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MONEY DEMAND IN AN EU ACCESSION COUNTRY: A VECM STUDY OF CROATIA

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ABSTRACT

The paper estimates the money demand in Croatia using monthly data from 1994 to 2002. A failure of the Fisher equation is found, and adjustment to the standard money-demand function is made to include the inflation rate as well as the nominal interest rate. In a two-equation cointegrated system, a stable money demand shows rapid convergence back to equilibrium after shocks. This function performs better than an alternative using the exchange rate instead of the inflation rate as in the 'pass-through' literature on exchange rates. The results provide a basis for inflation rate forecasting and suggest the ability to use inflation targeting goals in transition countries during the EU accession process. Finding a stable money demand also limits the scope for central bank 'inflation bias'.

Keywords: accession, error correction, Fisher equation, money demand, multivariate cointegration
JEL classification numbers: O42, E13, E41, E51

I. INTRODUCTION

Neoclassical money-demand functions underlie much theoretical and empirical work. Typically, the nominal interest rate is the price of

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money, and the income velocity of money moves in conjunction with this rate. This is as in Friedman's (1956) restatement of money-demand theory, although it contrasts with the institutionally fixed velocity in Fisher's (1911) quantity theory. Similar to Fisher, velocity has often been assumed to be exogenous (Lucas, 1980; Ireland, 1996; Alvarez *et al.*, 2001). Similar to Friedman, others have endeavoured to explain velocity and related phenomena within the model (Hodrick *et al.*, 1991; Eckstein and Leiderman, 1992; Ireland, 1995; Lucas, 2000; Gillman and Kejak, 2004, 2005).

Empirical work on money demand has focused on interest rate explanations as in the constant semi-interest elasticity model of Cagan (1956) (Eckstein and Leiderman, 1992; Mark and Sul, 2003) or the constant interest elasticity model of Baumol (1952) (Hoffman and Rasche, 1991; Hoffman *et al.*, 1995; Lucas, 2000). Apparent instability in empirical money-demand functions was found because of 'shifts' in demand in the 1980s; for example, Friedman and Kuttner (1992) found a break in cointegration around 1980. This instability literature was met with an effort to include, within the money-demand function, the prices of substitutes for money that may have been subjected to large changes and that may have caused money demand without these substitute prices to appear unstable. In particular, interest earning accounts with demand deposits that could be used in exchange, or 'exchange credit', were used to avoid the high inflation tax of the 1980s and seemed to cause a shift in money demand. Including proxies for financial service innovation led to renewed results of stable money-demand functions, even including the period of the big financial deregulations (Friedman and Schwartz, 1982; Gillman *et al.*, 1997; Gillman and Otto, 2003).

Money demand has become less visible in the policy debate because of interest in Taylor (1999)-type rules. The focus on nominal interest rate instruments has bred the perception of policy irrelevance of money-demand theory and the use of monetary aggregates. However, McCallum (1999) has disputed such conclusions by emphasizing that money demand and the use of rules based partly on money aggregates are being disregarded to the detriment of the ultimate monetary policy results. Alvarez *et al.* (2001) further advance the importance of money aggregates by providing a general equilibrium basis for the equivalence between interest rate rules and money supply rules. Similarly, Schabert (2005) establishes a liquidity effect in a general equilibrium neoclassical monetary model, in which there is also a direct relation between the money supply growth rate and the nominal interest rate. And empirical money-demand work has recently become more prominent in the central banks of developed nations (e.g., the euro-area studies of Brand and Cassola, 2004; Brand *et al.*, 2002; Kontolemis, 2002).

Developing nations tend to rely more on discretion rather than rules and often justify this just as central banks in developed nations did in the past: the money-demand function is unstable. This sort of discretion instead of rules can lead to an 'inflation bias' of the type described by Kydland and Prescott (1977). Empirically, evaluating the stability of money demand still remains a challenge in developing countries because of lack of confidence in the data quality and because of the many major changes that continue to occur in such economies.

In this paper, the key extension to a standard money-demand function results from an investigation of whether the Fisher equation of interest rates holds.[1] The myriad ways in which an unexpected acceleration or deceleration of the inflation rate can affect the real interest rate makes suspect the standard Fisher (1930) relation that underlies classical money-demand functions. In those, changes in the inflation rate are directly reflected in the nominal interest rate. But if this is not true, which can be a likely scenario in a transition country, then the standard money-demand function requires modification from only including the nominal interest rate as the price of money.

