Zagreb International Review of Economics & Business, Special Conference Issue, pp. 1-12, 2004

ISSN 1331-5609; UDC: 33+65

Constructing an Export Supply Function for Croatia

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Abstract: The main aim of this paper is an investigation of the interplay between Croatian export supply and the real exchange rate, i.e. the relative prices. We analyse here the price elasticity of export supply through modelling both the short- and the long-run structures of the model. Additionally, the relationship between export supply and production capacity is also subjected to an analysis. As a result, this paper provides some insights into the importance played by the real exchange rate and production capacity in explaining Croatian exports.

JEL Classification: F4, F41

Key words: export supply, unit root tests, cointegration, VECM model

Introduction

Exporting firms are free to supply their goods and services either on domestic market or abroad. They allocate their output on both markets according to the price signals received. The relative profitability is therefore defined as the ratio between the average price received on export markets and the one received on the domestic market. A rise in the relative profitability of exporting leads to an increase in exports. Therefore, our special attention is focused on the interplay between Croatian export supply and the real exchange rate, i.e. of relative prices. The price elasticity of export supply is to be analysed through modelling both the short- and long run structures. Additionally, the relationship between export supply and production capacity (business cycle variable) is also subjected to the analysis. Therefore, real exchange rate and production capacity variables are integrated into a defined trade model. A long-term model serves as a benchmark where lagged reactions do not exist.

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Obtaining long-run estimates of the cointegration relationship is only a first step towards estimating the complete model. The short-run structure of the model is also important in the terms of the information it conveys on the short adjustment behaviour of economic variables. The dynamic modelling enables us to analyse whether an adjustment with lags is characteristic for the total exports in the short-run, especially when we have reasons to assume that export reacts and responds to changes in the real exchange rate.

Theoretically currency devaluation can improve export, i.e. trade flows, if the relative prices among the country and its trading partners, as well as other factors, remain unchanged. Whether devaluation improves the trade flow is still unclear as shown by many empirical studies. Changes in the real exchange rate do affect the trade flows in some economies but not in all because the changes in nominal exchange rate might lead to changes in the relative prices in the same or different directions. In other words, when country has the exchange rate changes the real exchange rate may capture two effects, price effect and volume effect.

The next section presents data used in the analysis, unit root tests of the variables and a definition of export supply model. It continues with a cointegration analysis of Croatian data. The third section concludes.

Estimation: Data, Model and Results

The model estimation is based on monthly data covering the period[1] 1994/1-2002/12. In order to estimate export demand function, export supply (*exp*) is considered as a function of real domestic income (*ind*) and real effective exchange rate (*reft*). Letting lower-case letters denote logarithmic values of time series.

The real effective exchange rate in forms of indices is taken from International Financial Statistics (IFS). The real effective exchange index is derived from the nominal effective exchange rate index, adjusted for relative changes in consumer prices. A nominal effective exchange rate index (base period 1995=100) is a ratio of an index of a currency's period-average exchange rate to a weighted geometric average of exchange rate for the currencies of the selected countries and Euro area. A real effective exchange rate index represents a nominal effective exchange rate index adjusted for relative movements in national price or cost indicators of home country, selected countries, and Euro area. The average exchange rate expressed in terms of US$ of the national currencies is given as an index number based on 1995=100. In both cases, an increase in the index number reflects the appreciation.

So far, data on real export do not exist in Croatian statistics. Data on export prices and deflator of exports of goods and services are not available on monthly basis. Therefore, nominal export of goods and services in national currency from the

Croatian Bureau of Statistics (CBS) reports are taken as the best proxy for the overall export volume of the Croatia and are used in the estimation (e.g. Straus, 2001). Industrial production is a representative proxy for total economic activity (e.g. Filardo, 1997).

The empirical analysis is presented in two parts. The first part presents the estimation of the long-run equilibrium of the variables — the cointegration vector(s). In the second part this information is included into a model of short-run dynamics as an error-correction term.

A number of alternative ways to analyse integration and cointegration of time series as well as to estimate the cointegrating vectors have been proposed (e.g. Maddala and Kim, 1998, Harris, 1995 etc.). We employed Johansen's reduced-rank procedure, (e.g. Johansen, 1988, and Johansen and Juselius, 1990). Before testing for cointegration, the order of integration of the individual time-series must be determined.

