_1 ecm\_py_{-1}^{MONEY} + \alpha_2 ecm\_py_{-1}^{MARKUP} + \sum_i \beta_i \Delta py_{-i} + \sum_j \Gamma_j^T \Delta X_{-j} + \upsilon, \quad (9)$$

where $X = \left(m1, y, i, e^{OF}, w, ipp^{FUEL}, e^{MAR} + ipp^{USA}\right)^T$;

$ecm\_py^{MONEY} = py - m1 + 1{,}671y - 0{,}169i - 0{,}156e^{of} - 6{,}596$;

$ecm\_py^{MARKUP} = py - 0{,}745w + 1{,}433y - 0{,}195ipp^{FUEL} - 0{,}082\left(e^{MAR} + ipp^{USA}\right) - 9{,}077$;

$seas_s$ is a quarterly seasonal dummy that equals 1 for the quarter s and 0 for other quarters;

υ is a random error.

Taking into account the previously ascertained fact that the model variables are integrated of order one and, therefore, their first differences are stationary and that relationships (6) and (8) are cointegrated ones and, thus, variables ecm_py^{MONEY} and ecm_py^{MARKUP} are stationary as well, the use of the least squares method is valid and allows us to receive consistent estimates of the coefficients for the model. A number of tests for parameters were conducted which made it possible to ignore the insignificant factors and to obtain the model of the short-run GDP deflator dynamics. Estimates of the model parameters, its statistical characteristics and results of testing residuals of the model are given in Table 7.

The data of this table show that the model thus developed has good statistical characteristics and satisfies a set of criteria on its residuals. For example, the Lagrange multipliers tests (LM) prove the lack of autocorrelation, autoregressive conditional heteroskedasticity and the White's heteroskedasticity in residuals of the model, while the Jarque-Bera test confirms their normality. The determination coefficient R^2 is equal to 0.982, whereas $\overline{R^2}$ is 0.978, which characterizes high explanatory capability of the model. The model parameters have theoretically correct signs. The factor ecm_py^{MONEY} proved to be insignificant, whereas the speed of adjustment of the GDP deflator to the long-run level of production costs amounted to -0.444.

Table 7

Results of model (9) estimation

Variable	Estimates	t-statistics
$ecm\_py_{-1}^{MARKUP}$	-0.444	-7.423**
$\Delta m1$	0.393	5.296**
$\Delta m1_{-1}$	0.331	6.297**
Δy	-0.541	-10.002**
Δe^{OF}	0.116	3.265**
Δe_{-1}^{OF}	0.245	6.998**
Δe_{-1}^{MAR}	0.112	3.365**
$const$	-0.034	3.549**
R^2	0.982	
$\overline{R^2}$	0.978	
$s.e.$	0.025	
DW	2.438	
$AR:\chi^2(4)$	4.398	
$ARCH:\chi^2(4)$	1.165	
$Jarque-Bera:\chi^2(2)$	1.922	
$White:\chi^2(7)$	10.885	

The model thus developed allows us to make sufficiently accurate forecasts of the GDP deflator changes. For example, Figure 3 depicts dynamics of actual inflation in 1994–2002, its forecast made on the basis of the present model, as well as the forecast errors. It is worthwhile noting that, while the forecasts in the left-hand chart were developed on the basis of the model estimated on data for the period 1994–2002, i.e. are retrospective, the right-hand chart shows the dynamics of the model forecast estimated on data for the period 1994–2000. Thus, for the period 2001–2002 the right-hand chart gives a perspective forecast under known explanatory variables. Both charts confirm that the actual inflation is within the 95% confidence interval for the central forecasting value.

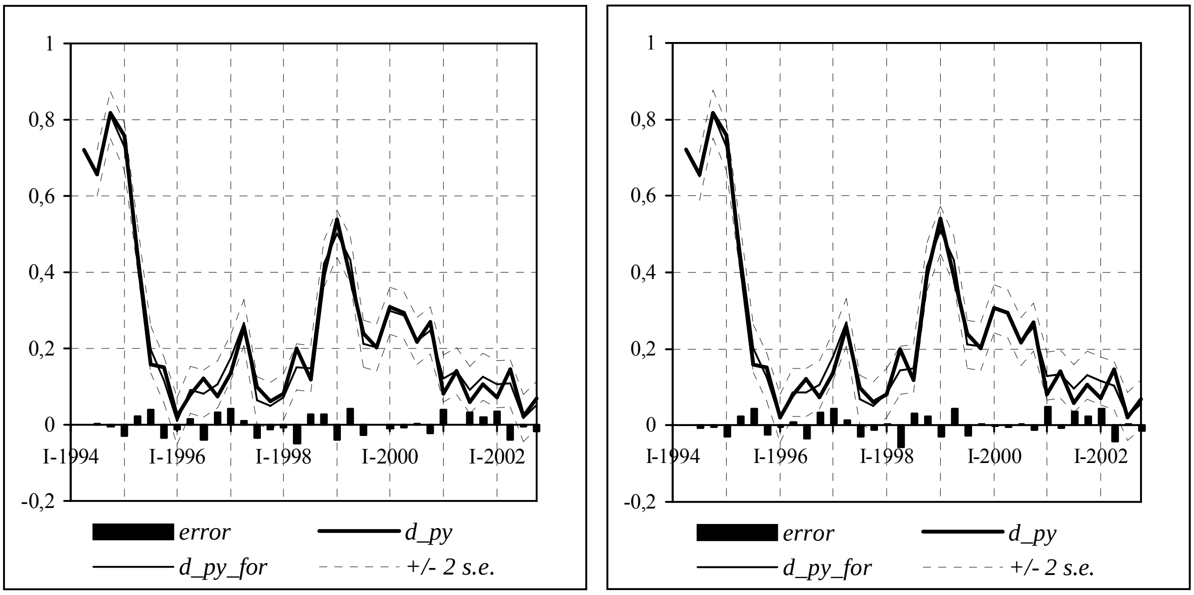
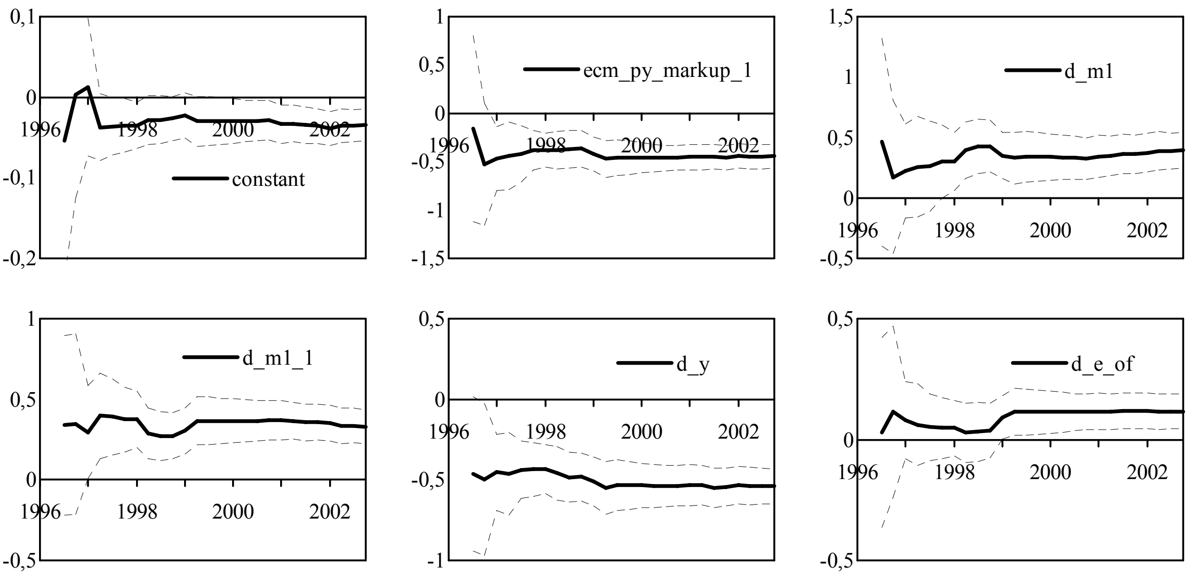


