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TESTING THE FISHER EFFECT IN CROATIA: AN EMPIRICAL INVESTIGATION

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ABSTRACT

In his celebrated book The Theory of Interest Irving Fisher asserted that a percentage increase in the expected rate of inflation would lead to a percentage increase in the nominal interest rates assuming that real interest rate is constant. The so-called Fisher effect suggests that changes in the nominal interest rate reflect the revised inflation expectations; and revised inflation expectations have an impact on the level of the nominal interest rate. Consequently, the monetary authorities should employ strategies that will prevent inflation from rising if nominal interest rates are to be kept at low levels in order to not discourage the borrowing. The aim of this paper is to empirically test the Fisher effect in Croatia using VEC (vector error correction) model. The results suggest that the "full" Fisher effect in Croatia may de facto hold only in the long-run. This paper is organized as follows. Section 2 after the introduction, reviews the literature. Section 3 reviews used data while Section 4 describes methodology, empirical analysis and the results. Finally, Section 5 provides some concluding remarks.

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I. INTRODUCTION

Relationship between the nominal interest rates and the expected inflation is one of the most important bonds that explain the functioning of financial markets. Fisher (1930) asserted that a percentage increase in the expected rate of inflation would lead to a percentage increase in the nominal interest rates assuming that real interest rate is constant. This further means that the nominal interest rate consists of an expected "real" rate plus an expected inflation rate. The mathematical representation of the Fisher effect is as follows:

$$(1+i_t) = (1+r_t^e)\cdot(1+\pi_t^e) \tag{1}$$

$$1+i_t = 1+r_t^e+\pi_t^e+(r_t^e\cdot\pi_t^e) \tag{2}$$

where i_t is nominal interest rate, π_t^e is expected inflation and r_t^e is real interest rate. The following equation represents the Fisher equation:

$$i_t = r_t^e+\pi_t^e+(r_t^e\cdot\pi_t^e) \tag{3}$$

The term $(r_t^e\cdot\pi_t^e)$ is usually very small and thus can be neglected. Therefore, the approximation of the Fisher equation is:

$$i_t \approx r_t^e+\pi_t^e \tag{4}$$

$$r_t^e \approx i_t-\pi_t^e \tag{5}$$

Finally, the Fisher effect can be written as $\Delta i_t = \Delta\pi_t^e$ proposing a one-to-one relationship between i_t and π_t^e.

Additionally, the Fisher equation has also some implications (Jareño and Tolentino 2012).

If $\pi_t^e = 0$, then $i_t = \pi_t^e$. If so, money is not loosing or gaining any value and thus the cost of holding money is equal to its opportunity cost, the real return on assets. Under this condition r_t^e cannot be negative, as $i_t > 0$.

If $\pi_t^e > 0$, then $i_t > r_t^e$. For a positive inflation rate, nominal interest rates will always exceed real interest rates.

If $\pi_t^e < 0$, then $i_t < r_t^e$. For a negative inflation rate, real interest rates will always exceed nominal interest rates.

For a given *i*, the higher π_t^e, the lower *r* $\partial r_t^e / \partial\pi_t^e = -1$.

The last implication is particularly relevant if an economy is in a liquidity trap where i_t cannot be influenced by the central bank anymore.

It should be emphasized that the Fisher effect is a phenomenon that appears in the long-run and may not be present in the short-run[1]. The quantity theory of money states that, in the long-run, change in the money supply result in corresponding amounts of inflation, and in addition, economists generally agree that in the long-run, change in the money supply does not have an effect on real variables. Therefore, a change in the money supply should not have an effect on the real interest rate. If it does, then all changes in inflation must be reflected in the nominal interest rate under the principle of one-to-one, as stated by Fisher. This means that the long-run real interest rate is established in the real sector of the market by means of "technology and preferences". The fact that in the long-run change in the money supply results only in corresponding amounts of inflation is known "the long-run neutrality of money". Hence, the Fisher effect implies the neutrality of money and therefore has important policy implications for the behavior of interest rates and the efficiency of financial markets. The monetary authorities should employ strategies that will prevent inflation from rising if nominal interest rates are to be kept at low levels in order to not discourage the borrowing.

Consequently, taxation can also affect the Fisher effect. Weidmann (1997) asserted that "full" (i.e. one-to-one) Fisher effect applies only to economies without taxes while Darby (1975) notes that when the nominal interest rate is taxed, the Fisher relationship implies that the change in the nominal interest rates is greater than the change in expected inflation so as to maintain the constant ex-ante real interest rate.

As the primarily focus of this paper is to empirically test the Fisher effect by applying the concept of cointegration, it is necessary to point out some of the important issues regarding the relationship between inflation and interest rate. Namely, we can discuss three following cases. First, if inflation and interest rate are not cointegrated, then there is no long-run relationship between them. Second, there is a positive long-run relationship between the inflation and interest rate, but the cointegration vector does not correspond to the "full" Fisher equation. This case corresponds to a so called "partial" (i.e. "weak") Fisher effect in which inflation affects interest rate in the amount larger or less than one. And third, there is a cointegration relationship between inflation and interest rate corresponding to the "full" (i.e. "strict") Fisher effect in which inflation and interest rate move together under the principle of one-to-one[2]. Congruently, it follows that cointegration is a necessary condition for the Fisher effect to hold in the long-run. To conclude, this paper will try to empirically test the possible presence of the "full" long-run Fisher effect not withstanding its possible presence in the short-run.

