

THE IMPACT OF KUNA EXCHANGE RATE VOLATILITY ON CROATIAN EXPORTS

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Abstract

The aim of this paper is to analyze the monetary transmission mechanism through the influence of exchange rate variability on export volume. To date it has been very common to use "historical volatility" as an approximation for exchange rate variability in empirical studies. However, many macroeconomic time series are characterized by heteroskedasticity, i.e. their variance is not constant over time. Thus in this paper the ARCH model is proposed as a model of conditional heteroskedasticity. Also, as an alternative to ARCH we will introduce historical volatility based not only on future but also on past exchange rate values. In exploring the influence of exchange rate volatility and domestic income on export volume, Johansen's multivariate cointegration approach and error-correction model (ECM) are used. The short run and long run relationships are analyzed separately. The results of econometric analysis draw attention to the different strengths of the relationship between kuna volatility and exports for the two proposed models. The first model shows a mild negative long-run relationship, while the second shows the much stronger aversion of Croatian exporters to volatility as a measure of exchange rate uncertainty.

Key words: ARCH model, Johansen's approach, ECM model, cointegration

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1 Introduction

The issue of export stimulation is an essential part not only of Croatian economic policy but also of wider development policy. This is a problem that has particularly been emphasized in the document “Croatian export offensive”, recently published by the Croatian Government, for it is precisely the poor and stagnating Croatian export results that necessitate further and more detailed analysis of exports and their determinants. Thus in this paper we will particularly examine the existence of the classical dichotomy or duality in the Croatian economy. The main issue is the possibility of influencing real economic trends with monetary measures and instruments. To be more specific, we will deal with the issue of manipulating the kuna exchange rate in order to influence the real economic sphere. Thus in this paper we will analyze the impact of kuna exchange rate volatility and domestic income on Croatian export volume. At this point it is necessary to mention that economic theory gives us quite direct and clear evidence in favour of a strong positive impact of domestic income on export volume, while the influence of exchange rate volatility is rather vague and continuously under discussion.

Economic theory bases the relationship between exchange rate volatility and export volume on an extremely simplified model: a corporation producing a sole commodity and, without importing any intermediate goods, exporting it to a single foreign market. The corporation is paid in foreign currency, and it cannot change its production volume because of the high costs arising. With all that in mind, we can then conclude that the firm’s profit fluctuations depend only on exchange rate volatility. From that point of view volatility can be considered an exchange rate risk; where higher risk leads to reduction of corporate exports (Clark, 1973; Hooper and Kohlhagen, 1978). Empirical research into the connection between exchange rate volatility and export volume have not yet resulted in unambiguous evidence of the strength-direction relationship (IMF, 1984; Cote, 1994; McKenzie, 1999). So far the only Croatian study into this matter has shown that there is no short run relationship between the two variables, while the relationship is positive in the long run (Erjavec, Cota and Bahovec, 2004). In other words, according to this paper, the current Croatian National Bank exchange rate policy of keeping the kuna exchange rate within the minimal range is not justified if it discourages exports. Thus the goal of this paper is to try to provide better statistical results that are, hopefully, economically more meaningful. At the same time, the results obtained should also reflect some recent results from analyzing the volatility of many macroeconomic series.

In analyzing the influence of exchange rate volatility on export volume it is common to use the moving standard deviation of the real effective exchange rate growth rate (historical volatility) (Brodsky, 1984; Kenen and Rodrick, 1986; Frankel and Wei, 1993; Dell’Ariccia, 1999; Rose, 2000; Erjavec, Cota and Bahovec, 2004). In this paper we propose two alternative models for volatility analysis. Concretely, the first model is an ARCH(1) process, introduced to overcome the homoskedasticity problem (Engle, 1982). The second one is actually a modification of historical volatility per se.

The paper is divided in four chapters. After arguments justifying the exploration of exports and their determinants in the introduction there are more detailed methodological explanations and reasons for using an ARCH model, as well as the other econometric techniques used in this paper. The methodology chapter is followed by empirical data

concerning the three analyzed time series (exchange rate volatility, domestic income and export volume). The way in which different approaches to modelling a real effective exchange rate risk can influence its relationship with export volume is analysed. The paper is concluded by a summary of the obtained results, their economic implications, and guidelines for future researches.

2 Methodological approach

In many empirical investigations the standard, so-called historical, volatility is used to approximate exchange rate volatility. It can be mathematically expressed in the following way (Kenen and Rodrick, 1986):

$$V_t = \left[\frac{1}{m} \sum_{i=1}^m (\ln E_{t+i-1} - \ln E_{t+i-2})^2 \right]^{\frac{1}{2}} \quad (1)$$

where E_t stands for real effective exchange rate.