Figure 3. Forecast of the GDP deflator changes on the basis of model (9)

Figure 4 shows dynamics of the recursive estimates of model (9) coefficients. As can be seen from the charts, the coefficients of model (9) are insensitive to changes in the sample size for model estimation which makes it possible to conclude that the model is stable. The results of various Chow tests on stability of the model also support this conclusion.



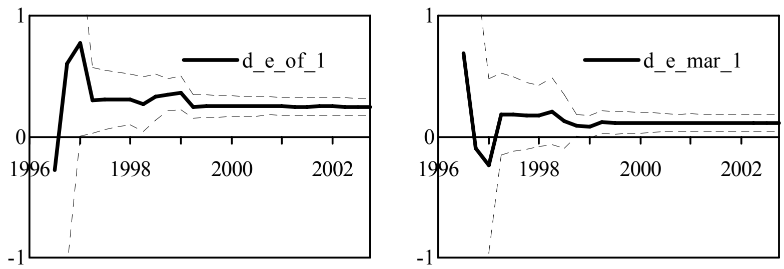


Figure 4. Recursive coefficients of model (9)

Econometric model of the producer price index

To develop a model of the producer price index (PPI), two cointegration relationships were also derived under this research which reflects the long-run factors setting the level of prices for industrial products.

By analogy with equation (8) for the GDP deflator, the first relationship characterizes the long-run level of the industrial costs adjusted for the markup. The main factors of the industrial costs are as follows: faster growth of the average monthly nominal wages in industry compared with the growth of real industrial production, price increases in the fuel industry, the devaluation of the market nominal exchange rate of the Belarusian ruble against the US dollar and increasing industrial products prices abroad. Thus, this relationship for the PPI will have the following form:

$$ipp = \alpha_0 + \alpha_1 w^{IND} + \alpha_2 y^{IND} + \alpha_3 ipp^{FUEL} + \alpha_4 \left(e^{MAR} + ipp^{USA}\right) + \varepsilon, \qquad (10)$$

where ipp is a logarithm of the producer price index, calculated on the December 1993 time base;

w^{IND} is a logarithm of the average quarterly monthly industrial wages and salaries;

y^{IND} is a logarithm of the real volume of industrial production;

ipp^{FUEL} is a logarithm of the price level in the fuel industry;

e^{MAR} is a logarithm of the average quarterly market nominal exchange rate of the Belarusian ruble against the US dollar;

ipp^{USA} is a logarithm of the producer price index in the USA;

ε is a random error, characterizing the temporal deviation of the markup value in industry from its long-run level.

Dynamics of model (10) indicators is shown in Figure 5.