Testing for Unit Roots

A set of unit root tests was performed on data attempting to classify the series based upon trend and unit root properties. To save the space, only the results of ADF tests (e.g. Dickey-Fuller, 1979) and KPSS tests (e.g. Kwiatkowski et al., 1992) are presented in Table 1.

The difference between these two types of tests is the specification of the null hypothesis. ADF test has a nonstationarity as a null hypothesis i.e. the null hypothesis is that the variable under investigation has a unit root. On the other hand, in the KPSS test it is assumed that the variable is stationary. It has been suggested that the tests using stationarity as a null can be used for confirmatory analysis, i.e. to confirm the conclusion about the order of integration suggested by other unit root tests, (e.g. Kwiatkowski et al., 1992). If both tests fail to reject the respective nulls or both reject the respective nulls, there is a confirmation[2]. The top part of Table1 reports tests of stationarity of the log-levels of the variables and the bottom half of their first differences[3]. The variables used in this study are given in the first column. Columns two, three and four contain test values for ADF tests with the information about adding a constant term or/and a deterministic trend to the model. The strategy of adding lags to the ADF regression is based on the objective to remove any autocorrelation from the residuals, which is tested applying Lagrange Multiplier test[5]. For each test the length of included lags is given in the square brackets after the test value. The fifth column contains KPSS test values for testing stationarity around level and in the sixth column KPSS test values for testing trend stationarity of the variables are reported[6].

Table 1.: Variables and Unit Root Tests

a) Levels

Variable	ADF value	ADF value Constant included	ADF value Constant and trend included	KPSS value H0 stationary around a level	KPSS value H0 trend stationary
Exp	1.0201(4)	-1.2647(3)	-3.6902* (2)	2.3742**	0.3087**
Ind	2.9328(12)	0.7642(12)	-1.7937(13)	0.9165**	0.0773**
Reft	0.1569(0)	-3.5653*(0)	-4.1614* (0)	2.4312**	0.6646**

b) First Differences

First Differences	ADF value	ADF value Constant included	ADF value Constant and trend included	KPPS value H0 stationary around a level	KPPS value H0 trend stationary
Δexp	-13.4273** (1)	-8.3377** (3)	-8.2865** (3)	0.0265	0,0263
Δind	-2.5962* (12)	-12.4349** (0)	-12.3768** (0)	0.0080	0.0068
Δreft	-14.1595** (0)	-14.0974** (0)	-14.0428** (0)	0.0210	0.0155

Notes: Δ is the first difference operator. One (two) asterisk(s) indicates a rejection of the Null at 5% (1%) significance level. The critical values for ADF tests are taken from Hamilton (1994) and for KPSS tests from Kwiatkowski and al. (1992).

Test results suggest that the unit root null cannot be rejected for the levels of *ind* and *exp* variables[7]. In the case of real effective rate (*reft*) variable the unit root hypothesis is rejected at 5% significance level if the model includes constant or trend (at 1% significance level unit root null is accepted). If the criteria for adding lags to the ADF regression is changed to AIC or BIC the null of the unit root for *reft* cannot be rejected at 5% level. In contrast to level forms, the unit root null is strongly rejected at 5% significance for the first differences of all variables. The results of KPSS tests confirm the results of ADF tests. In other words, all variables appear to be integrated of order one and in the remainder of this study all variables are treated as being I(1).

Modelling Export Supply Function

When analysing export supply in a multivariate framework we define a vector of variables $Z_t = (exp, reft, ind)$ and allow all three variables in Z_t to be potentially endogenous, i.e. we define the following unrestricted VAR model:

$$Z_t = A_1 Z_{t-1} + \ldots + A_k Z_{t-k} + \mu' + \mu \cdot t + \Psi D_t + u_t, \quad u_t \approx IN(0, \Sigma), \tag{1}$$

where D_t is a vector of non-stochastic seasonal dummy variables included in the model. Model (1) can be reformulated into a vector error-correction model, VECM model:

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \Pi Z_{t-1} + \mu + \Psi D_t + u_t \qquad (2)$$

where: Ψ and Γ_i are matrices of parameters[8]. Matrix Π contains information on the long-run relationships; in fact $\Pi = \alpha\beta'$ where β is 3 by r matrix of long-run coefficients (cointegration vectors) and α is 3 by r matrix of the respective loading of cointegrating vectors and represents the speed of adjustment towards the long-run equilibrium. r is a number of cointegrating vectors of the system, and k is a lag length of the VAR model[9].