Values obtained by inserting empirical real effective exchange rate data into this relation actually represent a series of simple moving averages, which are defined as mean values of m consecutive elements of the associated time series. They are used for smoothing the time series. The economic logic behind that relation in fact implies that the current volatility value at time t (V_t) actually gives an average of its m following consecutive values. The basic assumption of such formulation of volatility is heteroskedasticity. However, many financial and macroeconomic time series are characterized by a wide range of volatility in certain time periods. Practically, in the case of variance heteroskedasticity of the observed exchange rate volatility series, this would mean that the estimation of parameters in the model could also be inefficient unless the conditional heteroskedasticity of the observed time series is taken into consideration.

Many empirical researches have confirmed that more sophisticated models for volatility evaluation such as ARCH, GARCH or EGARCH lead to more accurate and statistically better estimates (Akgriray, 1989; Chu and Freud, 1996). Therefore on account of analysis results quality (in our case the analysis is referred to the prediction of exchange rate volatility) it would be wise to use a heteroskedastic model with a variance depending on past values of the series itself.

The first models used for empirical analysis of volatility were derived from financial analysis. By observing logarithms of some American stock returns it was found that they were characterized by serial independence of data and inconstancy of volatility through time, and that the distribution of data was asymmetrical with fat tails¹. In other words, the data do not originate from normal distribution, the heteroskedasticity of the variance is ignored, and thus such model is insufficiently realistic for the expression of volatility. The first more sophisticated concept was introduced, with the name of autoregressive conditional heteroskedasticity (Engle, 1982; Engle, 1980), in which the conditional variance is

¹ The so-called "stylized facts of financial data".

not constant through time and is characterized by an autoregressive structure². In ARCH models every observation (X_t) is equal to the product of white-noise process (Z_t) and a positive process σ_t , on the assumption that for every t the variables Z_t i σ_t are independent:

$$X_t = Z_t \sigma_t \quad (2)$$

On the assumption of Gaussian white-noise, i.e.. $Z_t \sim N(0,1)$, the ARCH (q) process is defined as:

$$\sigma_t^2 = \alpha_0 + \alpha_1 X_{t-1}^2 + \dots + \alpha_q X_{t-q}^2, \text{ for each } t \quad (3)$$

under the parameter conditions $\alpha_0, \alpha_i > 0, i = 1, \dots, q$.

A characteristic of such ARCH processes is the existence of the *conditional variance* σ_t , which depends on its one past value, i.e. the calculation is conditional on information available up to moment t . In other words, we can say that the model is based on a property of heteroskedasticity, which is also a characteristic of all macroeconomic time series used in our analysis. Precisely that heteroskedasticity was in fact the reason for introducing an ARCH(1) model for expressing exchange rate volatility in this analysis. If we denote $\log(E_t)$ with Y_t , where E_t stands for the real effective kuna/euro exchange rate, equations (2) and (3) can be therefore be expressed in the following way:

$$Y_t = C + V_t Z_t \quad (4)$$

$$V_t^2 = \omega + \alpha(Y_{t-1} - C)^2 \quad (5)$$

where

$$\omega > 0 \text{ and } \alpha > 0, \text{ and } \alpha \leq 1 \quad (6)$$

According to Jondeau, Rockinger and Poon (2006), it follows that in the case of an ARCH(1) model the necessary condition and the sufficient condition for the process V_t^2 to be strongly stationary is:

$$E[\ln(\alpha Z_t^2)] > 0 \quad (7)^3$$

One of the key advantages of ARCH models is the possibility of predicting certain values for one observed time period in advance. In this paper we are, thus, interested in predicting the value of volatility V_{t+1} with the data available up to moment t .

The estimation of the observed model is based on 132 monthly data for real effective kuna/euro exchange rate in the period from January 1996. to December 2006. Accord-

² It is shown in many empirical studies that the GARCH(1,1) process is able to present most time series, and that data groups that require modeling of processes of higher order than GARCH(1,2) or GARCH(2,1) are extremely rare (Berra, Higgins, 1993). That is why in this paper we first considered a GARCH(1,1) model. However, the estimation showed that the second parameter in the model is equal to zero, so such a model was actually reduced to an ARCH(1) (Jondeau, Rockinger, 2007). That is the reason why in this paper we analyzed only an ARCH(1) model as one of the possible alternative approaches for expressing volatility.