With an extended money-demand specification, the paper shows that a stable money-demand function can be found for Croatia, despite tumultuous changes there over the transition period. This presents a good case study in that the finding of a stable money demand may be surprising. Both the emphasis of the Croatian central bank on the exchange rate in its monetary policy and the high fraction of private foreign exchange use in the country have led to the expectation that Croatian money demand is unstable (see Kraft, 2003). However, with a money demand that accounts for failure of the Fisher relation, a stable function is estimated with vector error correction model (VECM) techniques using monthly International Financial Statistics (IFS) data from 1994 to 2002. Over this period the data are reliable, and several robustness checks are conducted, including a focus on exchange rates within the money-demand function.

The data begin only after the Croatian hyperinflation of 1993 and near to the beginning of the issuing of the new Croatian currency, providing confidence in the data. The data's stationarity and seasonal properties are tested carefully (Section III). After finding the Croatian income velocity of money non-stationary (Section IV.1), in contrast to Fisher's (1911) concept, the paper focuses on whether the Fisher (1930) equation of interest rates holds in Croatia. Researchers such as Baba *et al.* (1992) have included the inflation rate as well as the nominal interest rate in the money-demand function. This strategy is justified here in that evidence suggests a failure of the long-run Fisher relation in which the nominal interest rate and inflation rate move together and are

[1] For example, see Crowder's (2003) panel testing of this equation; see also Brand and Cassola (2004) for an alternative multi-equation money-demand approach that includes a Fisher equation.

interchangeable in the money-demand function (Section IV.2). This extension of money demand to include the inflation rate along with the nominal interest rate, and hence capture deviations from the Fisher equation, constitutes the baseline model (Section IV.3).

Petrovic and Mladenovic (2000) estimate Yugoslavian money demand using the exchange rate rather than the inflation rate or the nominal interest rate. The idea is that exchange rates reflect the inflation rate changes as in the uncovered interest rate parity concept (see, for example, Walsh, 2003). This approach to money demand is sometimes used to support a monetary policy of exchange rate targeting even when the goal is to decrease the inflation rate. As part of the robustness investigation, the paper compares this alternative money-demand approach to the baseline model (Section IV.3.2).

While data limitations in terms of the length of the time series are an important qualification, implications can still be cautiously deduced. A stable money demand is useful because it suggests the variables that can be used to forecast inflation. And all central banks appear to engage in inflation rate forecasting as one of their crucial tasks. Croatia in 2001, along with Hungary in 2001 and Poland in 1997, established new central bank chartering acts that state price stability as the primary goal of the central bank. Croatia has recently had very low inflation, and low inflation in Croatia remains the goal, even if it may be using the exchange rate as its primary instrument.

The results show that the inflation rate enters a stable money-demand function that exhibits fast readjustment to shocks. This suggests that an inflation targeting policy (Svensson, 1999) will not 'de-stabilize' money demand. In contrast to the baseline model, including the exchange rate instead of the inflation rate yields a near-zero adjustment to shocks. This implies that an exchange rate targeting strategy may induce a perceived instability in the money-demand function if, for example, such a policy results in substantial inflation-rate volatility that keeps the money demand constantly readjusting (Section V).

II. CROATIAN MONEY, POLICY AND BANKING BACKGROUND

We first consider some descriptive facts about Croatian real money use, nominal interest rates and the inflation rate over the 1994–2002 period. These help indicate whether it is likely that a classic money-demand function will be operative. The money aggregate M1 comprises the new Croatian kuna currency, as of 30 May 1994, and kuna-denominated demand deposits.[2] Figure 1 shows that the quantity of real money

[2] The kuna replaced the Croatian dinar that had been introduced on 23 December 1991, when Croatia became an independent state.