The analysis starts with determination of a number of cointegrating vectors, r. The results are presented in Table 2.

Table 2.: Johansen's Test for the Number of Cointegrating Vectors[10]

H_0: r =	p-r	λ_{max}	λ_{trace}	λ	λ_{max} - 10% critical value	λ_{trace} - 10% critical value
0	3	17.69*	25.97	0.1550	13.39	26.70
1	2	8.16	8.29	0.0748	10.60	13.31
2	1	0.12	0.12	0.0012	2.71	2,71

BETA (transposed)

exp	reft	ind
1.000	5.737	-3.464

	ALPHA	T-VALUES FOR ALPHA
Δexp	-0.256	-3.847
Δreft	-0.019	-1.849
Δin per cent	0.001	0.066

Note: '*' indicates a rejection of the Null at 10 per cent. The critical values are from Johansen and Nielsen, (1993).

From the results in Table 2 it can be concluded that there is one cointegrating vector, $\hat{\beta}_1$. The outcome of the cointegration analysis is the following long-run relationship:

$$exp_t = -5.737\, reft_t + 3.464\, ind_t \qquad (3)$$

The diagnostic statistics of the model (Table 3) is generally satisfactory except for the variable real effective exchange rate. When modelling *reft* in multivariate settings there is a violation of normality and heteroscedascity that affects the whole system as well. Introducing additional dummy variables can solve the problem of heteroscedascity. On the other hand, non-normality in *reft* is not such a problem if variable prove to be weakly exogenous, (e.g. Johansen and Juselius, 1992). In that situation we analyse a partial system with variable *reft* treated as weakly exogenous[11] and thus improve stochastic properties of the model.

Table 3.: Residual Analysis of the VECM Model[12]

TEST FOR AUTOCORRELATION				TEST FOR NORMALITY		
L-B(26),CHISQ(213)=237.959, p-val = 0.12						
LM(1), CHISQ(9) = 5.124, p-val = 0.82				CHISQ(6)=126.183,p-val=0.00		
LM(4), CHISQ(9) = 5.892, p-val = 0.75						
UNIVARIATE STATISTICS						
VARIABLE	STD.DEV	SKEWNESS	KURTOSIS	ARCH(3)	Normality	R-squared
exp	0.104454	0.414624	3.273736	6.956	3.208	0.692
reft	0.016087	0.868498	13.462776	22.085	121.849	0.280
ind	0.033120	0.194632	2.427514	6.027	2.296	0.750

Using our findings of one cointegrating vector, we continue the analysis with imposing restrictions on weak exogeneity of variables[13], *reft* and *ind*. Performing likelihood ratio tests of the adjustment parameters involving the eigenvalues of the system, we can not reject the null of weak exogeneity for both *reft* and *ind*.

Table 4.: Testing Weak Exogeneity of Variables *reft* and *ind*

Weakly exogenous	reft	Ind	reft and ind
Restrictions on $\hat{\alpha}_1'$	$\hat{\alpha}_1' = (*,0,*)$	$\hat{\alpha}_1' = (*,*,0)$	$\hat{\alpha}_1' = (*,0,0)$
Test statistics	$\chi^2(1) = 1,87$	$\chi^2(1) = 0,00$	$\chi^2(2) = 2,20$
p-value	0.17	0.95	0.33

The next step is to estimate a dynamic short-run model. Treating variables *reft* and *ind* as weakly exogenous we analyse export supply function in a single equation model. Thus, model (2) can be reformulated into a conditional model[14] defined as:

$$\Delta \exp_t = a + \sum_{i=1}^{2} b_i \Delta \exp_{t-1} + \sum_{i=0}^{2} c_i \Delta reft_{t-1} + \sum_{i=0}^{2} d_i \Delta ind_{t-1} + \alpha ECM_{t-1} + \sum_{i=1}^{11} g_i SEAS_{t-1} + u_t \quad (4)$$

The estimation results of conditional model (4) are summarised in Table 5 and diagnostic statistics in Table 6.