³ It can be instantly noticed that, in the case of $\alpha = 1$, condition (7) is fulfilled.

ing to the proposition, Z_t is a series of independently and identically distributed variables, such that $Z_t \sim N(0, 1)$, so the likelihood function has the following form:

$$L_t = \frac{1}{T} \sum_{i=1}^T \left[-\frac{1}{2} \ln(2\pi) - \frac{1}{2} \ln(V_t^2) - \frac{1}{2} \frac{Y_t}{V_t^2} \right] \quad (8)$$

where T is the number of the observed data.

Let us denote the unknown parameters vector with $\theta = (C, \omega, \alpha)$. It is then necessary to find vector θ for which the function L reaches the maximal value under the conditions given in equation (6). The maximization of function L_t is carried out using a numerical logarithm for finding the function maximum under the given parameter conditions. Their estimated values are given in Table 1⁴:

Table 1 Estimated ARCH(1) model parameters

Parameter	Value	Standard error	t-value
C	4.5613	0.0021	2170.6687
ω	$7.04 \cdot 10^{-5}$	$4.67 \cdot 10^{-5}$	1.5903
α	1	0.36297	2.7551

Source: authors' calculations

From this econometric analysis it is quite clear that the parameter α is equal to 1, which indicates the nonstationarity of the observed time series. This implies that the conditional variance will approach infinity as time variable t approaches infinity. A situation of this kind is commonly known in econometric literature as a series "explosion".

Economically, the fact that $\alpha = 1$ could mean that, on average, process Y_t (logarithm of the real effective kuna exchange rate) is characterized by extremely large fluctuations in short time periods.

The above mentioned stationarity will later be additionally corroborated with the results of unit root tests. Thus, the volatility of a time series obtained by an ARCH model shows characteristics of nonstationarity. The existence of a unit root in the conditional variance does not affect the distribution of estimators because in that case they are also normally distributed. Therefore it is possible to draw conclusions about the model on the basis of standard test statistics (Lumsdaine, 1996). Being aware of the topicality and the necessity of analyzing Croatia's exports and their determinants, we decided to include the volatility obtained in this way in our analysis.

The other suggested alternative model for analyzing volatility is given here:

$$V_t = \left[\frac{1}{2m+1} \sum_{i=-m}^m (\ln E_{t+i} - \ln E_{t+i-1})^2 \right]^{\frac{1}{2}} \quad (9)$$

⁴ Parameters are obtained using Matlab software.

From this relation it is clearly visible that 12 successive past values and 12 successive following values of the real exchange rate are used to calculate a single element of the V12 time series. Introducing this formulation of volatility is justified by the need of monetary policy holders to make decisions based not only on future but also on past exchange rate values (which is the case in relation 1).

The main reason for introducing these two alternative volatility models into the analysis of the effect of kuna exchange rate variability on export volume is the fact that their application represents a methodological innovation in exploring the relationship between the two variables in Croatia. The ultimate goal that we want to achieve with this step is the achievement of better statistical results in the estimation of the explained relationship.

After exchange rate volatility, the second explanatory variable in our model is industrial production volume, which has been proven to be a good indicator of the total economic activity in a certain country. In spite of the fact that the share of industrial production in total Croatian production has decreased through the past fifteen years, this variable's short-term variability cyclically follows and indicates Croatian real GDP movements extremely well (Cerovac, 2005).

Also, we have to emphasize that real export data do not exist in official Croatian statistics. Therefore, we have used the nominal export of goods and services expressed in the Croatian national currency (the kuna) as an approximation of export level. The data for all three mentioned variables are monthly and apply to the time period 1996/1-2006/12.

Since we want to examine the dependence of export volume on industrial production and exchange rate volatility, we will here define the observed model:

$$export = f(ind, Vol) \quad (10)$$

where the values of *ind* time series are expressed in the form of a base index (average 2000=100). The real kuna/euro exchange rate data used in both suggested volatility series is also in the form of index, with 2001. as the base year. The *export* and *ind* represent the logarithm values of the same name time series.