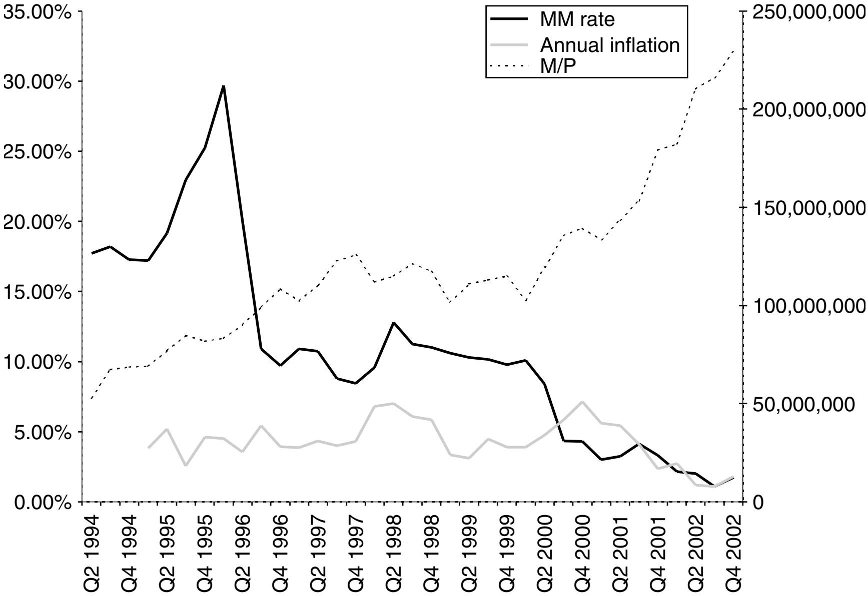

Fig. 1. Real money, inflation and nominal interest rate.

(M/P), defined here in terms of M1, and the nominal interest rate (money market rate (MM)) move inversely as in a classical money-demand function. However, the figure also shows that the inflation rate and nominal interest rate do not move together as in a Fisher equation.

Between 1994 and 2001, the inflation rate was fairly stable around 5 percent; it then moved downwards steadily towards very low levels by 2003, and it has remained in the 1.5 percent range. With such low rates, the Croatian National Bank has begun succeeding in its 'primary objective to achieve and maintain price stability' (2001 National Bank Act). This low inflation has been achieved while the Bank has been described as being engaged in 'strict exchange rate targeting' (Billmeier and Bonato, 2004). Or as Kraft (2003) puts it, 'The main intermediate target is the exchange rate, not any monetary aggregate. In that respect, Croatia's monetary policy resembles an exchange rate fix more than a float of any sort' (p. 14). These different perspectives suggest that the exchange rate may have been an important instrument in the Bank's realization of its low inflation goal.

An interesting banking aspect of the M1 aggregate can be seen in Figure 2. Currency constitutes the lion's share of M1. The demand deposit to currency ratio averages well below 1. In comparison, for example, the US demand deposit to currency ratio has trended downwards steadily from 4 in 1959 to near 1 in 2002. Low- or non-interest-bearing demand deposits have been used significantly less in Croatia than in the USA.

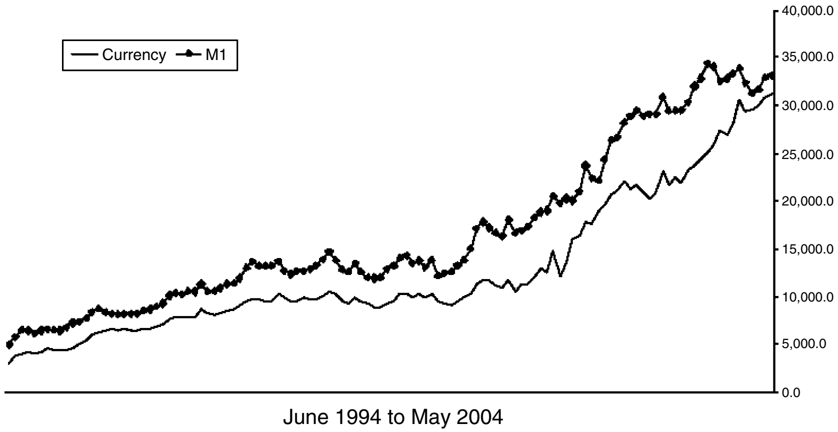

Fig. 2. Croatian monetary aggregates.

Another banking feature is that there have been significant foreign-currency-denominated deposits, now primarily in euros. These deposits have accounted for some 75 percent of total new deposits (Kraft, 2003). Kraft suggests that these holdings may imply a 'lack of credible monetary policy', adding that Croatians have a 'habit of saving in foreign exchange' (p. 4). Such a habit can be because of inflation avoidance and, in addition, may reflect a lesser use of banking for exchange purposes.