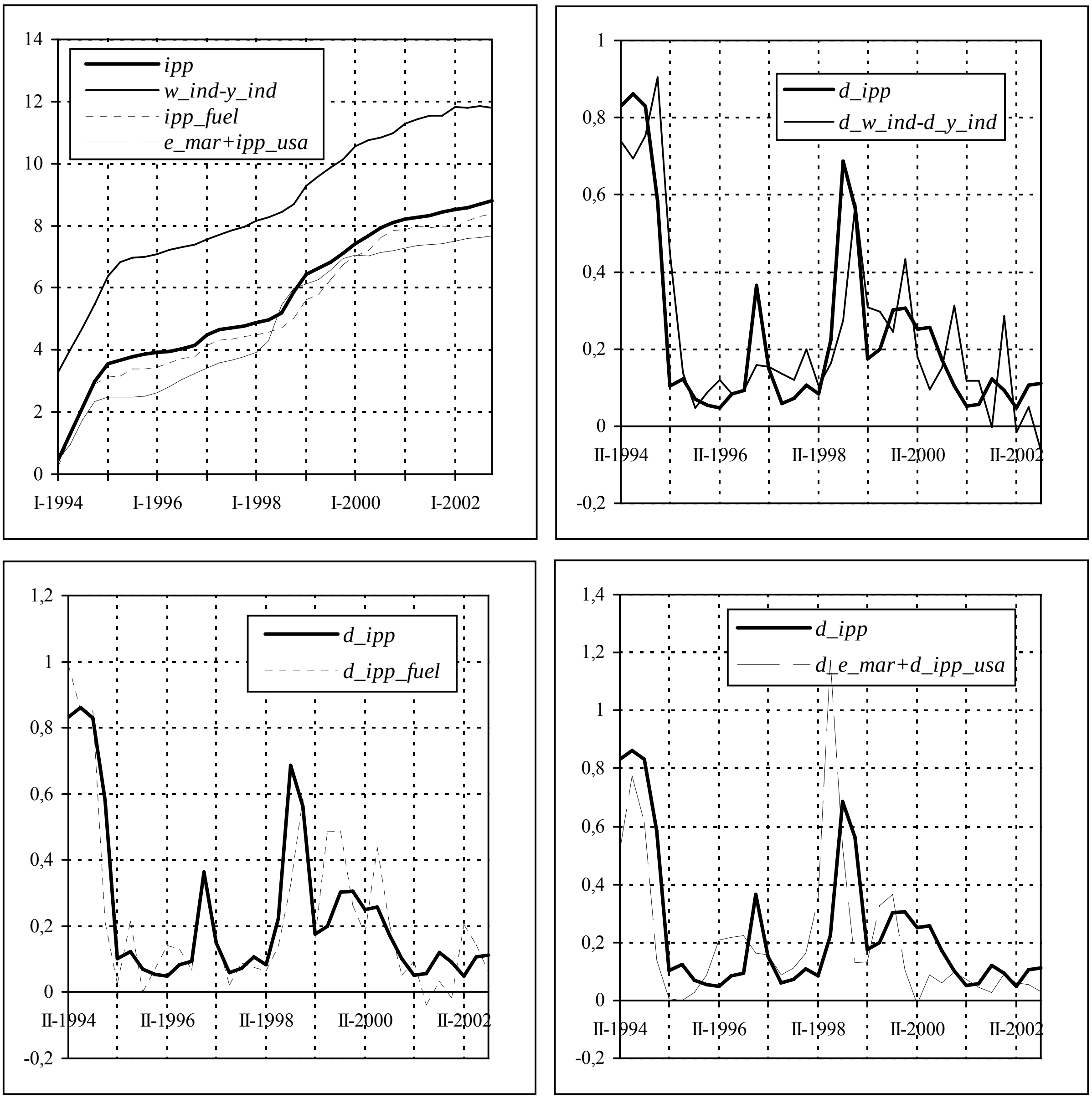


Figure 5. Dynamics of model (10) indicators

The Johansen cointegration analysis was used for estimating long-run relationship (10). Analysis of the VAR-model lag structure on the basis of a number of sequential Lagrange multipliers tests (LMF) as well as the use of the Akaike information criterion (AIC) and the Schwarz information criterion (SIC) showed that there is no autocorrelation in residuals for the VAR(4) model.

Results of the Johansen $\lambda_{\max}$ and λ_{trace} cointegration tests are shown in Table 8, whose data confirm the existence of three cointegration vectors. Only the first cointegration relationship has theoretically correct parameters under the relevant normalization:

$$ipp = -1.645 + 0.601w^{IND} - 0.479y^{IND} + 0.297ipp^{FUEL} + 0.096\left(e^{MAR} + ipp^{USA}\right) + error. \qquad (11)$$

Table 8

Results of the cointegration analysis of model (10)

(A) Determining the cointegration rank

Eigenvalue	0.918	0.832	0.540	0.200	0.061
Null hypothesis	$r=0$	$r \le 1$	$r \le 2$	$r \le 3$	$r \le 4$
$\lambda_{\max}$	79.947**	57.132**	24.815**	7.144	2.018
λ_{trace}	171.056**	91.109**	33.977**	9.162	2.018

(B) Standardized cointegration vectors β'

ipp	w^{IND}	y^{IND}	ipp^{FUEL}	$e^{MAR}+ipp^{USA}$	const
1	-0.601	0.479	-0.297	-0.096	1.645
-1.100	1	-1.504	-0.126	0.263	-3.043
0.188	-0.225	1	-0.144	0.077	1.136

(C) Standardized adjustment coefficients α

ipp	-0.968	0.651	-0.310
w^{IND}	0.084	0.207	0.076
y^{IND}	-0.229	0.295	0.190
ipp^{FUEL}	0.237	1.858	0.310
$e^{MAR}+ipp^{USA}$	-0.933	-0.029	4.165

On the basis of the results of cointegration analysis the following conclusions could be drawn. The level of industrial prices depends on the growth of the nominal wages to a lesser degree than the GDP deflator: the elasticity with respect to the variable w^{IND} turned out to be equal to 0.601 as compared to 0.745 with respect to the variable w in equation (8). Indeed, the share of labour remuneration costs in production costs is lower in industry than in agriculture, construction and services sectors, whose product prices are captured by the GDP deflator. At the same time, the level of industrial prices is to a greater extent influenced by the dynamics of energy inputs and intermediate import prices: the long-run elasticity coefficient amounted to 0.297 with respect to the variable ipp^{FUEL} and 0.096 with respect to the variable $e^{MAR} + ipp^{USA}$ (0.195 and 0.082, respectively, in the equation for GDP the deflator).

The signs of relationship (11) parameters estimates coincide with theoretically assumed ones. The homogeneity condition $\alpha_1 + \alpha_3 + \alpha_4 = 1$ is satisfied as well, the actual total of the elasticities of the industrial price level with respect to the main cost factors amounted to 0.994. As distinct from the analogous relationship for the GDP deflator, the restriction $\alpha_1 = -\alpha_2$ is also acceptable for this equation. Indeed, the absolute value of the elasticity with respect to the wages (0.601) is close to the absolute value of the elasticity with respect to the real industrial production (-0.479).

Interestingly, the dynamics of deviations of the actual level of industrial prices from the long-run level determined by relationship (11) is practically reproducing the dynamics of profitability rate in industry (in the logarithmical form). That is evidenced by Figure 6. So, this fact proves the economic content of cointegration relationship (11).