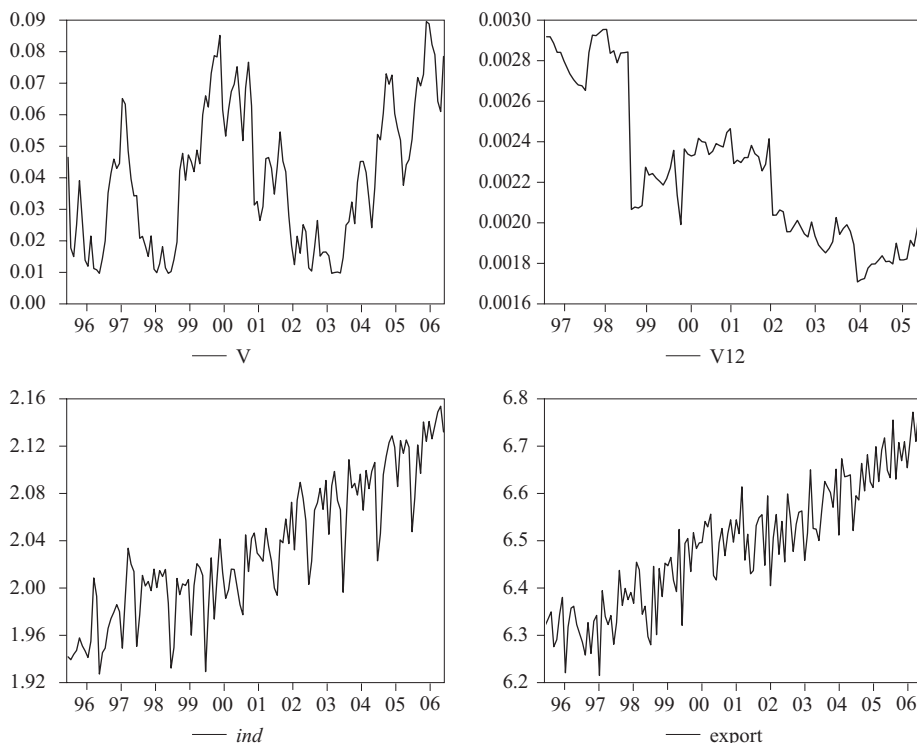
Most macroeconomic time series are unstationary by their nature (Asteriou, 2006). One of the causes of unstationarity of the time series can be the presence of unit root. Then the application of a classical linear regression model to that series could lead to unrealistically high values of R^2 and t -statistics (the so-called "spurious regression"). In this case, we can dispose of some more sophisticated econometric techniques like the Engle-Granger or the Johansen approach. Both mentioned techniques are based on the principle of cointegration, i.e. the existence of the long-term relationship between two or more variables. However, Johansen's approach is multivariate and treats the possibility of more than one cointegration vector (Johansen, 1991), so we decided to use exactly that approach.

3 Empirical data

As the first step of the cointegration analysis (examining the long-run relationship between our variables) it is necessary to determine whether all the observed variables integrated are of the same order. In stationary time series the effects of short-term shocks

are eliminated and the series revert to their mean values in the long run (Asteriou, 2006). Since Figure 1 clearly shows that for all four observed time series that is not the case, this kind of plot can indicate the unstationarity of our time series. Such indications will later be confirmed by a formal unit root test.

Figure 1 Observed time series in levels



Source: authors' calculations

In order to analyze the empirical data an ADF⁵ unit root is performed (Dickey-Fuller, 1979). It is formed in such a way that the null hypothesis (H_0) implies unstationarity, i.e. the presence of unit root in the time series of interest.

Therefore we can say that the test results lead us to the conclusion that all variables of interest to us are integrated of first order, (I(1)), i.e. they are unstationary in levels and stationary in first differences. It is important to note that in further analysis we will be observing two different models. In the first one we will use an ARCH model for volatility approximation (time series V), and in the second one a modification of historical volatility will be used. The modification will include the simple moving averages of the standard

⁵ All data obtained using E-Views software and Mackinnon's critical values table (1996)

deviation of 12 previous and 12 following values of real effective exchange rate growth rate (time series V12). Let us then begin with the first model.

Table 2 ADF test in levels

Variable	Constant and trend	Constant	Model without a constant or trend
export	-2.107934 (11)	-0.934560 (11)	3.519571 (11)
ind	-1.324893 (12)	1.277858 (12)	3.451840 (12)
V12	-3.252733 (0)	-1.956908 (0)	-1.183389 (0)
V	-2.813696 (1)	-1.988642 (2)	-0.416649 (2)

Source: authors' calculations

Table 3: ADF test in first differences

Variable	Constant and trend	Constant	Model without a constant or trend
Δ export	-5.695471*(10)	-5.706614*(10)	-2.753500*(12)
Δ ind	-5.739134*(11)	-5.415932*(11)	-2.990529*(12)
Δ V12	-11.18570* (0)	-11.20553* (0)	-11.17002* (0)
Δ V	-9.246821* (1)	-9.273375* (1)	-9.281676* (1)

Remark: Values denoted by * indicate rejection of the null hypothesis of unstationarity at the 5% significance level. The optimal lag length for each variable is given in brackets after each test statistic, and is obtained using the Akaike information criterion (Asteriou, 2006).