The commercial bank sector has seen significant turmoil. The banks started out as state owned and have gradually become privatized in the face of many disruptions to activity. Stringent restrictions have been imposed on the banks at times, for example, with reserve requirements as high as 31 percent during the war period of 1995 and with punitive levels of reserves if bank credit exceeded a certain threshold in recent years. Deregulation and liberalization started in earnest in 1996, when the government received an investment grade rating on its debt and a large commercial bank consolidation took place. A crisis occurred in 1998–99 with some bank insolvencies, and bank reform acts were passed in 1999 and 2002. Restructuring and privatization were largely finished by 2001.[3]

[3] For example, three banks, Bjelovarska Banka, Trgovacka Banka and Cakovecka Banka, were merged into Erste and Sleiermarkische Bank in September 2000 to make one of the ten largest banks in Croatia. Erste then bought 85 percent of Rijecka Banka in April 2002 and merged it with Erste and Sleiermarkische Bank in August 2003 to make the third largest bank group in Croatia. Another example is Slavoska Banka, which started in 1955 and sold some 35 percent of its shares in 1999 to the EBRD and Hypo Alpe Adria Bank. Zagrebacka Banka was the first Croatian bank to be registered as a joint stock company, with limited liability, in 1989, the first bank rated by the three major international rating companies in 1997, and a bank that recently accounted for 25 percent of total Croatian banking assets. It partnered internationally in 2002, with UniCredito Italiano and Allianz.

A weak, gradually reforming commercial bank sector could help explain the greater use of currency and foreign-exchange-denominated deposits. But a lesser use of banking does not necessarily threaten the stability of money demand. The main factors that have caused large shifts in money demand in developed countries have been the big, sudden financial deregulations of the 1980s (Gillman and Kejak, 2004; Benk *et al*., 2005). Such changes in financial sector productivity can be incorporated in money-demand functions to stabilize an otherwise seemingly unstable money-demand function, as Gillman and Otto (2003) show in time series estimations for the USA and Australia. However, deregulation has been gradual in Croatia and inclusion of financial sector variables in the money-demand function appears less necessary.

III. DATA AND DESCRIPTIVE ANALYSIS

The data used in the estimation are IFS time series with monthly frequency and seasonal adjustment: industrial production for the output variable, M1 money, consumer prices, a Croatian kuna (HRK)–euro exchange rate and the money market interest rate (Table 1).

The variables are in natural logarithms of the indices with base year 1995. The data span is from April 1994 till August 2002 for all series, which are plotted in Figure 3 along with velocity (output divided by real money).

III.1 Seasonal unit root tests

To ensure that the use of seasonally adjusted data is appropriate, we first consider Figure 4, which compares seasonally unadjusted series with seasonally adjusted series. Only modest differences emerge. However, it is useful to test whether explicit modelling of seasonality is requisite. In particular, if the series are stochastic and there exist

TABLE 1
Definition of variables

Symbol	*Definition*
i_t	Industrial production
m_t	Money
$(m-p)_t$	Real money
p_t	CPI prices
Δp_t	Inflation
ex_t	HRK–euro exchange rate
r_t	Money market interest rate
v_t	Money velocity ($p_t + y_t - m_t$)