This paper is organized as follows. Section 2 after the introduction, reviews the literature. Section 3 reviews used data while Section 4 describes methodology, empirical analysis and the results. Finally, Section 5 provides some concluding remarks.

II. LITERATURE REVIEW

This section chronologically summarizes the most relevant papers on the Fisher effect across different time span and countries.

Mishkin (1991) tested the Fisher effect for the US economy using the concept of

[1] For an explanation of why the Fisher hypothesis does not necessarily hold in the short-run see Fisher (1930).
[2] Also see Mishkin (1991), Mishkin (1994), Weidmann (1997), Westerlund (2006), Obi, Nurudeen and Wafure (2009), Argyro (2010), Ray (2012) etc.

cointegration and monthly data in the period from February 1964 till December 1986. By using the data on inflation rate and one to twelve-month US Treasury bills rate, he found evidence that does not support the presence of a short-run Fisher effect but does support the existence of a long-run Fisher effect.

King and Watson (1992) tested the long run neutrality and superneutrality of money, a vertical long run Philips curve and the Fisher effect for the postwar US economy using the VAR (vector autoregressive) model and quarterly data from March 1949 till December 1990. By testing the linkages between gross national product, M2, unemployment, inflation rate and the nominal interest rate on three-month Treasury bills, they found that the US postwar data are consistent with the neutrality of money and a vertical long run Philips curve, but also found evidence against the superneutrality of money and the long run Fisher relation.

Mishkin and Simon (1994) tested the Fisher effect in Australia using the concept of cointegration and quarterly data from September 1962 till December 1993. By using the data of inflation rates and thirteen-week Treasury note interest rates, he found evidence that does not support the presence of a short-run Fisher effect but does support the existence of a long-run Fisher effect.

Crowder and Hoffman (1996) investigated the relationship between nominal interest rates and inflation, i.e. the Fisher equation for the US economy using the VEC (vector error correction) model and quarterly data in the period from March 1952 till December 1991. By using the data of three-month Treasury bills rate and annualized log changes in the price deflator for total consumption expenditures as a proxy for expected inflation, they found considerable support for a Fisher equation.

Weidmann (1997) tested the long-run relationship between nominal twelve-month interest rate on Treasury bills and inflation using the VEC (vector error correction) model and threshold cointegration with German monthly data from January 1967 till June 1996. By using the data of the German twelve-month interest rate on Treasury bills and inflation, he found the full Fisher effect only when a threshold cointegration model is estimated.

Paleologos and Georgantelis (1999) tested the joint hypothesis of the Fisher effect and rationality of inflation expectations in Greece applying cointegration technique and quarterly data in the period from March 1980 till Jun 1996. By using the data of three-month Treasury bills rate and inflation rate, they found the invalidity of the Fisher relationship as a long-run equilibrium phenomenon.

Wesso (2000) tested the relationship between the long-term yield bonds and future inflation, i.e. the Fisher effect in South Africa using the VEC (vector error correction) model and quarterly data in the period from March 1985 till Jun 1999. Using the yield on long-term government bonds, the actual inflation rate, the repurchase rate, and the output gap, he showed that the bond yield is cointegrated with the inflation rate, which indicates that in the long-run, permanent movements in actual inflation have been associated with permanent movements in the long-term bond yield. This further indicates that long-term bond yield movements are driven by expected inflation.

Wong and Wu (2001) investigated the Fisher hypothesis in long horizons for G7 and several Asian countries using the OLS (ordinary least squares) and instrumental variables regression estimates and monthly data in the period from March 1985 till Jun 1999. By using the data on inflation rates and stock returns, they found support for Fisher hypothesis, as well as a positive relation between long-horizon nominal stock returns and expected inflation but not between long-horizon stock returns and contemporaneous inflation.

Berument and Mehdi (2002) examined a multi-country Fisher hypothesis in a sample of 26 developed and developing countries using the instrumental variable method and monthly data in the period from April 1957 till Jun 1981 and end in 1998 (for developed countries) and form May 1957 till February 1985 an end in 1998 (for developing countries). By using the data on Treasury bills rates (or landing rates) and inflation rates, they found support for Fisher hypothesis in a various number of developed and developing countries proving that the Fisher hypothesis holds more for the developed countries than the developing ones.

Westerlund (2006) tested the Fisher effect in 20 OECD countries using panel cointegration tests and quarterly data in the period from March 1980 till December 2004. By using the data on nominal short-term interest rates and inflation rates, he found support for the Fisher effect across countries.

Sathye, Sharma and Liu (2008) tested the Fisher effect in India using the Engle-Granger method for cointegration with monthly data from April 1996 till August 2004. By using the data of inflation rate and yield rates on three-month Treasury bills, they confirmed the presence of the Fisher effect in the Indian financial market.

Obi, Nurudeen and Wafure (2009) tested the Fisher effect in Nigeria using co-Integration and error correction approach and annual data from 1971 till 2007. By using the data of interest rate, money supply, overall fiscal deficit and inflation rate, they confirm the presence of partial Fisher effect.