Source: authors' calculations

Due to the obtained results of the ADF test, a VAR (vector autoregressive) model with all variables included is given in the following way (Harris, 1995):

$$Z_t = A_1 Z_{t-1} + A_2 Z_{t-2} + \dots + A_k Z_{t-k} + \psi D_t + u_t \quad (11)$$

where Z is a vector of all n system variables (in our case the variables *export*, *ind* and *Vol*), and u_t represents the n -dimensional vector of error terms with the mean equal to zero and a covariation matrix Σ . Vector D_t includes 11 dummy variables used to overcome the seasonal influences. That VAR model is defined in a general way, while we will obtain the optimal lag length for variables in relation (11) by carrying out an adequate test. Guided by the Akaike information criterion we concluded that 3 is the optimal lag length for our model.

Variables are cointegrated if and only if we can reformulate equation (11) to define the so-called VECM (vector error correction model) (Asteriou, 2006):

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \Gamma_2 \Delta Z_{t-2} + \dots + \Gamma_{k-1} \Delta Z_{t-k+1} + \Pi Z_{t-1} + \psi D_t + u_t \quad (12)$$

the Π matrix can additionally be written as

$$\Pi = \alpha x \beta' \quad (13)$$

where α includes the speed of adjustment to equilibrium coefficients. That is, due to certain negative shocks the economy moves away from the long-run equilibrium, and the elements of the α matrix measure the speed of return to the equilibrium. β' is the long-run coefficients matrix, i.e. its elements show the long-run relationship between the variables in the model.

One of the most important steps of the cointegration analysis is determining the rank of the long-run matrix Π , or determining the maximal number of linearly independent columns in this matrix. This number actually represents the number of cointegration vectors. There are two methods for determining the number of cointegration relationships: the trace statistic and the maximum eigenvalue statistic. The first one tests the null hypothesis that the number of cointegration vectors is less than or equal to r , while the alternative suggests that the number is equal to k , where k is the number of endogenous variables. The second method tests the null hypothesis that $r=0$, while the alternative suggests that $r=1$, then $r=1$ opposite to the alternative that $r=2$, etc. The results of the determination of the cointegration vectors number is shown in the following tables:

Table 4 Determination of the number of cointegration vectors (trace statistic)

$H_0: r =$	Eigenvalue	Trace statistic	0.05 Critical value	0.1 Critical value
0	0.184941	40.36490 ^a	42.91525	39.75526
1	0.082896	14.18963	25.87211	23.34234
2	0.024028	3.113155	12.51798	10.66637

^a denotes rejection of the null hypothesis at the 10% significance level

Source: authors' calculations

Table 5 Determination of the number of cointegration vectors (maximum eigenvalue statistic)

$H_0: r =$	Eigenvalue	Max-eigenvalue	0.05 Critical value	0.1 Critical value
0	0.184941	26.17527 ^a	25.82321	23.44089
1	0.082896	11.07647	19.38704	17.23410
2	0.024028	3.113155	12.51798	10.66637

^a denotes rejection of the null hypothesis at the 5% significance level

Source: authors' calculations

Here both test statistics lead us to the conclusion that the long-run relationship between our variables is determined by only one cointegration vector. We also applied the test for including the deterministic elements in the model, and the obtained results are given in the following table:

Table 6 Including the deterministic elements in the model⁶

Coint. equation		Constant	Constant	Constant & linear trend
VAR		–	Constant	Constant
Number of equations	trace	0	0	1
	max. eigen	0	0	1

Source: authors' calculations

As both indicators (the maximum eigenvalue statistic and the trace statistic) indicate the existence of a constant and a linear trend in the cointegration relationship, we can express this with the following equation:

$$export = -7.589764 - 0.003994 t + 0.689808 ind - 0.860479 V \quad (14)$$

From this relation we can conclude that Croatian exports are positively correlated with domestic income, while the relationship with exchange rate volatility is negative. At the same time, using *t*-values for *ind* (1.60540) and *V* (-3.56157) we can conclude that, at the 10% significance level, the impact of volatility and industrial income is statistically significant. Rather high trend *t*-values (-6.68589) indicate that the trend is also statistically significant, but at the 10% significance level.