Table 5.: Estimates of the Conditional Model

Variable	Coefficient	't-value'	Variable	Coefficient	't-value'
Constant	5.257	4.090			
Δexp_{t-1}	-0.449	-4.327	SEA(1)	-0.140	-1.957
Δexp_{t-2}	-0.204	-2.239	SEA(2)	-0.107	-1.542
$\Delta reft_{t}$	0.298	0.482	SEA(3)	-0.096	-1.693
$\Delta reft_{t-1}$	-0.296	-0.450	SEA(4)	-0.053	-0.854
$\Delta reft_{t-2}$	0.874	1.456	SEA(5)	-0.233	-3.431
Δind_{t}	0.773	2.540	SEA(6)	-0.113	-1.920
Δind_{t-1}	-0.698	-1.914	SEA(7)	-0.111	-2.102
Δind_{t-2}	-0.189	-0.583	SEA(8)	-0.180	-2.643
			SEA(9)	-0.177	-2.327
ECM_{t-1}	-0.326	-4.085	SEA(10)	-0.270	-3.290
			SEA(11)	-0.181	-2.455

The error correction term, *ECM*, is a difference between actual export and their long-run value as predicted by the cointegration relationship[15]:

$$exp_t = -3{,}306\, reft_t + 2{,}939\, ind_t \qquad (5)$$

The error-correction coefficient a equals -0.326 with t-value of -4.085. It presents a measure of the average speed of convergence towards the long-run equilibrium. It has an expected negative sign, implying that a deviation from long-run equilibrium exerts pressure on export adjusting 32.6 per cent of the resulting disequilibrium in each period.

Obtained results of the VECM model show that in the long-run export supply is affected by real domestic activity positively while real effective exchange rate has a negative impact. The long-run elasticity of the production capacity is 2.939 and related exchange rate elasticity is 3.306. Thus, the conclusion is that in the long-run the main determinants of Croatian export are domestic demand and real exchange rate. However, it should be pointed out that, as suggested by neo-classical theory, the real exchange rate is crucial for export supply. In each case the results show that domestic production capacity and real depreciation do all have right signs and significant (as expected) impact on export supply in the long-run[16]. In the short-run, apart from export itself with one and two period lags effects, domestic demand (that is proxied by industrial production) affects export supply. Its influence is positive in the current period and negative in the previous one. The influence of real effective exchange rate on export changes signs through time, though without statistical significance. The error-correction term, ECM_{t-1}, is statistically significant in the model, as well as a constant. The significance of the constant indicates an existence of a linear trend in the data. The residual statistics of conditional model is generally satisfactory.

Table 6: Residual Analysis of the Conditional Model

MULTIVARIATE STATISTICS TEST FOR AUTOCORRELATION	TEST FOR NORMALITY
L-B(26), CHISQ(23)=26.344, p-val = 0.28 LM(1), CHISQ(1) = 0.003, p-val = 0.96 LM(4), CHISQ(1) = 0.280, p-val = 0.60	CHISQ(2) = 5.859, p-val = 0.05

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VARIABLE	STD.DEV	SKEWNESS	KURTOSIS	ARCH(3)	Normality	R-squared
exp	0.100575	0.577158	3.514066	4.204	5.859	0.714

To check stability of the final one-equation model and constancy of cointegrating vector β we performed a sequence of statistical tests. We chose a sub-sample from 1994/1 to 2001/1 as a base period. When testing the constancy of β, the parameter $\hat{\beta}$ is first calculated for the base period. The constancy of the cointegrating space is then tested using a sequence of tests of the 'known vector' β, where the known vector is represented by chosen $\hat{\beta}$, sub-sample estimate have β.