The second relationship that was developed in the PPI model is based on the well-known concept of the relative purchasing power parity (PPP). Since industrial products are traded internationally (tradable goods), the dynamics of prices for domestic and foreign goods must be close. Otherwise, the international arbitrage mechanism would be switched on and, ultimately, prices would level off. To examine the relative PPP law, the following variable was introduced:

$$rer = ipp - e^{MAR} - ipp^{USA}, \tag{12}$$

where rer is the real exchange rate;
ipp is the PPI in Belarus, calculated on the December 1993 time base;
ipp^{USA} is the PPI in the USA, calculated on the December 1993 time base;
e^{MAR} is the market nominal exchange rate of the Belarusian ruble against the US dollar.

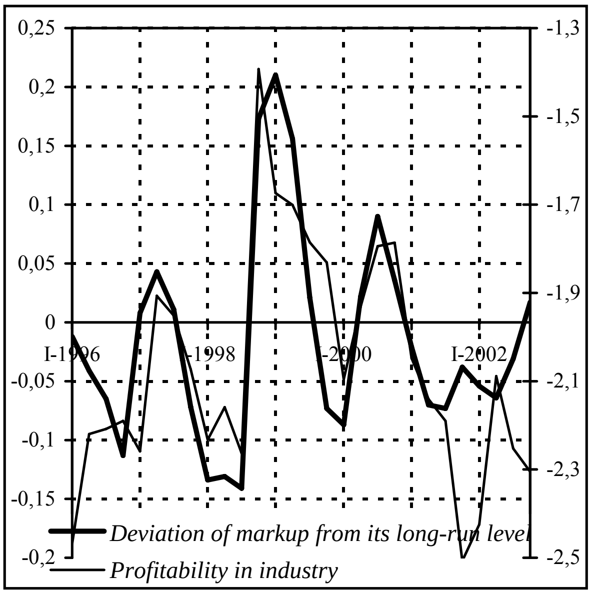


Figure 6. Profitability in industry and residuals of equation (11)

For the PPP law to be valid for the Republic of Belarus, the variable rer must be stationary. Thus, unilateral deviations of the real exchange rate from its long-run constant level can be only temporal. The actual dynamics of this variable is shown in Figure 7.

The variable rer was tested for stationarity on the basis of the augmented Dickey-Fuller test (ADF). The results are shown in Table 9. As analysis demonstrated, the real exchange rate is not a stationary variable and is integrated of order 1. At the same time, given a low power of this criterion and a relatively short period of analysis, we will assume that the long-run cointegration relationship for the variables ipp, e^{MAR} and ipp^{USA} exists. Moreover, two cases of a sharp devaluation of the nominal exchange rate of the Belarusian ruble in 1994 and 1998 prove that a significant depreciation of the real exchange rate is followed by its subsequent appreciation up to the level preceding the currency crisis, that is the levelling off of prices for domestic and imported industrial products is observed.

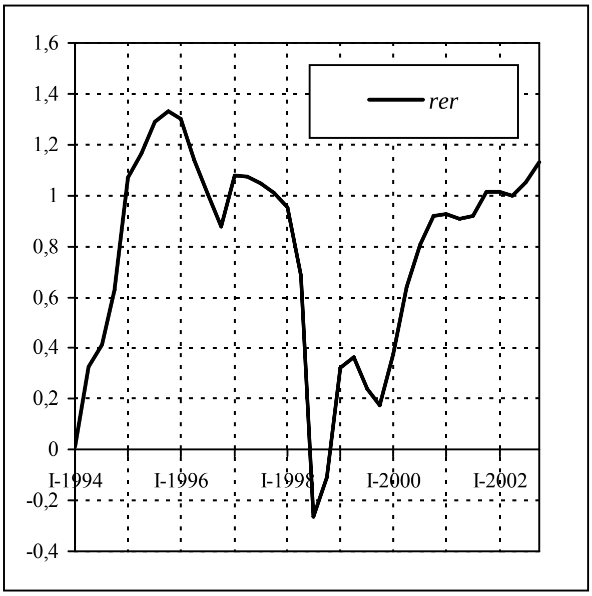


Figure 7. Dynamics of the real exchange rate in the Republic of Belarus

By analogy with the GDP deflator model, derived cointegration relationships (11) and (12) were used for developing a short-run model of the PPI. To this end, the following error-correction model was examined:

$$\Delta ipp = \pi_0 + \sum_{s=1}^{3} \pi_s seas_s + \alpha_1 ecm\_ipp_{-1}^{MARKUP} + \alpha_2 ecm\_ipp_{-1}^{PPP} + \sum_i \beta_i \Delta ipp_{-i} + \sum_j \Gamma_j^T \Delta X_{-j} + \upsilon, \quad (13)$$

where $X = \left(w^{IND}, y^{IND}, ipp^{FUEL}, e^{OF}, e^{MAR} + ipp^{USA}\right)^T$;

$ecm\_ipp^{MARKUP} = ipp - 0{,}601 w^{IND} + 0{,}479 y^{IND} - 0{,}297 ipp^{FUEL} - 0{,}096\left(e^{MAR} + ipp^{USA}\right) + 1{,}645$;

$ecm\_ipp^{PPP} = ipp - e^{MAR} - ipp^{USA}$;

$seas_s$ is a seasonal quarterly dummy variable that is equal to 1 for the quarter s and 0 for other quarters;

υ is a random error.

Table 9

Results of real exchange rate testing for stationarity:
the augmented Dickey-Fuller test (ADF)

Variable	t_{ADF}	Test	Variable	t_{ADF}	Test
rer	-2.346	ADF(1) with a constant	Δrer	-4.385**	DF

Equation (13) was estimated on the basis of statistical data for the period Q1 1994 – QIV 2002 using the least squares method. A number of parameters tests made it possible to ignore insignificant factors and to derive a short-run PPI model. Estimates of the model parameters, statistical characteristics and the results of residual tests are shown in Table 10.