Argyro (2010) tested the Fisher effect in OECD countries using the concept of cointegration and quarterly data. By using the data of short-term interest rates and inflation rates, they found that the full Fisher effect is present for only one country, Canada, while the partial Fisher effect holds for Belgium and Korea.

Piccinino (2011) tested the Fisher hypothesis in the Euro Area using OLS (ordinary least squares) and the concept of cointegration and monthly data from January 1999 till March 2011. By using the data on the European interbank offered rate as a measure for interest rates, and the six-month maturing German Federal Securities as a measure of expected inflation, he provided evidence that the Fisher hypothesis may hold for the Euro Area.

Ramadanović (2011) tested the long-run Fisher effect in United Kingdom, Switzerland and Germany using the VEC (vector error correction) model and monthly data from March 2001 till Jun 2012. By using the data of yield on long-term government bonds and inflation rates, he found little empirical support for a long-run equilibrium relationship between inflation and nominal interest rates in selected countries.

Fatima and Sahibzada (2012) tested the Fisher effect in Pakistan using the VEC (vector error correction) model and annual data from 1980 till 2010. By using the data of real interest rate, money supply and inflation rate, they confirm the presence of Fisher effect.

Jareño and Tolentino (2012) tested the Fisher effect in Spain economy using the OLS (ordinary least squares) and monthly data from February 1993 till December 2004. By using the data of approximated expected inflation rate and the returns of the one-year Treasury debt securities, they found support for the partial Fisher effect.

Ray (2012) tested the international Fisher effect in USA and selected Asian economies using the concept of cointegration and quarterly data from March 2001 till Jun 2012. By using the data of interest rates and inflation rates, they confirmed the presence of partial Fisher effect in USA.

Wallace (2012) tested for a nonlinear Fisher relationship in US economy using nonlinear least squares and monthly data in the period from January 1934 till May 2011. By using the data on actual inflation rate and three-month Treasury bills rate, he showed that the US real interest rate is stationary which supports the existence of the Fisher effect.

It is evident that conclusions regarding the evidence on the Fisher effect vary among countries. Furthermore, different authors use different periods and measures of interest rates and inflation as well as different data observations and estimation techniques. In addition, a good overview of papers regarding the empirical evidence on the Fisher effect can be found in Cooray (2002), Ray (2012), Jareño and Tolentino (2012) etc.

III. DATA

In order to observe the interdependence between the variables of interest a VEC (vector error correction) model is estimated[3]. However, before estimating the model, a brief discussion regarding the selection of variables is made. Although the Fisher effect implies the relationship between expected rate of inflation and interest rates, it should be noted that the expected rate of inflation and the corresponding ex-ante real interest rate are unobservable and hence must be proxied (Weidmann, 1997 and Ramadanović, 2011). In this sense, Fisher was also criticized for utilizing a backward-looking model of inflationary expectations, which induces systematic errors in forecasting that rational individuals would be capable of rectifying (Piccinino, 2011).

Gordon (1971) and Lahiri (1976) suggested that the expected inflation rate, which is unobservable, may be systematically related to past rates of inflation wherein price expectations are determined by a considerably longer history of past price changes. The same principle was also applied by Wesso (2000) who used lagged inflation rates as a proxy for inflation expectations.

Crowder and Hoffman (1996) used the price deflator for total consumption expenditures as a proxy for expected inflation.

Wong and Wu (2001) stressed that because expectation of inflation is not available in general, the estimation has to rely on a regression model of observables.

Ray (2012) assuming the rational expectations said that that expected inflation is equal to actual inflation. Wallace (2012) also used actual inflation as a proxy for expected inflation noting that if inflation expectations are rational, then actual and expected inflation differ only by a white noise error term and will have identical integration properties, thus the actual inflation rate will be a suitable proxy.

On the other side, Mankiw et al. (2003) outlined a number of survey measures which could serve the purpose of providing data for inflation expectations. These include the Michigan and the Livingston Survey, as well as the Survey of Professional Forecasters. Whereas the former encompasses qualitative questions to the general public, the latter two asks the opinion of academics and labour economists. Beside just stated and by exploring the available literature,

[3] The Structural VAR (SVAR), JMulTi and Gretl econometric software's are used for the multiple time series analysis.

Ramadanović (2011) and Jareño and Tolentino (2012) found that different authors suggest several proxies and methodologies for measuring the expected component of inflation rate such as the use of market prices, ARMA forecasts and regressions, VAR models, data filters (like the Kalman and the Hodrick – Prescott filter), government inflation-indexed bonds, naïve models etc.

However, when choosing the proxies for expected inflation one should be careful. Brhel and Smant (2009) utilized a survey of inflation expectations in four markets: the Euro Area, the US, Switzerland and Sweden and argued that the results emanating from a test of the Fisher hypothesis in the euro area depended largely on the chosen proxy for expected inflation.

It is obvious that the majority of authors (due to the mentioned restrictions) use the current rate of inflation as a proxy for the expected inflation, so the same will be done in this paper.

Regarding the interest rates, there also a few things that should be kept in mind when testing the Fisher effect. Different authors use different measures of interest rates wherein some of them use long-term rates while other use short-term rates. For example, Smant (2011) mentions two reasons to use long-term rates rather than short-term rates. First, short term rates are likely to be directed by monetary policy, whereas the relationship between monetary policy and longer-term interest rates is weak. Second, important financial decisions tend to be medium-term to long-term and should therefore be linked to interest rates of corresponding maturity. In addition, there are authors who use overall interest rates (for example on credits, deposits or Treasury bills yields). For additional review, see Ray (2012), Berument and Jelassi (2002), Berument and Mehdi (2002) Berument and Mehdi (2002), Brhel and Smant (2009), Argyro (2010), Ramadanović (2011) etc.