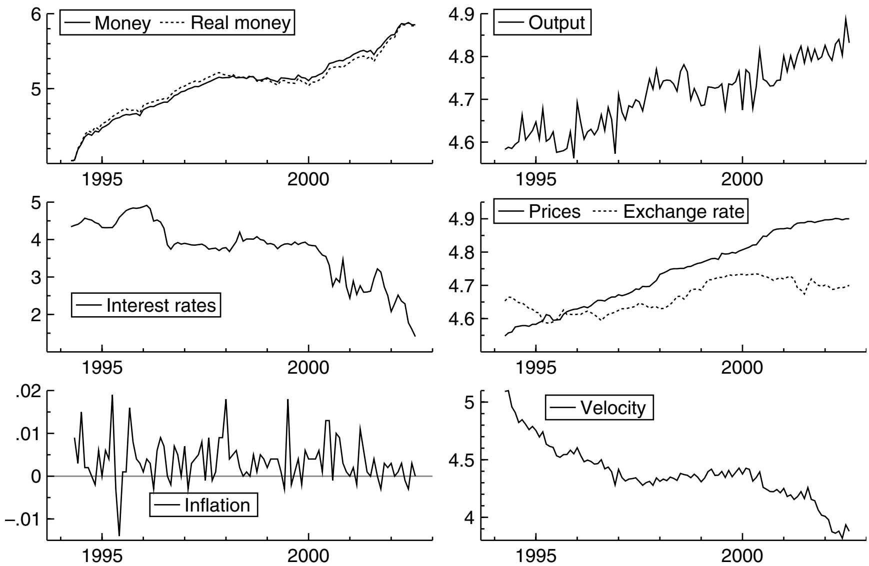

Fig. 3. Money-demand variables.

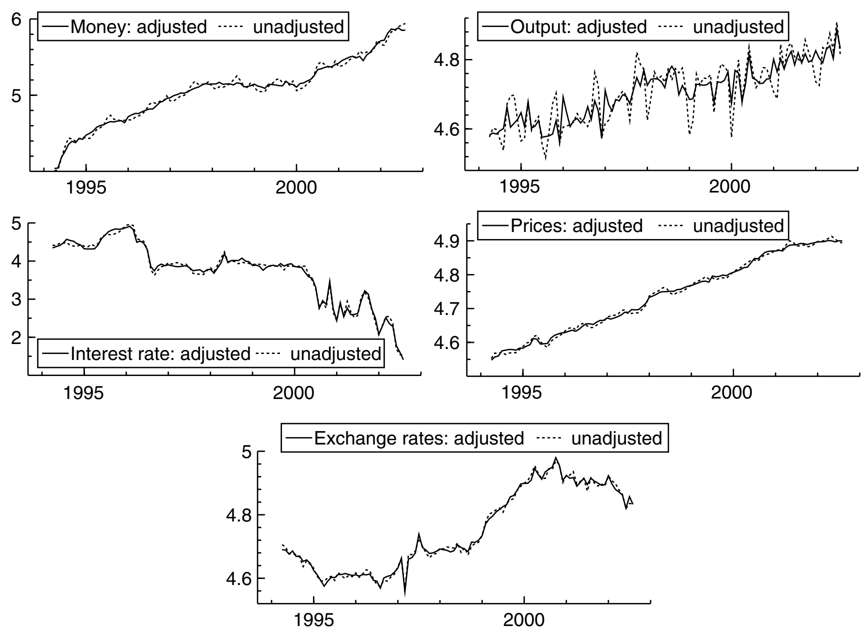

Fig. 4. The effect of seasonal adjustment.

seasonal unit roots, then these unit roots would need to be adjusted for through seasonal differencing (Davidson *et al.*, 1978; Dickey *et al.*, 1984; Beaulieu and Miron, 1993; Canova and Hansen, 1995).

The particular test used for seasonal unit roots is that of Hylleberg *et al.* (1990) as adapted to monthly data by Franses (1990), based on the following ordinary least squares (OLS) regression:

$$\begin{aligned}\Delta_{12}y_t = \; & \pi_1 y_{1,t-1} + \pi_2 y_{2,t-1} + \pi_3 y_{3,t-1} + \pi_4 y_{3,t-2} + \pi_5 y_{4,t-1} \\ & + \pi_6 y_{4,t-2} + \pi_7 y_{5,t-1} + \pi_8 y_{5,t-2} + \pi_9 y_{6,t-1} + \pi_{10} y_{6,t-2} \\ & + \pi_{11} y_{7,t-1} + \pi_{12} y_{7,t-2} + \sum_{i=1}^{12} \alpha_i D_{it} + \gamma t + u_t,\end{aligned}$$

where

$$\begin{aligned}y_{1,t} &= (1+L)(1+L^2)(1+L^4+L^8)y_t,\\ y_{2,t} &= -(1-L)(1+L^2)(1+L^4+L^8)y_t,\\ y_{3,t} &= -(1-L^2)(1+L^4+L^8)y_t,\\ y_{4,t} &= -(1-L^4)(1-\sqrt{3}L+L^2)(1+L^2+L^4)y_t,\\ y_{5,t} &= -(1-L^4)(1+\sqrt{3}L+L^2)(1+L^2+L^4)y_t,\\ y_{6,t} &= -(1-L^4)(1-L^2+L^4)(1-L+L^2)y_t,\\ y_{7,t} &= -(1-L^4)(1-L^2+L^4)(1+L+L^2)y_t.\end{aligned}$$