Table 10

Results of model (13) estimation

Variable	Estimates	t-statistics
$ecm\_ipp_{-1}^{MARKUP}$	-0.490	-4,446**
Δipp_{-1}	0.485	7,534**
Δipp^{FUEL}	0.344	7,610**
Δe_{-1}^{MAR}	0.283	6,437**
$seas_2$	-0.081	-4,750**
R^2	0.970	
$\overline{R}^2$	0.966	
$s.e.$	0.041	
DW	2.339	
$AR:\chi^2(4)$	4.180	
$ARCH:\chi^2(4)$	12.334**	
$Jarque-Bera:\chi^2(2)$	0.633	
$White:\chi^2(4)$	12.908	

The model has adequate statistical characteristics and meets a number of criteria for properties of residuals. The Lagrange multipliers tests (LM) provide evidence that there is no autocorrelation and the White's heteroskedasticity. The Jarque-Bera test confirms normality of residuals. At the same time, the null hypothesis of no autoregressive conditional heteroskedasticity is rejected and a more thorough analysis proves an insignificant presence of the ARCH(1)-effect. The determination coefficient R^2 for the model with included insignificant constant is equal to 0.972, whereas $\overline{R^2}$ is 0.966 which characterizes high explanatory capability of the model. The factor $ecm\_ipp_{-1}^{PPP}$ was found to be insignificant, and the rate of PPI adjustment to the long-run level of production costs amounted to -0.49.

The model in question allows us to forecast the dynamics of industrial prices with sufficient accuracy. For instance, Figure 8 shows dynamics of the actual PPI during the period 1994–2002, its forecast made on the basis of the present model as well the forecast errors. Note that the right-hand chart describes dynamics of the forecast for the model estimated on data for the period 1994–2000. Thus, the right-hand chart for the period 2001–2002 illustrates the perspective prediction under known explanatory variables. Both charts confirm that the actual inflation measured by the PPI is within the 95% confidence interval for the central forecasting value.

Figure 9 shows the dynamics of the recursive estimates of model (13) coefficients. As we can see, the model coefficients are stable to changes in the sample size for the estimation which makes it possible to conclude that the model is stable.

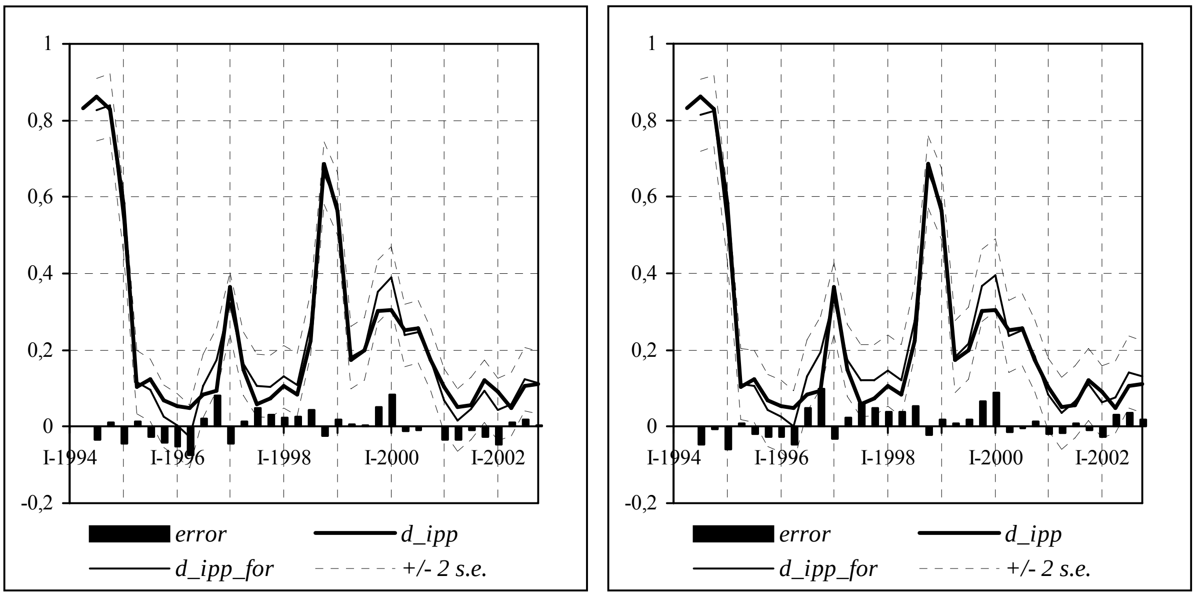


Figure 8. Forecasts of the PPI changes on basis of model (13)

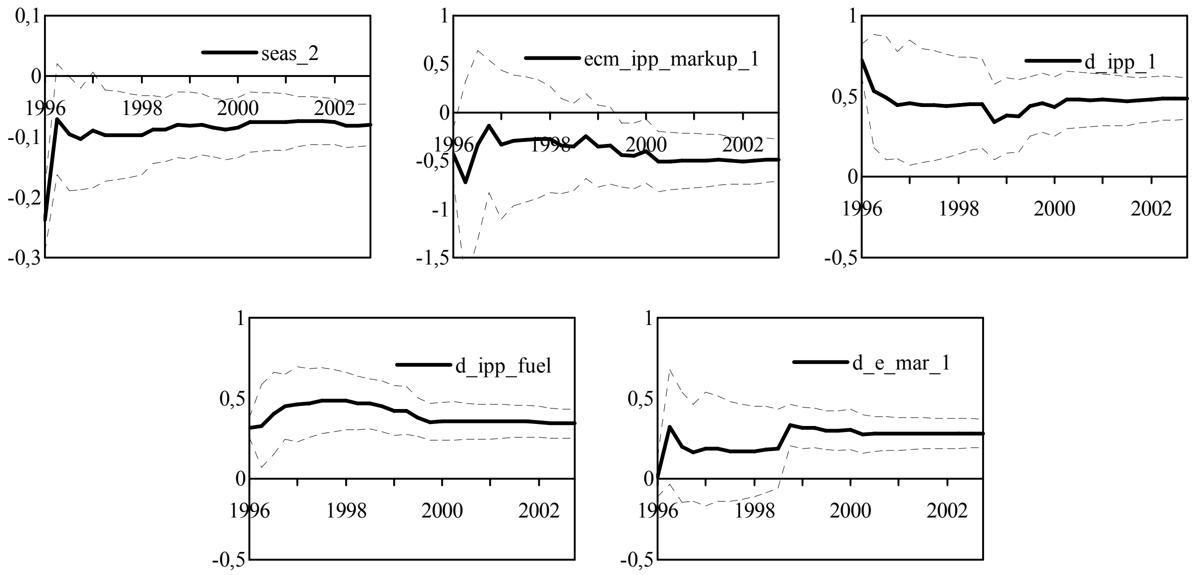


Figure 9. Recursive coefficients for model (13)

Econometric model of the consumer price index

In contrast to the GDP deflator and producer price index models discussed in the previous sections, the long-run relationships included into the consumer prices index model do not reflect theoretical interrelations between inflation and its main factors from the demand or cost side. Consumer prices dynamics is determined by changes in some other price indicators taken with certain weights.