According to the above discussion, data on non-taxable interest rates on kuna credits indexed to foreign currency (*INF_fc*)[4] is taken as a proxy for nominal interest rate and inflation rate (*INF*)[5] is taken as a proxy for expected inflation. Used data are analyzed on a quarterly base from March 1996 till September 2012 and their movements can be perceived within the following figure.

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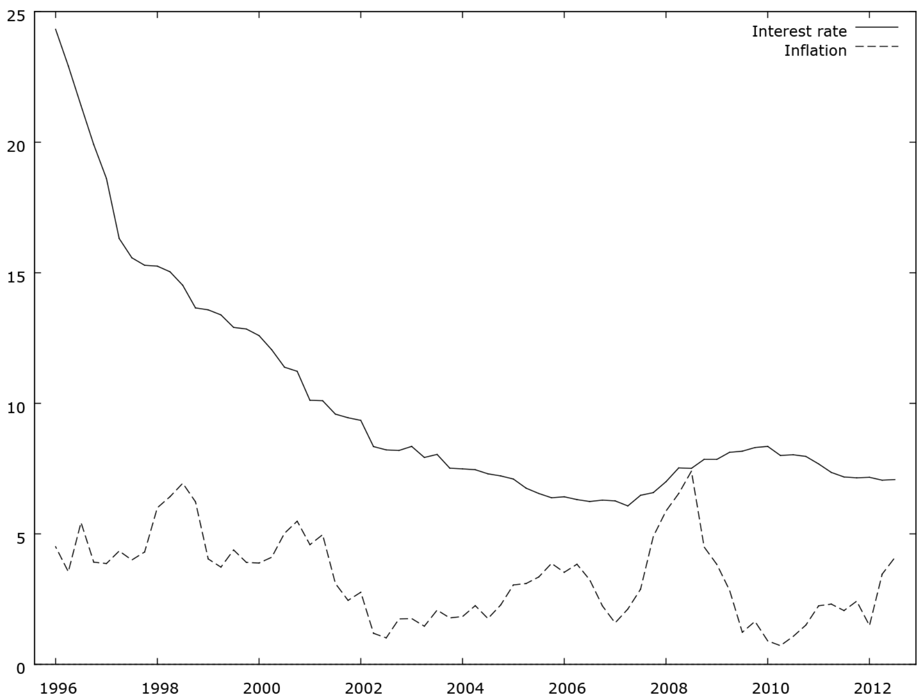
[4] It should be noted that CNB (Croatian National Bank) does not announce the overall average lending interest rate. The largest share of commercial bank loans in Croatia refers to kuna credits indexed to foreign currency and according to this, named interest rate presents the most significant interest rate in Croatian economy. It is important to note that the highest number of loans is granted with variable interest rates which changes according to loan market conditions or bank management decisions. This is very important since the Fisher effect states that nominal interest rates adjust to changes in expected inflation which is not the case if loans are granted with fixed interest rates. Although a number of authors who have analyzed the Fisher effect used Treasury bills yields or some other interest rates under the control of the central bank (benchmark or official rate), Croatian statistics publishes data on Treasury bills only since 2003 and additionally, Croatian money market is undeveloped and therefore a CNB's (Croatian National Bank) role in determining the official interest rate is very limited. Problems steams from money market illiquidity, money market interest rate volatility, under-representation of the money market compared to other sources of funds, the complex interactions between the regulatory costs and market costs of funds, banking contracts which contains clauses on variable interest rates according to the decisions of bank management etc. (HUB, 2007 and HUB 2008). Also, there are other constraints that hinder the formation of the official interest rates and those are related to the fact that Croatia is a small and open economy, highly euroized, import dependent with high external debt. Because of these limitations, this analyses uses interest rate on kuna credits indexed to foreign currency.

[5] Inflation rate is calculated as percent change over corresponding period of previous year.

FIGURE 1. INTEREST RATES ON KUNA CREDITS INDEXED TO FOREIGN CURRENCY (TOTAL AVERAGE) AND INFLATION RATE (SERIES ARE EXPRESSED IN % ON ANNUAL BASIS)



Source: CNB (Croatian National Bank) and IFS (International Financial Statistics).

Until the 2007, interest rate was in fact displaying the downward trend but due to a spillover effect from the global crisis it started to rise afterwards. Inflation has declined until 2002 and then began to rise till 2008 because of rising energy and food prices on the world market, which in turn strongly affected Croatia as an opened and highly import-dependent economy. After 2008 inflation fell again due to a global crisis.