Using the obtained cointegration vector we can now define the error correction model:

$$\Delta export = \zeta_1 + \sum_{i=1}^3 b_i \Delta export_{t-i} + \sum_{i=1}^3 c_i \Delta ind_{t-i} + \sum_{i=1}^3 d_i \Delta V_{t-i} + \alpha_1 ECM_{t-1} + \sum_{i=1}^{11} g_i D_{t-i} + u_t \quad (15)$$

where D_{t-i} represents the seasonal dummy variables introduced to solve the seasonality problem in the data. Parameter ECM ensures convergence towards the long-run equilibrium, while the α_1 coefficient measures the speed of adjustment to the mentioned long-run steady-state. Evaluation of parameters from the VEC model is given in Table 7:

Because the VEC model represents the approximation of the short-run export function, on the basis of the *t*-value of the ECM coefficient ($t=-4.27997$) we can conclude that the error correction term is also statistically significant at the 5% level. The value of the ECM parameter ($\alpha_1 = -0.757946$) implies that 75.79% of the long-run equilibrium devia-

⁶ The MacKinnon-Haug-Michelis critical values table (1999) was used.

tion is corrected monthly. The obtained t-value for the ΔV_{t-1} parameter ($t=-2,14512$) suggests that the short-run relationship between kuna exchange rate volatility and Croatian exports in this model is determined by only a one-month time lag. Of course, this conclusion applies to the 5% significance level, for this is the extent to which exports are sensitive to kuna volatility changes.

Although the determination of the Π matrix rank gives us the answer to the question about the number of cointegration vectors, it is necessary to complete this with the analysis of the weak exogeneity of the model variables. All included variables have up to this point been treated as potentially endogenous. Thus with that goal we applied the χ^2 test, which tests the hypothesis that the adjustment parameter for the observed variable in the VEC model is equal to zero (Harris, 1995). That is to say, in the case of accepting the null hypothesis we can conclude that the observed variable is weakly exogenous.

Table 7 VECM model

Variable	Coefficient	t-value	Variable	Coefficient	t-value
ζ	0.006287	0.35300	D_{t-1}	-0.077519	-3.30510
$\Delta \text{export}_{t-1}$	-0.331254	-2.02016	D_{t-2}	-0.047702	-1.40771
$\Delta \text{export}_{t-2}$	-0.134547	-0.96135	D_{t-3}	0.034553	0.99283
$\Delta \text{export}_{t-3}$	-0.034325	-0.35165	D_{t-4}	0.025413	0.84337
Δind_{t-1}	0.054044	0.19015	D_{t-5}	-0.003448	-0.15848
Δind_{t-2}	0.093412	0.31812	D_{t-6}	-0.007162	-0.34057
Δind_{t-3}	0.405528	1.49440	D_{t-7}	0.031083	1.38548
ΔV_{t-1}	-1.087370	-2.14512	D_{t-8}	-0.050553	-2.44957
ΔV_{t-2}	-0.347279	-0.70503	D_{t-9}	0.010997	0.45026
ΔV_{t-3}	-0.350109	-0.71238	D_{t-10}	0.051901	2.10260
ECM	-0.757946	-4.27997	D_{t-11}	0.013952	0.59721

Source: authors' calculations

Table 8 Weak exogeneity test

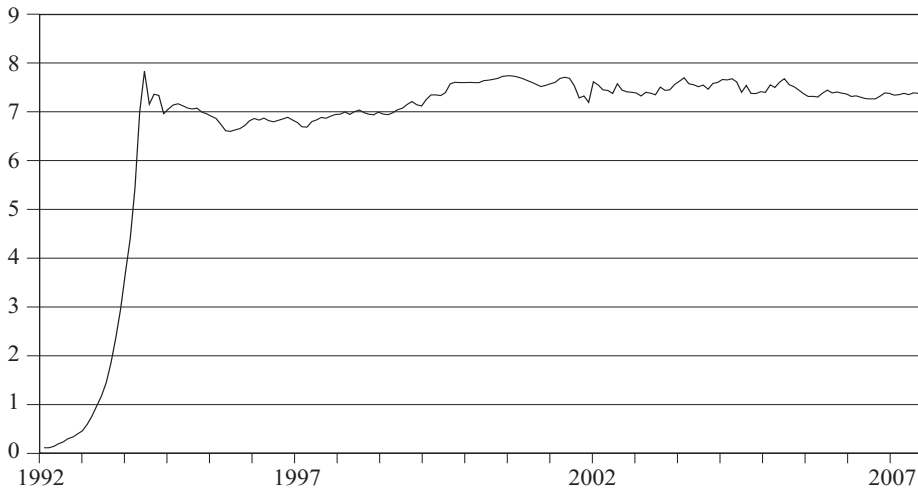
Variable	export	ind	V
Test statistic	$\chi^2 = 12.93052$	$\chi^2 = 5.536632$	$\chi^2 = 0.852173$
p-value	0.000323	0.018622	0.355938