The first relationship reflects the structure of the consumer basket used for calculations of the consumer price index. As discussed earlier, the dynamics of consumer services tariffs in the present model is considered to be exogenously assigned since the state regulates tariffs on housing and public utilities, passenger transportation and communication services to a considerable degree. At the same time, prices for consumer goods traded internationally are subject to administrative regulation to a lesser degree and, on the whole, they are formed under the impact of market forces. That is why the dynamics of the consumer goods price level in the present model is determined endogenously. Thus, the general level of the consumer prices in the present model is determined by the following relationship:

$$cpi = \alpha_1 cpi^G + \alpha_2 cpi^S \,,\;\; \alpha_1 + \alpha_2 = 1, \tag{14}$$

where cpi, cpi^G and cpi^S are, respectively, logarithms of the indexes of consumer prices, prices for consumer goods and tariffs on consumer services, calculated on the December 1993 time base.

Model (14) indexes are shown in Figure 10 which evidences that the dynamics of the consumer price level during the period 1994-2002 was basically determined by the changes in the consumer goods prices which reflects their prevailing weight in the consumer basket.

As before, the Johansen cointegration analysis was applied for the estimation of structural relationship (14). Putting the VAR-model to a number of the Lagrange multipliers tests (LMF) and using the Akaike information criterion (AIC) and the Schwarz information criterion (SIC) made it possible to ascertain that the maximum length of the lag must be equal to 1.

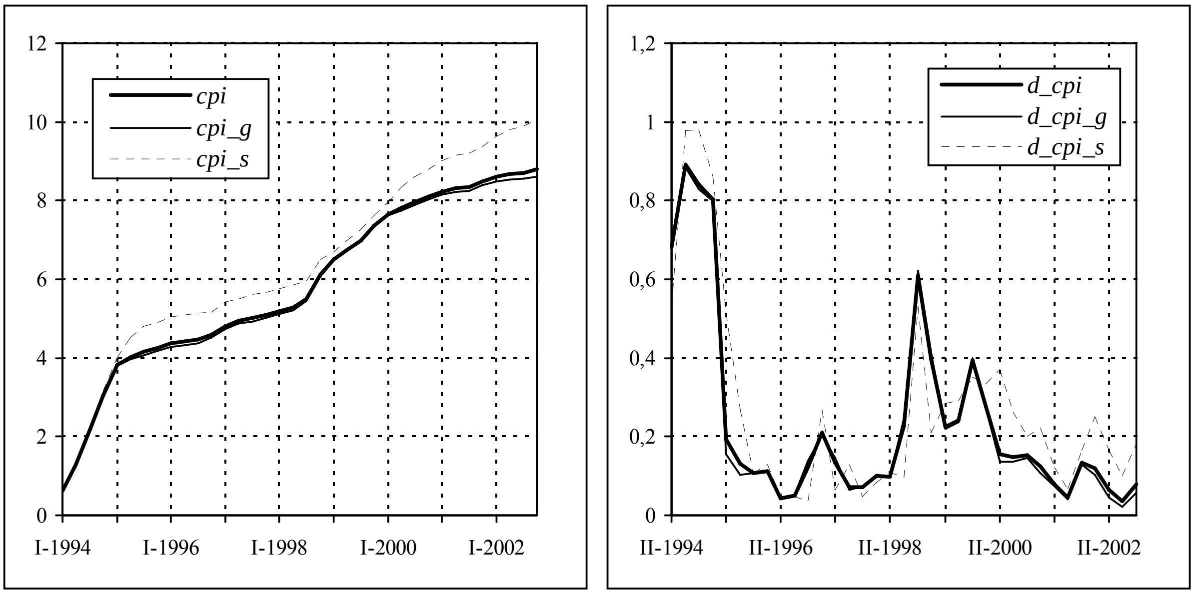


Figure 10. Dynamics of model (14) indicators

Results of the Johansen $\lambda_{\max}$ and λ_{trace} tests are shown in Table 11 and they confirm the presence of two cointegration vectors. The homogeneity condition $\alpha_1 + \alpha_2 = 1$ is satisfied for the second cointegration relationship (after appropriate normalization) only and thus equation (14) would is as follows:

$$cpi = 0{,}850 cpi^G + 0{,}147 cpi^S + error. \tag{15}$$

As the analysis shows, the share of consumer services in the consumer basket is insignificant and on the average makes up 15%, whereas the overwhelming part of the CPI dynamics is determined by changes in consumer goods prices – 85%.

To find out the dynamics of the consumer goods prices, one more structural relationship was developed taking into account changes in prices both for domestic and imported goods. This relationship is as follows:

$$cpi^G = \beta_0 + \beta_1 py + \beta_2 \left(e^{MAR} + ipp^{USA}\right) + \varepsilon, \tag{16}$$

where py is a logarithm of the GDP deflator which characterizes the level of prices for domestic goods;

$e^{MAR} + ipp^{USA}$ is a logarithm of the imported goods price level determined by the level of prices abroad ipp^{USA} and by the market nominal exchange rate e^{MAR}.

Table 11

Results of cointegration analysis of model (14)

	(A) Determining the cointegration rank		
Eigenvalue	0.484	0.285	0.021
Null hypothesis	$r = 0$	$r \le 1$	$r \le 2$
$\lambda_{\max}$	23.187**	11.742**	0.743
λ_{trace}	35.673**	12.486**	0.743
	(B) Standardized cointegration vectors β'		
cpi	cpi^G	cpi^S	
1	-0.740	-0.238	
-1.176	1	0.172	
	(C) Standardized adjustment coefficients α		
cpi	3.443	-0.708	
cpi^G	3.357	-0.957	
cpi^S	4.181	0.012	

Dynamics of the model variables is shown in Figure 11.