Initially, the growth rate of money supply (M1) was also included in the analysis. However, exclusion and weak exogeneity tests suggested that this variable can be omitted from the model, i.e. it is found to be insignificant in the long-run and short-run[6]. This may point to the fact that short-run and long-run changes in interest rate and inflation are caused by other factors, including those from abroad. For example, changes in the inflation rate in Croatia can be affected by the changes in the price of oil, energy and food prices on international markets (Krtalić and Benazić 2008) and fiscal policy while changes in interest rates can be affected by international capital flows which is further confirmed by the fact that Croatian money market is inefficient, and as such, does not form a official interest rate which in turn affects commercial banks interest rates (HUB 2007, HUB 2008 and Benazić 2012). As already said, just stated is not unusual since Croatia is small and open economy. Finally, there is always a possibility that monetary policy in Croatia

[6] Testing exclusion restrictions in cointegration space for M1 (for rank = 1): LR = 2,7436, df = 1, p-value = 0,0976; Weak exogeneity test for M1 (for rank = 1): df = 1, W = 0,3739, p-value = 0,5409.

operates through several monetary transmission channels and therefore affects prices and other macroeconomic variables in different ways (Erjavec, Cota and Bahovec 1999, Billmeier and Bonato 2002, Erjavec and Cota 2003, Lang and Krznar 2004, Žigman and Lovrinčević 2005, Vizek 2006, Benazić 2009, HUB 2009, Dumičić, Čibarić and Horvat 2010, Benazić and Učkar 2011 etc.).

To determine whether there is a long-run relationship between the variables, it is necessary to apply the concept of cointegration, but prior to testing cointegration and defining the model, it is necessary to observe the degree of integration of variables as well. If nominal interest rates and inflation are indeed driven by nonstationarities, the textbook Fisher relation implies cointegration or that they share a common stochastic trend (Crowder and Hoffman, 1996).

So, in order to test the degree of integration of variables, standard univariate unit-root tests available in all conventional econometric packages and used by the greatest number of authors who investigated the Fisher effect are employed, i.e. ADF test (Dickey and Fuller 1979), PP test (Phillips and Perron 1988) and KPSS test (Kwiatkowski, Phillips, Schmidt and Shin 1992) are considered. Because of the brakes in data, a unit root test with structural break (Saikkonen and Lütkepohl 2002) is employed also[7]. Namely, when structural breaks are present, unit root tests are biased and may lead to false conclusions regarding integration of a time series which can ultimately affect the results of the model. Results of unit root tests are shown in Table 1. Before testing, the variables were seasonally adjusted in order to eliminate the influence of seasonal factors[8].

TABLE 1 - UNIT ROOT TESTS

	Level		***First difference***	
Variable and test	Constant	Constant and trend	Constant	Constant and trend
ADF test		t-stat.		
INT_fc	-9,225439	-5,333599	-2,942911	-5,353975
INF	-2,256586	-2,459368	-5,409715	-5,368786
PP test		Adj. t-stat.		
INT_fc	-7,734221	-4,602585	-3,961426	-5,420338
INF	-2,721868	-2,837255	-7,469838	-7,420318
KPSS test		LM-stat.		
INT_fc	0,818118	0,264665	0,780381	0,154090
INF	0,363499	0,086030	0,049640	0,034896
Test with structural break		t-stat.		
INT_fc	-3,3508	0,1148	-4,0108	-3,4692
INF	-2,2870	-2,4764	-5,4420	-4,7373

Source: Research results

Note: For the implementation of ADF test and test with structural brake the Schwarz Information Criterion (AIC) is used. ADF test critical values (MacKinnon, 1996); constant: 1% level (-3,49), 5% level (-2,89), 10% level (-2,58); constant and trend: 1% level (-4,04), 5% level (-3,45), 10% level (-3,15). PP test critical values (MacKinnon, 1996): constant: 1% level (-3,49), 5% level (-2,89), 10% level (-2,58); constant and trend: 1% level (-4,04), 5% level (-3,45), 10% level (-3,15). KPSS asymptotic critical values (Kwiatkowski-Phillips-Schmidt-Shin, 1992); constant: 1% level (0,739), 5% level (0,463), 10% level (0,347); constant and trend: 1% level (0,216), 5% level (0,146), 10% level (0,119). Critical values for a test with structural brake (Lanne, Lütkepohl and Saikkonen, 2002): constant: 1% level (-3,48), 5% level (-2,88), 10% level (-2,58); constant and trend: 1% level (-3,55), 5% level (-3,03), 10% level (-2,76).

[7] Whereby the following break dates are detected: for interest rate Jun 1997 and March 2001; for inflation December 2008.

[8] Using the Arima X12 method.

Although ADF test and PP test indicate possible stationarity of interest rate in level, KPSS test and test with structural break clearly reject this possibility. Therefore, obtained results and insight into Figure 1 suggests that series are integrated of order I(1), i.e. they are stationary in their first differences. Despite the fact that standard unit root tests are known for their low confirmation power since they do not take into the consideration the variance shift in time series which could significantly affect the empirical results of unit root and cointegration tests (Hamori and Tokihisa, 1997), for the sake of this analysis, it is assumed that the variables under consideration are nonstationary in levels, i.e. they are stationary in their first differences.

IV. METHODOLOGY, EMPIRICAL ANALYSIS AND RESULTS

Engle and Granger (1987) indicated that a linear combination of two or more nonstationary series may be stationary. If so, for these series are said to be cointegrated. This linear stationary combination shows the long-run relationship among the variables and is called cointegrated equation. As already postulated in the introduction, if inflation and interest rate are not cointegrated, we conclude that there is no Fisher effect and proceed no further. Contrary, if cointegration is accepted and if the relationship between inflation and interest rate is positive, then the “partial” Fisher effect holds. When the “partial” Fisher effect holds then the presence of the full” effect can be tested, which, in addition to the “partial” effect, requires testing whether the estimated relationship is in fact one-for-one.