Source: authors' calculations

The obtained p-values and the χ^2 statistics unambiguously lead us to the conclusion that *export* and *ind* are endogenous variables, while volatility V, on the other hand, is weakly exogenous. Thus, as a result of a deviation from the long-run model equilibrium,

volatility reacts just slightly.⁷ Such results are in fact just the expected due to the fact that the Croatian National Bank, under pressure from the Maastricht criteria, is carrying out a quasi-fixed monetary policy. In other words, it regulates the exchange rate movements within a very narrow range. If we observe the movement of the kuna/euro exchange rate over the past fifteen years, we will very easily see that, since the Stabilization Programme of 1993 it has been held at a practically fixed level. (www.hnb.hr, 2007).

Figure 2 Midpoint kuna/euro exchange rates of Croatian National Bank (end of period), January 1992- March 2007



Source: authors' calculations

Let us now observe the second model, where we expressed volatility as a modification of the standard historical volatility. Here we denote the volatility time series by V12. With the previously proved fact that the variable $V12 \sim I(1)$, in the next step we analyze the optimal lag length in the VAR model. Just as in the first model, the information criteria here also suggest 3 as the optimal lag length. The third step of the Johansen approach is to determine the number of cointegration equations:

Table 9 Determination of the number of cointegration vectors (trace statistic)

$H_0: r =$	Eigenvalue	Trace statistic	0.05 Critical value	0.1 Critical value
0	0.214373	47.70672 ^a	42.91525	39.75526
1	0.134460	19.60371	25.87211	23.34234
2	0.043895	4.33089	12.51798	10.66637

^a Denotes rejection of the null hypothesis at the 5% significance level

Source: authors' calculations

⁷ Similar results are given in other papers; for a more detailed review, see (Vizek, 2006).

Table 10 Determination of the number of cointegration vectors (Maximum eigenvalue statistic)

H ₀ : r =	Eigenvalue	Max-eigenvalue	0.05 Critical value	0.1 Critical value
0	0.214373	28.10301 ^a	25.82321	23.44089
1	0.134460	15.27284	19.38704	17.23410
2	0.043895	4.330869	12.51798	10.66637

^a Denotes rejection of the null hypothesis at the 5% significance level

Source: authors' calculations

Both sets of statistics (trace and maximum eigenvalue) again lead us to posit the existence of a single cointegration relationship. At the same time, the results of the test for including the deterministic elements in the model (given in the next table) point to the need to include a linear trend and a constant in the model.

Table 11 Including the deterministic elements in the model⁸

Coint. equation		Constant	Constant	Constant & linear trend
VAR		–	constant	constant
Number of equations	trace	0	0	1
	max-eigen	0	0	1

Source: authors' calculations

Thus the cointegration vector can be approximated in the following way:

$$export = -10.56368 - 0.007272 t + 2.326924 ind - 117.3238 V12 \quad (16)$$

The obtained *t*-values (-8.60845 for *t*, 4.92680 for *ind* i -3.32224 for V12) show that all three variables in the model are significant in the long run (5% significance level). The estimated impact of the domestic income is, expectedly, again positive. However, the highly negative effect of volatility on export volume is quite intriguing. Such results obviously imply that Croatian exporters respond very negatively to an increase in kuna volatility. In other words, the result of the exchange rate risk growth in the observed period was the downfall of export. The estimated cointegration vector is expressed in the following equation:

$$\Delta export = \zeta_1 + \sum_{i=1}^3 b_i \Delta izvoz_{t-i} + \sum_{i=1}^3 c_i \Delta ind_{t-i} + \sum_{i=1}^3 d_i \Delta V_{t-i} + \alpha_1 ECM_{t-1} + \sum_{i=1}^{11} g_i D_{t-i} + u_t \quad (17)$$

Here are also the estimated parameters from the previous equation:

⁸ The MacKinnon-Haug-Michelis critical values table (1999) was used.