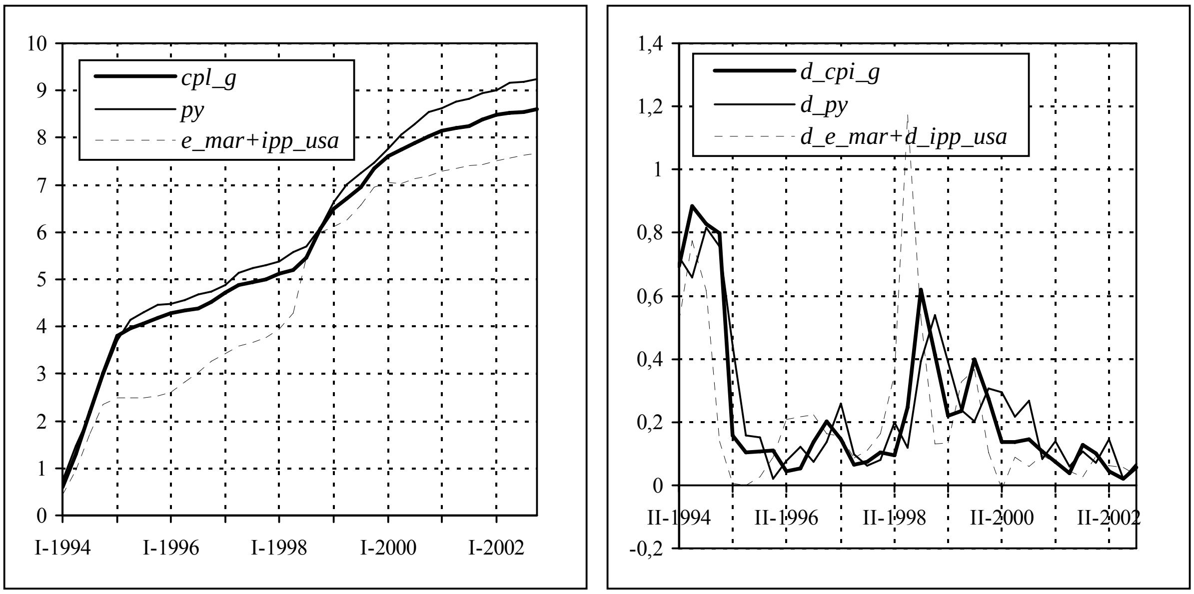


Figure 11. Dynamics of model (16) indicators

The results of cointegration analysis for the given relationship are shown in Table 12. The VAR-model with the maximum length of lag equal to three was selected. According to the data of the table, the Johansen $\lambda_{\max}$ and λ_{trace} tests showed presence of a single cointegration relationship which has economically sound values of parameters:

$$cpi^{G} = 0{,}883 + 0{,}495\,py + 0{,}395\left(e^{MAR} + ipp^{USA}\right) + error. \tag{17}$$

According to this equation, 50% of the consumer price index long-term dynamics is determined by changes in the domestic goods prices and 40% by changes in the consumer imports prices. At the same time, the sum of these elasticities is less than 1 testifying that in the Republic of Belarus the consumer prices growth rate is falling behind the growth rate of the GDP deflator and the nominal exchange rate. Probably this is due to the fact that the domestic consumer goods prices are subject to administrative regulation to a much greater extent than prices for capital goods and exports (these are accounted for in the GDP deflator) and prices for imports (these depend upon the exchange rate dynamics).

Table 12

Results of cointegration analysis of model (16)

(A) Determining the cointegration rank			
Eigenvalue	0.523	0.162	0.091
Null hypothesis	$r=0$	$r\le 1$	$r\le 2$
$\lambda_{\max}$	24.440**	5.834	3.140
λ_{trace}	33.414**	8.974	3.140
(B) Standardized cointegration vectors β'			
cpi^G	py	$e^{MAR}+ipp^{USA}$	const
1	-0.495	-0.395	-0.883
(C) Standardized adjustment coefficients α			
cpi^G	-0.658		
py	-0.301		
$e^{MAR}+ipp^{USA}$	-1.698		

In order to estimate the consumer price index short-term dynamics, the following error-correction model was considered:

$$\Delta cpi^G = \pi_0 + \sum_{s=1}^{3}\pi_s seas_s + \alpha ecm\_cpi_{-1}^G + \sum_i \beta_i \Delta cpi_{-i}^G + \sum_j \Gamma_j^T \Delta X_{-j} + \upsilon, \tag{18}$$

where $X=\left(py, e^{MAR}+ipp^{USA}\right)^T$;

$ecm\_cpi^G = cpi^G - 0{,}495\,py - 0{,}395\left(e^{MAR}+ipp^{USA}\right) - 0{,}883$;

$seas_s$ is a seasonal quarterly dummy variable equal to 1 for the quarter s and 0 for other quarters;

υ is a random error.

Model (18) was estimated on the basis of quarterly statistical data for the period 1994-2002 by means of the least squares method. Some insignificant factors were ignored and the CPI model was developed following a number of parameters tests. The estimates of the model parameters, statistical characteristics and results of the residual tests are shown in Table 13.

Table 13

Results of model (18) estimation

Variable	Estimates	t-statistics
$ecm\_cpi^G_{-1}$	-0.632	-5.118**
Δpy	0.622	7.755**
Δe^{MAR}_{-1}	0.127	1.850*
Δe^{MAR}_{-2}	-0.189	-3.510**
$seas_1$	0.073	3.726**
$const$	0.092	3.651**
R^2	0.958	
$\overline{R^2}$	0.950	
$s.e.$	0.045	
DW	2.042	
$AR:\chi^2(4)$	1.572	
$ARCH:\chi^2(4)$	1.324	
$Jarque-Bera:\chi^2(2)$	3.277	
$White:\chi^2(4)$	13.616	

According to the data of this table, the model thus developed possesses adequate statistical characteristics and satisfies a number of criteria for the residual properties. The Lagrange multipliers tests (LMF) confirm that there is no autocorrelation, autoregressive conditional heteroskedasticity and the White heteroskedasticity in residuals, while the Jarque-Bera test proves their normality. Coefficients R^2 and $\overline{R^2}$ are equal to 0.958 and 0.950, and although these are somewhat lower compared to the GDP deflator and PPI models, they nevertheless characterize high explanatory capability of the model. The coefficient of adjustment of deviations in the level of the consumer goods prices from its long-term level makes up -0.632.