The t-tests for the significance of the coefficients are given in Table 2, which can be compared to the critical values tabulated by Franses (1990). The π_1 coefficients are below their 95 percent critical values indicating that a unit root hypothesis cannot be rejected at the zero frequency, using the standard Dickey–Fuller tests. Yet the existence of seasonal unit roots is rejected for all π_i coefficients. Note that seasonal dummies and a time trend are included in the test regressions. These results together indicate that there is a stochastic trend within the series and that seasonality is deterministic. This means that seasonality need not be modelled explicitly. Using seasonally adjusted data directly, without having to remove any seasonal unit roots, allows us to save degrees of freedom with a limited data set.

III.2 Unit root tests of seasonally adjusted series

Augmented Dickey–Fuller (ADF) unit root tests for the order of integration (Table 3) do not reject the hypothesis that the tested series have a unit root and are thus I(1). The ADF tests were performed by considering all options regarding deterministic components (i.e., trend and constant), and in all cases the unit root hypothesis could not be rejected. Additional

TABLE 2
Seasonal unit root tests (t-values)

Coefficient	m_t	p_t	i_t	ex_t	r_t
π_1	−2.997	−1.647	−1.798	−1.685	−1.891
π_2	−4.963	−4.069	−1.536	−4.959	−3.392
π_3	−2.649	−2.584	−3.845	−2.653	−2.097
π_4	−5.337	−4.693	−3.967	−3.419	−3.392
π_5	−11.503	−8.545	−6.766	−10.734	−8.241
π_6	−5.393	−4.171	−3.356	−3.230	−4.063
π_7	−3.640	−3.709	−5.500	−5.289	−3.634
π_8	−12.289	−9.077	−6.733	−11.215	−8.508
π_9	−8.902	−6.795	−5.670	−8.720	−6.821
π_{10}	−4.684	−3.728	−2.932	−4.827	−3.607
π_{11}	−4.500	−4.257	−5.303	−5.577	−4.343
π_{12}	−12.729	−8.328	−4.694	−11.143	−7.838

ADF tests on first differences find strong rejection of the unit root null in all series.

The inflation rate series deserves careful consideration, in that evidence on the integration order of the inflation rate tends to be mixed between unit root and stationarity findings (Culver and Papell, 1997; Benati and Kapetanios, 2003). Perron (1989)-type tests for structural breaks can

TABLE 3
Augmented Dickey–Fuller unit root tests

Variable	t-ADF	$\beta(y_{t-1})$	$\hat{\sigma}$	j*	$t - \Delta y_{t-j}$	p value
†m_t	−1.815	0.937	0.026	9	3.333	0.001
‡m_t	0.166	1.002	0.026	9	3.029	0.003
†p_t	−1.542	0.883	0.005	5	−1.847	0.068
‡p_t	−1.018	0.995	0.005	4	−1.847	0.068
†i_t	−2.326	0.589	0.032	9	2.097	0.039
‡i_t	−0.644	0.966	0.033	2	−4.708	0.000
†$(m-p)_t$	−1.794	0.939	0.026	9	3.623	0.001
‡$(m-p)_t$	0.113	1.002	0.026	9	3.299	0.002
‡Δp_t	−6.999	0.038	0.005	1	−0.125	0.901
†ex_t	−1.033	0.966	0.007	2	−2.183	0.032
‡ex_t	−1.725	0.971	0.007	2	−2.333	0.022
†r_t	−1.485	0.893	0.195	5	1.975	0.052
‡r_t	0.713	1.023	0.199	5	1.707	0.092

*Highest significant lag in the ADF regression.
†Trend and constant included; 5% c.v. = −3.461, 1% c.v. = −4.066.
‡Constant included; 5% c.v. = −2.895, 1% c.v. = −3.507.

indicate if an apparent unit root is break-adjusted stationary. While such an investigation is limited within the short time periods available for transition countries, the Croatian inflation rate (Δp_t) does not appear to be trending (Figure 1). The visual impression is further confirmed by the unit root tests (Table 3), which strongly reject the null up to the fifth lag in the ADF regression (as no lagged differences are significant, a simple DF test suffices; the t-DF value is -9.653, with $\beta(y_{t-1}) = 0.026$).