First we perform the Trace test. When analysing the Trace statistic, Figure 1, one would expect the time path of Trace statistic to be upward sloping for $j \leq r$ and constant for $j > r$. Since our conditional model has rank 1 we expect the time path of Trace statistic to be upward sloping. Although it seems to be some changes by the end of the series the general impression is that the time path is indeed upward sloping. The maximised log-likelihood function, which consists of two factors, is given in Figure 2. The path for the log-likelihood value is well inside the 95 per cent confidence bounds for the full sample. Figure 3 shows a plot of the test of constancy of the cointegration space. The hypothesis is accepted in the investigated sub-sample. This supports the hypothesis of parameter constancy for the analysed period. Figure 4 gives plot of the time path of the non-zero eigenvalue together with the asymptotic 95 per cent error bounds for a sub-sample. As can be seen from Figure 4 the plot does not indicate non-constancy in our model. Figure 5 presents various plots associated with diagnostic testing: actual and fitted values, the standardised residuals, a histogram of the standardised residuals with histogram of the standard Normal distribution on the background and the correlogram of residuals. There are some evidences of residual autocorrelation at the beginning of the period, but on the average the diagnostic statistics supports model adequacy. At the end, we can conclude that the final model is generally satisfactory, being a well interpretable and statistically acceptable model for describing export supply function in Croatia.