In order to test for cointegration, the methodology proposed by Johansen (1991, 1995) is used. A vector of variables y_t = *(INT_fc, INF)* is defined allowing variables in y_t to be potentially endogenous, i.e. the following unrestricted VAR model is defined:

$$y_t = A_1 y_{t-1} + \cdots + A_p y_{t-p} + Bx_t + \varepsilon_t, \quad \varepsilon_t \approx IN(0, \Sigma) \tag{6}$$

where y_t is a *k*-vector of nonstationary I(1) variables, x_t is a *d*-vector of deterministic variables, ε_t is a vector of independently normally distributed errors with mean zero and covariance matrix Σ, while *A* and *B* are matrices of parameters.

Model (6) can be reformulated into a vector error correction model, i.e. VEC model:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + Bx_t + \varepsilon_t \tag{7}$$

where

$$\Pi = \sum_{i=1}^{p} A_i - I, \quad \Gamma_i = -\sum_{j=i+1}^{p} A_j \tag{8}$$

Granger’s representation theorem asserts that if the coefficient matrix Π has reduced rank $r < k$, then there exist $k \times r$ matrices α and β each with rank r such that $\Pi = \alpha\beta'$ and $\beta' y_t$ is I(0). r is the number of cointegrating relations (the cointegrating rank) and each column of β is the cointegrating vector. The elements of α are known as the adjustment parameters in the VEC model and p is the number of lags.

Number of lags is determined using standard information criteria (AIC, log (FPE), HQ, SC, and FML)[9] and although the criteria indicated different number of lags, the final model is estimated using one lag as suggested by SC and FML criterion.

To determine the existence of cointegrating vector the Johansen's reduced rank procedure is employed whereby *"case 1"* model[10] is tested.

TABLE 2 - JOHANSEN TRACE TEST

Number of cointegration vectors (rank)	*Eingevalue*	*LR trace*	*p-value***
None*	0,6875	82,9357	0,0000
At most 1	0,0893	6,1746	0,1777

Source: Research results

Note: (*) indicates rejection of the hypothesis at the 5% significance level. p-values are taken from several sources including MacKinnon, Haug and Michelis (1999), SVAR simulation (see Harbo, Johansen, Nielsen and Rahbek 1998) and using the method proposed by Boswijk and Doornik (2005).

First column displays the number of cointegrating relations under the null hypothesis, second column displays the ordered eigenvalues of the Π matrix, third column displays the test statistic, while the fourth column represents the probability. The trace statistic tests the null hypothesis of r cointegrating relations against the alternative of k cointegrating relations, where k is the number of endogenous variables, for $r = 0, 1, \ldots, k-1$.

Results in Table 2 clearly suggest the existence of one significant cointegrating vector[11]. Due to breaks in the data, Johansen cointegration test which allows setting breaks (Johansen, Mosconi and Nielsen 2000, Trenkler 2004) is employed with two breaks in levels, i.e. for Jun 1997 and December 2008 whereby *"case 1"* model (*"restricted constant"* model) is tested. Namely, it is well known that breaks in the data might lead to parameter instability and thus may affect the findings about cointegration analysis based on a stable long-run equilibrium relation (Gregory and Hansen 1996).

Again, number of lags is determined using standard information criteria and although the criteria indicated different number of lags, the model is estimated using one lag as suggested by HQ and SC criterion. The results are presented in Table 3 and once again suggest the existence of one cointegrating vector.

9 AIC-Akaike Information Criterion, FPE-Final Prediction Error, HQ-Hannan-Quinn Information Criterion, SC- Schwarz Information Criterion, FML- Fractional Marginal Likelihood Criterion.

10 Since trend was not significant, "case 1" model (i.e. "restricted constant" model) is tested in which the presence of the intercept in the variables in levels is not allowed, while the cointegration relation contains only intercept. As stated by Ramadanović (2011), this specification is consistent with the Fisher effect, as it includes a constant and no deterministic trends.

11 Also confirmed by the analysis of eigenvalues of associated matrix A, where the number of cointegrating vectors (r) is equal to the number of variables in the model (n) minus the number of eigenvalues of the matrix A which are close to one. In this case, r = 2 – 1 = 1.

TABLE 3 - JOHANSEN TRACE TEST WITH STRUCTURAL BRAKES (IN LEVELS ONLY)

Number of cointegration vectors (rank)	*LR trace*	*p-value***
None*	87,39	0,0000
At most 1	3,74	0,8732

Source: Research results

Note: (*) indicates rejection of the hypothesis at the 5% significance level; (**) p-values are taken from Johansen, Mosconi and Nielsen (2000).

Due to structural brakes, three impulse dummy variables for June 1997 (D_1), March 2001 (D_2) and December 2008 (D_3) are included and tested using exclusion test which showed that dummy variables are significant.

Model diagnostic tests shown in Table 4 include joint and individual tests for autocorrelation, ARCH effects and non-normality suggesting that the model is adequately estimated with acceptable characteristics.