The forecasting capability of the developed model is sufficiently good as well. Figure 12 shows the actual price index dynamics for the consumer goods in the period 1994-2002, its forecast data on the basis of this model as well as forecast errors. In this case, as before, the right-hand chart again shows the forecast dynamics on the basis of the model estimated on data for a shorter period (1994–2000). Both charts confirm that the actual price index is found within the 95% confidence interval for the central forecasting value.

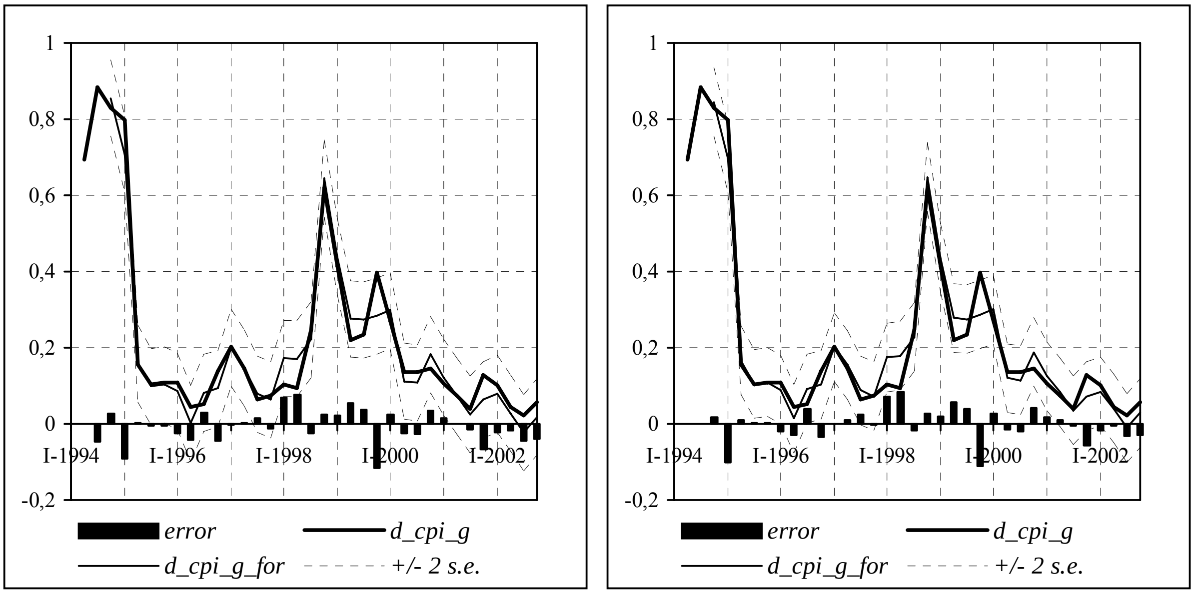


Figure 12. Forecasts on the basis of model (18)

Figure 13 shows the dynamics of the recursive estimates of model (18) coefficients. As could be seen from these charts, the model coefficients are stable to changes in the sample size for the estimation. Stability of the model is confirmed by the results of various kinds of the Chow tests as well.

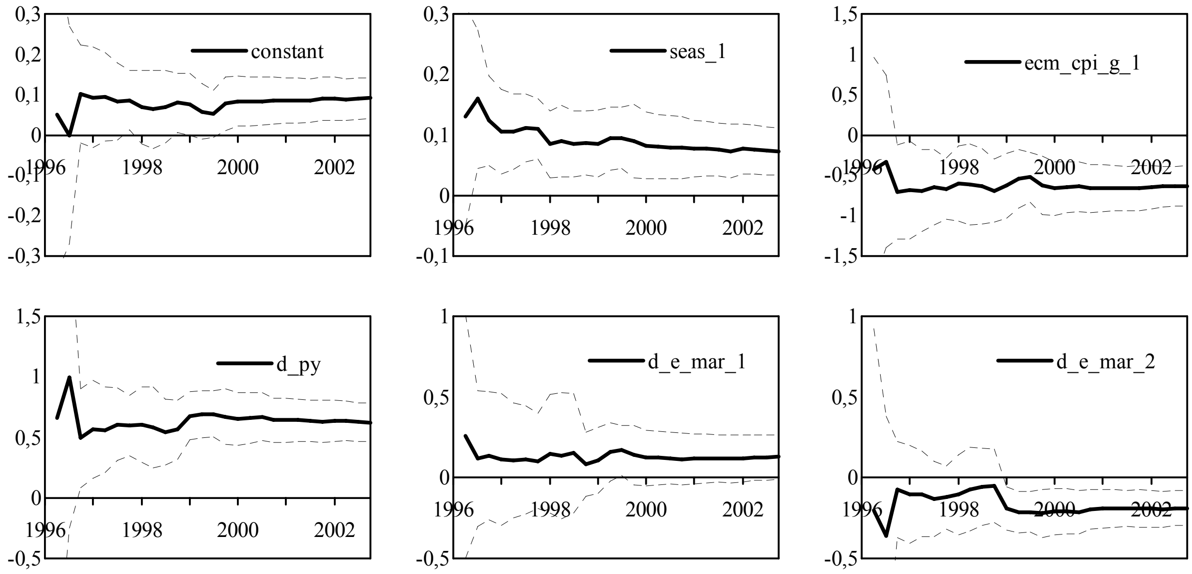


Figure 13. Recursive coefficients for model (18)

* * *

Thus, a model of inflation processes in the Republic of Belarus was developed which consists of equations (9), (13), (15) and (18), and can be used for forecasting the GDP deflator, producer price index and consumer price index. The model presented in this research paper possesses adequate statistical characteristics, demonstrates stability of the coefficients and provides a possibility to analyse various choices in the field of monetary and currency regulations as well as in the area of labour remuneration, prices and tariffs.