Figure 1.: Plot of the Trace Statistic

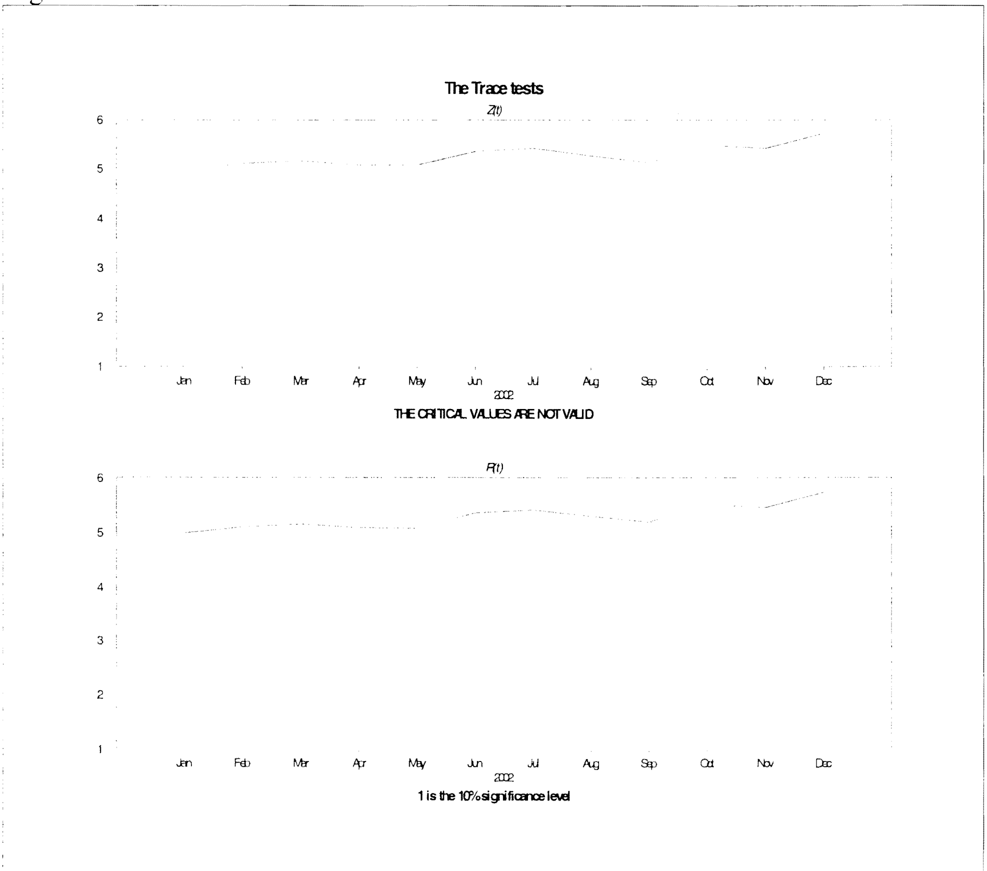


Figure 2.: The Log-Likelihood Value

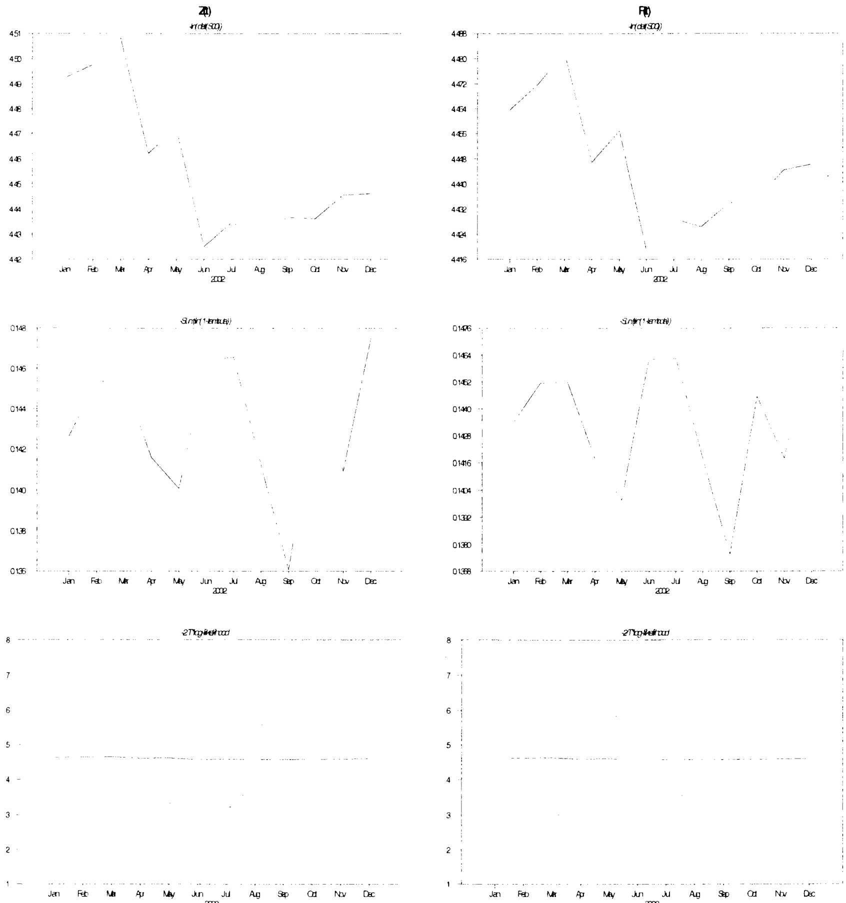


Figure 3.: Test of Constancy of $\hat{\beta}$

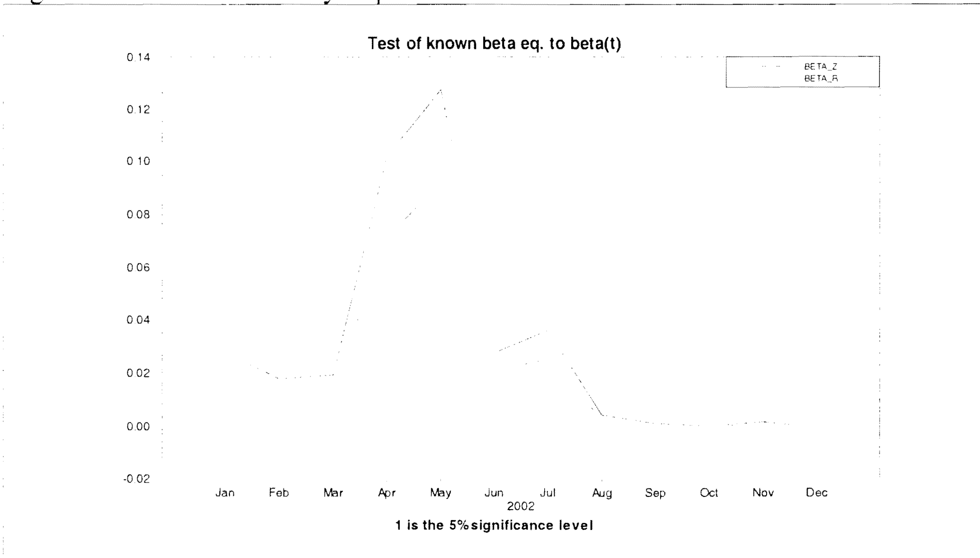


Figure 4.: The non-zero Eigenvalue[17]