TABLE 4 - VECM DIAGNOSTIC TESTS

Serial correlation tests	
LM (1) = 6,1022, p-value = 0,1916 LM (4) = 9,2475, p-value = 0,0552 Ljung-Box test LB (28) = 40,2790, p-value = 0,0624 Portmanteau test (4) = 16,1123, p-value = 0,3066, adjusted test statistic = 16,7914, p-value = 0,2675	
ARCH-LM tests	
VARCHLM (4) = 41,3764, p-value = 0,2475	
ARCH (1): INT_fc = 0,7615, p-value = 0,3829 INF = 0,0100, p-value = 0,9205	ARCH (4): INT_fc = 4,7467, p-value = 0,3143 INF = 1,2990, p-value = 0,8615
Normality tests	
Wald test: Skewness = 0: W(2) = 0,4143, p-value = 0,8129 Kurtosis = 3: W(2) = 0,9358, p-value = 0,6263	Skewness and Kurtosis = W(4) = 1,3501, p-value = 0,8528 Shenton-Bowman/Doornik-Hansen Omnibus test: Skewness and Kurtosis = E_2(4) = 3,6473, p-value = 0,4558
Doornik & Hansen test (lags = 4): Joint test statistic = 1,7264, p-value = 0,7859 Skewness only = 0,6616, p-value = 0,7183 Kurtosis only = 1,0648, p-value = 0,5872	Lütkepohl test (lags = 4): Joint test statistic = 2,4391, p-value = 0,6556 Skewness only = 1,2777, p-value = 0,5279 Kurtosis only = 1,1614, p-value = 0,5595
Jarque-Bera test (lags = 4): u1 = 1,0083, p-value = 0,6040, Skewness = -0,2946, Kurtosis = 3,1399 u2 = 1,3119, p-value = 0,5189, Skewness = -0,1900, Kurtosis = 3,5768	

Source: Research results

Estimated VEC model is normalized by interest rate and is shown in the next table.

TABLE 5 - COINTEGRATION VECTOR AND ADJUSTMENT PARAMETERS

Cointegration vector β			**Adjustment parameters α**		
INT_fc	INF	Constant		Parameters	t-ratios
1	-0,7090	-5,0053	D(INT_fc)	-0,0849	-12,7821
			D(INF)	-0,0038	-0,1683

Source: Research results

"D" indicates the first difference.

Conducted LR tests for exclusion and stationarity in the cointegration space (Johansen 1996) showed that variables in the long-run are significant and non-stationary[12].

The result of cointegration analysis is presented through following long-run relation:

$$INT\_fc_t = 5,0053 + 0,7090 \cdot INF \tag{9}$$

It is evident that an increase in inflation in the long-run leads to an increase in nominal interest rate in Croatia. This result suggests that the "partial" long-run Fisher effect holds. However, the "full" long-run Fisher effect implies that in the long-run nominal interest rate increases one-for-one with inflation and from above equation is obvious that this increase is less then unit. Therefore, in the next step restrictions are imposed on cointegrating vector to test whether the nominal interest rate increases one-for-one with inflation. Likewise, insight into adjustment parameters and their *t*-ratios indicate the presence of weak exogeneity of inflation. Simultaneously, restrictions are imposed on adjustment parameters to test weather inflation can be excluded from the short-run model dynamics. The results are presented in the next table.

TABLE 6 - TESTING RESTRICTIONS ON COINTEGRATING VECTOR AND ADJUSTMENT PARAMETERS

***Restrictions on* β**	**β = *(1,-1)***	**β = *(1,-1)***
Restrictions on **α**	Without restrictions	α = (*,0)
χ2 (1)	1,49165	1,49287
p-value	0,221961	0,474054

Source: Research results

Note: (*) unrestricted.

Results confirm the plausibility of imposed restrictions, i.e. that the "full" long-run Fisher effect in Croatia may hold as well as weak exogeneity for inflation, which means that inflation can be excluded from the short-run model dynamics, but it remains in the long-run equation. Having in mind that interest rates in Croatia are non-taxable, results suggesting the possible "full" long-run Fisher effect in Croatia are in line with Weidmann (1997) who asserted that "full" Fisher effect is attributive only to the economies without such taxes.

Obtained results are very similar with those of Crowder and Hoffman (1996). The results imply that lagged deviations from long-run equilibrium, deviations from the equilibrium interest rate, have no significant effect on future inflation. Therefore interest rates may not be good predictors of future inflation.

According to obtained results, the analysis continues with imposed restrictions through conditional model, i.e. through dynamic short-run model[13]:

$$\Delta INT\_fc_t = a + b_i \Delta INF_t + \alpha ECM_{t-1} + \varepsilon_t \tag{10}$$

where Δ is first difference operator, *a* is constant term, *b* is coefficient estimate, *α* refers to the error-correction term while Ɛ is the error term.

The estimation results are presented in the following table[14].

TABLE 7 - SHORT-RUN CONDITIONAL MODEL

Dependent variable: D(INT_fc)				
Variable	Coefficient	Std. Error	t-ratio	p-value
Constant	0,0147363	0,0378594	0,3892	0,69848
D(INF)	0,0777908	0,0348676	2,2310	0,02943**
D_1	-1,10485	0,263636	-4,1908	0,00009***
D_2	-1,04776	0,253197	-4,1381	0,00011***
ECM_1	-0,0851777	0,00843506	-10,0981	0,00001***

Source: Research results

Note: (***) indicates statistical significance at the 1% level; (**) indicates significance between 1% and 5%; (*) indicates significance between the 5% and 10% levels. "D" indicates the first difference.