Table 12 VEC model

Variable	Coefficient	t-value	Variable	Coefficient	t-value
ζ	-0.009858	-0.46432	D_{t-1}	-0.087964	-3.09002
$\Delta\text{export}_{t-1}$	-0.711433	-3.56114	D_{t-2}	-0.024136	-0.57571
$\Delta\text{export}_{t-2}$	-0.467361	-2.55494	D_{t-3}	0.080059	1.76756
$\Delta\text{export}_{t-3}$	-0.158888	-1.29776	D_{t-4}	0.068921	1.64257
Δind_{t-1}	0.112195	0.24613	D_{t-5}	0.017258	0.56464
Δind_{t-2}	0.426054	1.03057	D_{t-6}	0.002205	0.08751
Δind_{t-3}	0.615321	1.74282	D_{t-7}	0.040842	1.50557
ΔV12_{t-1}	-13.43439	-0.28750	D_{t-8}	-0.017665	-0.71882
ΔV12_{t-2}	78.09946	1.69311	D_{t-9}	0.037621	1.31631
ΔV12_{t-3}	-29.28236	-0.62879	D_{t-10}	0.0069820	2.31139
ECM	-0.278838	-1.48725	D_{t-11}	0.022028	0.75151

Source: author's calculation

Low volatility t-values instruct us that for all three time lags, unlike the long run, in the short run there does not exist a statistically significant relationship between volatility and export volume.

Table 13 Weak exogeneity test

Variable	export	ind	V12
test statistic	$\chi^2 = 1.353201$	$\chi^2 = 10.07904$	$\chi^2 = 5.042692$
p-value	0.244719	0.001500	0.024730

Source: author's calculation

This time the results of the weak exogeneity test are somewhat different. In our first model export was proven to be an endogenous variable. However, here on the basis of high χ^2 statistic and p-value export can be considered as a weakly exogenous variable, which in fact leads us to the existence of a relationship in the opposite direction. In other words, exchange rate volatility should be examined as a function of export volume. Such evidence clearly show us that the analysis results significantly differ for various ways of modelling the volatility itself.

4 Conclusion

The aim of this paper was to examine the functioning of the monetary transmission mechanism in Croatia using econometric analysis. To be more precise, we questioned the influence of monetary policy on real economic movements (export volume) through exchange rate manipulations. For this two models for the approximation of the exchange rate volatility were introduced: the conditional heteroskedastic (ARCH model, and a modification of the standard formula for volatility modelling. The latter does not include in the calculation only the future but also the past exchange rate values. Such a model is then

used to predict a certain volatility value in the near future. The long-run effect of volatility obtained by an ARCH model on export volume has been shown to be negative, but rather small, which would mean that Croatian exporters react to a kuna volatility rise by a small, but still existing export reduction. Concretely, an increase in volatility of 1%, *ceteris paribus*, induces an export decline in this model by 0.86%. It is, therefore, possible to conclude that the Croatian export sector is characterized by an aversion to exchange rate risk. Such results are completely identical to the principles of economic theory (Clark, 1973; Hooper and Kohlhagen, 1978), but contradict the only Croatian empirical research of this kind known to the author (Erjavec, Cota and Bahovec, 2004). Furthermore, the weak exogeneity test in that model has shown that volatility must be observed as a weakly exogenous variable, which means that in the short run it has not shown the ability to adapt to steady-state deviations. Also, it was shown that in the short run the relationship between exchange rate volatility and export volume is again negative, and statistically significant for only a one-month time lag.

When it is a question of the long-run relationship between the time series V12 and export volume, it was shown that the estimated elasticity coefficient of exports is very high and can amount to as much as -117.3238. Here, thus, it is shown even more convincingly that the increase of kuna volatility discourages Croatian exporters, so they respond by a reduction of export volume. In that sense the quasi-fixed monetary policy carried out by the Croatian National Bank can be characterized as fully justified, because otherwise the consequence would be a significant export destimulation. Besides, economic theory also suggests a fixed exchange rate regime as an optimal way to overcome exogenous shocks for a small open economy like that of Croatia (Babić, 2003). Opposite to the long-run case, the short-run effect of kuna volatility on export volume is not statistically significant. The argument of weak exogeneity of export volume in this model vividly shows the extent of the influence of the way in which volatility is modelled on the analysis results per se.

To conclude, this paper has examined only one small aspect of the influence of monetary policy on the real economy sphere. A holistic approach to the research of the current CNB exchange rate policy would demand a much wider perspective. Just for example, it would include the influence of exchange rate variability on the movements of Croatian external debt, production volume, the lending policy of commercial banks, as well as Croatian convergence on the Maastricht economic criteria of the European Union. In that sense it would be rather interesting to analyze the way in which kuna exchange rate movements might impact all the above mentioned variables in the long run, as well as in the short run.

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