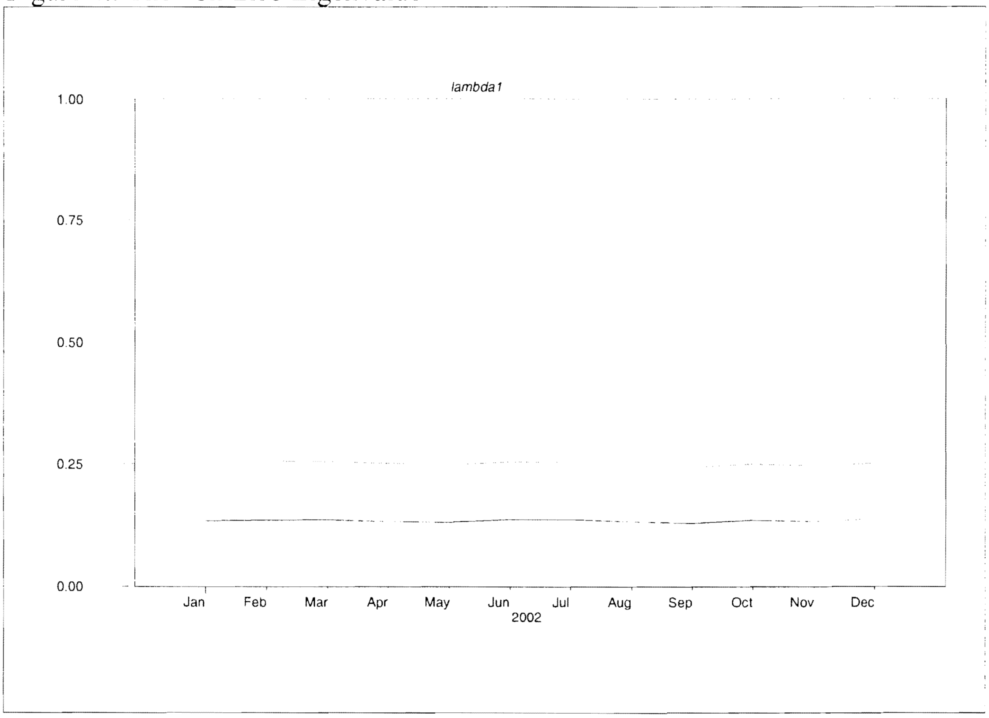
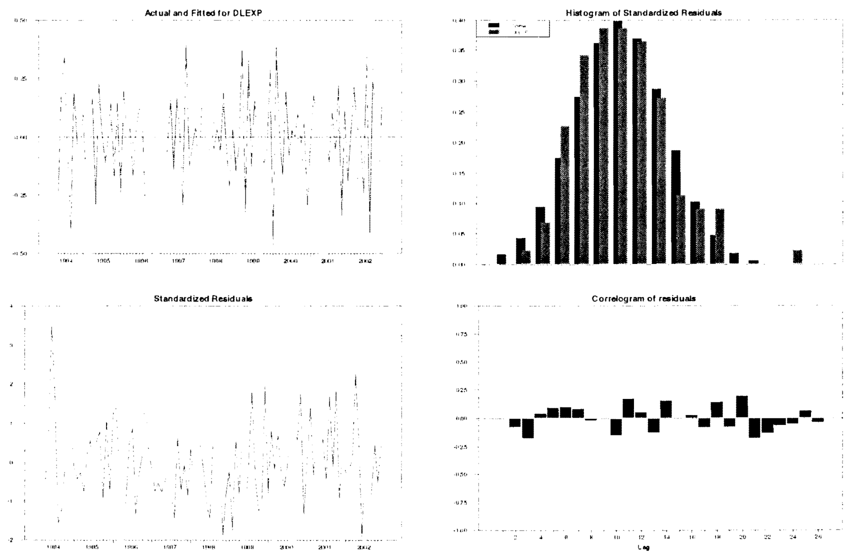


Figure 5.: Residual Analysis for *exp*



Conclusion

In this paper, through the analysis of Croatian export supply, we tried to evaluate the role and significance of the real exchange rate and income in the short- and the long-run. Therefore, this study provides some insights into the importance played by the real exchange rate and production capacity in explaining total Croatian export. The model is tested over the period 1/1994-12/2002. In the long-run variables income and real exchange rate have expected signs and their influence on export supply is statistically significant. Lagged changes of the real exchange rate are not statistically significant for the current total export trade in the short-run. This implies that Croatian total export reacts parsimoniously and slowly to the changes in the real exchange rate. The phenomenon could be due to the large share those products in Croatian exports that are exchange rate-inelastic. On the other hand, income has significant impact on export supply in the short-run.