IV. ECONOMETRIC MODELLING

IV.1 The income velocity of money

The observed downward trend in velocity in Figure 3 may be deterministic or stochastic. A stochastic trend can be tested for using an unrestricted vector autoregression (VAR) in levels. The resulting VECM system is given in equation (1), and the results are summarized in Table 4.[4]

$$\begin{pmatrix} \Delta m_t \\ \Delta y_t \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \end{pmatrix} + \sum_{i=1}^{11} \begin{pmatrix} \gamma_{11}^{(i)} & \gamma_{12}^{(i)} \\ \gamma_{21}^{(i)} & \gamma_{22}^{(i)} \end{pmatrix} \begin{pmatrix} \Delta m_{t-i} \\ \Delta y_{t-i} \end{pmatrix} + \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{pmatrix} \begin{pmatrix} \beta_{11} & \beta_{12} \\ \beta_{21} & \beta_{22} \end{pmatrix}' \begin{pmatrix} m_t \\ y_t \end{pmatrix} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix}. \tag{1}$$

The Johansen (1995) cointegration tests suggest that there is one cointegrating vector between money and output. Both λ-max and λ-trace statistics are above their 95 percent critical values with λ-max being significant at the 1 percent level. The (first) cointegrating vector including coefficients of m_t, y_t and t (a time trend) is estimated as $\boldsymbol{\beta}' = (1, -6.6, 0.01)$ with the accompanying adjustment coefficient vector $\alpha = (0.05, -0.02)$. This implies a long-run relationship $p_t + 6.6y_t - m_t$. Imposing the restriction[5] $\boldsymbol{\beta}' = (1, -1, *)$ results in the estimated trend coefficient of -0.05 and $\alpha = (-0.03, 0.1)$, which, however, is strongly rejected by the likelihood ratio (LR) $\chi^2_{(1)}$ of 22.3. It is clear then that the restriction $\boldsymbol{\beta}' = (1, -1, 0)$, being even more restricted, cannot hold either (which is confirmed by the highly significant LR $\chi^2_{(2)}$ test of 24.64).

The findings imply that $\upsilon_t \sim \mathrm{I}(1)$ regardless of the presence of a deterministic trend in the cointegration space. That is, an apparently systematic decline in the money velocity is in fact stochastic, and no fixed per annum percent decline or deterministic downward trend can be

[4] The lag length of the VAR was determined by sequential testing for the validity of the system's reduction, starting with 12 lags (i.e., 1 year of data) and reducing one lag at a time. The reduction from 12 to 11 lags was not rejected, while all further reductions were strongly rejected by the system reduction F-tests.

[5] The asterisk implies an unrestricted coefficient.

TABLE 4
*Johansen cointegration tests: VAR(11) with $y' = (m_t, y_t)$**

H_0: $r = p$	λ-*max*	*95% CV*	λ-*trace*	*95% CV*
$p = 0$	29.53	19.0	25.79	25.3
$p \leq 1$	4.60	12.3	4.60	12.3

*Eigenvalues: $\lambda_1 = 0.280$, $\lambda_2 = 0.050$.

claimed. It follows that long-run stability of the money-demand equation requires consideration of additional variables such as the interest, inflation or exchange rates.

IV.2 The Fisher equation

Denoting the nominal interest rate in period t by r_t, the real interest rate by ρ_t and inflation by Π_t, the Fisher equation (Dimand, 1999; see also Fisher, 1930) can be written as $r_t = \rho_t + \Pi_t$. With the additional assumption that $\rho_t = \hat{\alpha} + \hat{\varepsilon}_t$ (i.e., real interest rate is constant), where $\hat{\varepsilon}_t$ is independently and identically distributed (i.i.d.), the Fisher equation becomes $r_t = \hat{\alpha} + \Pi_t + \hat{\varepsilon}_t$, which implies independence of the real interest rate and inflation. The equation is usually estimated in log levels as $\ln(r_t) = \hat{\beta}_0 + \hat{\beta}_1 \ln(\Pi_t) + \hat{u}_t$, and a test of the restriction $\hat{\beta}_1 = 1$ is taken to be the test of the (long-run) validity of the Fisher equation. The constant $\hat{\beta}_0$ can be interpreted as the long-run equilibrium real rate of interest. Note that when the variables are in logarithms, inflation measured as $\ln(\Pi_t) = \ln(p_t/p_{t-1})$ is equivalent to a simple difference in the log of the price index, i.e., $\Delta\ln(p_t)$. Hence, the Fisher equation can be stated as[6]

$$\ln(r_t) = \hat{\beta}_0 + \hat{\beta}_1 \Delta \ln(p_t) + \hat{u}_t, \quad \hat{u}_t \sim \text{i.i.d.}, \quad \hat{\beta}_1 = 1. \tag{2}$$