The ECM is highly significant and has the expected negative sign indicating a relatively slow adjustment to the long-run equilibrium. Nearly 8,5% of the disequilibria of the previous quarter's

[13] Represents re-evaluated cointegration relation with imposed restrictions on cointegration vector while treating inflation as an exogenous variable. The conditioning of inflation means that inflation remains in the long-run i.e. in cointegration vector, while the same is excluded from the short-run dynamics of the model.

[14] Dummy variable D_3 was insignificant and was omitted from the conditional model.

shock adjust back to the long-run equilibrium in the current quarter. Additionally, the changes in inflation have a statistically significant and positive impact on the nominal interest rate, although this impact is very small. It can be concluded that inflation affects nominal interest rate in the long-run and short-run, while this effect is much stronger in the long-run. This further implies the possible existence of the "full" long-run and "partial" short-run Fisher effect in Croatia which is in accordance with the fact that the Fisher effect is a phenomenon that appears in the long-run and may not be present in the short-run. According to this, Croatian banks adjust their interest rates in line with inflation changes. *Ditto*, the monetary authorities, i.e. the CNB should employ strategies to keep inflation at low levels in order to not discourage the borrowing if nominal interest rates grow because of inflation. Due to the fact that Croatia is a small and open economy, highly euroized, import dependent with high external debt, this could be done by keeping the exchange rate stable.

Finally, conditional model diagnostic tests presented in the following table include joint and individual tests for autocorrelation, ARCH effects and non-normality suggesting that the model is adequately estimated with acceptable characteristics.

TABLE 8 - SHORT-RUN CONDITIONAL MODEL DIAGNOSTIC TESTS

Serial correlation	LM test: LMF = 0,781781, p-value = 0,542 Alternative statistic: TR^2 = 3,437729, p-value = 0,487 Ljung-Box test: Q' 3,21071, p-value = 0,523
Normality	χ2(2) = 0,254, p-value = 0,88060
Heteroscedasticity	White's test: LM = 6,45043, p-value = 0,488245 White's test (squares only): LM = 6,27733, p-value = 0,392851 Breusch-Pagan test: LM = 4,69607, p-value = 0,319928 Breusch-Pagan test (Koenker robust variant): LM = 6,210290, p-value = 0,183984
ARCH	LM = 8,60743, p-value = 0,071697
Structural break	QLR test: max F(3, 57) = 3,72793 occurs at observation 1998:3, p-value = 0,1422 Chow test for structural break at observation 2004:1, F(3, 57) = 2,72631, p-value 0,0524
Parameter stability	CUSUM test: Harvey-Collier t(59) = -1,59224, p-value = 0,116675
Non-linearity	Ramsey's RESET test: F(2, 58) = 2,55233, p-value = 0,0866221 Squared terms: TR^2 = 4,83336, p-value = 0,0892171

Source: Research results

Although the analysis showed the possible existence of the long-run Fisher effect in Croatia, there are some limitations to bear in mind. Firstly, this analysis is based on the proxy variables for nominal interest rate and expected inflation. Secondly, time period used in the analysis is relatively short and includes structural breaks. Due to a significant break in 2008, there is a possibility that after 2008 nominal interest rate and inflation do not achieve common trends or are no longer moving under the principle of one-to-one, as proposed by Fisher (see Figure 1). Therefore, the future research of the Fisher effect in Croatia should take into account just stated limitation by splitting the analysis into two parts, i.e. till 2008 and after, as similarly done by Piccinino (2011). And finally, most of the standard unit root tests have low confirmation power since they do not take into consideration the variance shift in time series which could significantly affect the empirical results of unit root and cointegration tests (Hamori and Tokihisa, 1997). Despite these limitations, this analysis can serve as a base for further research regarding the Fisher effect in Croatia.

IV. CONCLUSION

The main goal of this paper was to test the Fisher effect in Croatia using the cointegration model. The Fisher effect implies that a percentage increase in the expected rate of inflation would lead to a percentage increase in the nominal interest rates assuming that real interest rate is constant. According to this, the monetary authorities should employ strategies that will prevent inflation from rising if nominal interest rates are to be kept at low levels in order to not discourage the borrowing.

The Fisher effect is tested using the VEC model. Results suggest that inflation affects nominal interest rate in the long-run (under the principle of one-to-one) pointing to the fact that in the long-run the Fisher effect in Croatia may hold. Since imposed restrictions on cointegrated vector and adjustment parameters showed that inflation is weakly exogenous, and thus can be excluded from the short-run model dynamics, the analysis is conducted through conditional model, i.e. through dynamic short-run model showing that changes in inflation have a statistically significant and positive impact on the nominal interest rate, although this impact is very small. Therefore, it can be concluded that inflation affects nominal interest rate in the long-run and short-run, however this effect is much stronger in the long-run further implying the possible existence of the "full" long-run and the "partial" short-run Fisher effect in Croatia. Hence, Croatian banks adjust their interest rates in line with inflation changes. Non-the-less, Croatian monetary authority i.e. the CNB should keep inflation at low levels due to an unstable macroeconomic situation.

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