Our study shows that the real exchange rate and production capacities are important determinants of Croatian export supply. Therefore, attention should be paid to the development of the real exchange rate, especially whenever the promotion of manufactured exports is considered important.

NOTES

1 In the fourth quarter of 1993 Croatia introduced a stabilisation program, followed by a currency change. The exchange rate was stabilised and inflation was stopped.

2 The situation is similar to the tests of non-nested hypotheses.

3 All empirical work was performed using RATS and CATS statistical packages of Doan (1992).

4 The appropriate number of lagged differences is determined by adding lags until a LM test fails to reject no serial correlation of order 12 at 5 per cent level.

5 Additionally non-parametric Phillips-Perron unit root tests were performed on all variables in the study as well as unit root tests under structural change in the intercept or the slope, (e.g. Perron, 1997). The results support the conclusion that all variables are integrated of order one. In order to save space, they are not reported here but can be obtained under request.

6 There is an evidence of rejecting a unit root null hypothesis for variable *exp* if model contains trend at 5 per cent significance level. However, all other unit root tests support the unit root null as well as changing a significance level to 1 per cent.

7 $\Gamma_i = -(I - A_1 - ... - A_i), i = 1,...,k-1$ and $\Pi = \alpha\beta' = -(I - A_1 - ... - A_k)$

8 The lag length of the VAR model is determined to solve the trade-off between improving the fit of the model (which requires additional lags) and granting a sufficiently high number of degrees of freedom (which requires parsimonious parameterisation). Minimising SC and HQ information criteria and trying to reduce auto-correlation, but at the same time using common sense, the lag length is set to 3.

9 For the Johansen procedure, there are two test statistics for the number of cointegrating vectors: the trace (λ_{trace}) and the maximum value statistics, ($\lambda_{\max}$). In the trace test, the null hypothesis is that the number of cointegrating vectors is less than or equal to r, where $r = 0$ to 3 . In each case the null hypothesisis is tested againest the general alternative. The maximum eigenvalue test is similar, except that alternative hypothesis is explicit. The null hypotheisis $r=0$ is tested againest the alternative that $r=1$, $r=1$ againest $r=2$, *etc*.

10 Variable remains in the cointegration space, and in current and lagged differences in short-run dynamics.

11 L-B is Ljung-Box test for residual autocorrelation based on the estimated auto- and cross-correlation on the first (T/4) lags, (e.g. Ljung and Box, 1978). LM(1) and LM(4) are LM- type tests for the first and the fourth order autocorrelation (e.g. Godfrey, 1988). The test for normality is Shenton-Bowman test, (e.g. Doornik and Hansen, 1994). ARCH is a test for Auto Regressive Conditional Heteroscedasticity, (e.g. Engle, 1982).

12 Testing weak exogeneity of variables equals testing hypotheses about the rows of vector when the parameters of interest are the long-run parameters α and β.

[13] Conditioning on *reft* and *ind* variables means that variables remain in the long-run model (i.e., cointegration vector) although their short-run behaviour is not modelled because of their exclusion from the vector of the left-hand side of the equation (2). By conditioning on weakly exogenous variables, the rest of the system is likely to behave better statistically, (e.g. Johansen, 1992).

[14] This is re-estimation of cointegration relationship with variables *reft* and *ind* treated as weakly exogenous.

[15] When testing for the exclusion of the variable *reft* from the long-run the obtained value of the LR test is CHISQ(1) = 4.58 (p-value = 0.03) and for the variable *ind*, CHISQ(1) = 13.93 (p-value = 0.00). This confirms that both variables are significant in the long-run.

[16] The time path of the non-zero eigenvalue is well inside the asimptotic 95 per cent error bounds for each sub-sample which do not indicate non-constancy in the partial model.

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