Initially ignoring the order of integration, the estimated equation is

$$\ln(r_t) = \underset{(0.10)}{3.65} + \underset{(15.46)}{20.03}\, \Delta \ln(p_t),$$

where standard errors are in parentheses and $R^2 = 0.017$, $\hat{\sigma} = 0.783$ and DW $= 0.081$. It is evident that the null hypothesis $H_0: \hat{\beta}_1 = 0$ cannot be rejected, and in addition a low Durbin–Watson statistic implies dynamic misspecification. The ADF unit root test on $\hat{u}_t$ produced a t-value of 0.519 where the highest significant lag is 4, which clearly cannot reject that $\hat{u}_t \sim I(1)$. Note that this can also be inferred from the fact that $\ln(r_t) \sim I(1)$ while $\Delta\ln(p_t) \sim I(0)$; therefore, it must be that, for all γ, $\ln(r_t) - \gamma\Delta\ln(p_t) \sim I(1)$.

[6] An alternative version of the Fisher equation, given constant money velocity, is $\Delta m_t = \Delta p_t$ (see, for example, Monnet and Weber, 2001). This, however, is not suitable for the cases where velocity is not constant.

Alternatively, estimation of Sargent's (1972) extended Fisher equation, with $n = m = 3$, yields

$$\ln(r_t) = \underset{(0.45)}{13.74} - \underset{(1.15)}{2.17}\ln(m_t) - \underset{(1.62)}{2.16}\ln(m_{t-1}) - \underset{(1.61)}{1.08}\ln(m_{t-2}) + \underset{(1.05)}{3.46}\ln(m_{t-3})$$
$$+ \underset{(6.23)}{6.44}\Delta\ln(p_t) + \underset{(6.12)}{0.74}\Delta\ln(p_{t-1}) + \underset{(6.13)}{0.70}\Delta\ln(p_{t-2}) - \underset{(6.15)}{2.68}\Delta\ln(p_{t-3})$$

with $R^2 = 0.863$, $\sigma = 0.305$ and DW $= 0.570$. Here, while the Durbin–Watson statistic is still indicative of some remaining residual autocorrelation, the fit is improved and the residuals are close to stationary.[7] However, inflation is not significant at any lag. This is also seen in the long-run solution

$$\ln(r_t) = \underset{(0.45)}{13.74} - \underset{(0.08)}{1.95}\ln(m_t) + \underset{(12.95)}{5.19}\Delta\ln(p_t),$$

where Wald $\chi^2_{(2)} = 548.32$, which is highly significant. Individually, only the money variable is significant; inflation is not. Similar results are obtained by estimating the distributed lag version of the Fisher equation (Sargent, 1973), which is specified as a special case of the 'extended' equation, i.e., $\ln(r_t) = \tilde{\alpha} + \sum_{i=1}^{m} \tilde{v}_i \Delta\ln(p_{t-i}) + \breve{\varepsilon}_t$. Estimation of this equation produces insignificant coefficients of inflation at all lags (including up to 12 lags) and similarly insignificant long-run coefficients (not shown). In addition, the residuals are non-stationary which confirms the previous conclusion about the integration orders.

Alternatively, following Crowder and Hoffman (1996) and Crowder (1997), we can consider a bivariate VECM system using the Johansen technique. The specification is

$$\begin{pmatrix} \Delta r_t \\ \Delta^2 p_t \end{pmatrix} = \begin{pmatrix} \tau_1 \\ \tau_2 \end{pmatrix} + \sum_{i=1}^{11} \begin{pmatrix} k_{11}^{(i)} & k_{12}^{(i)} \\ k_{21}^{(i)} & k_{22}^{(i)} \end{pmatrix} \begin{pmatrix} \Delta r_{t-i} \\ \Delta^2 p_{t-i} \end{pmatrix}$$
$$+ \begin{pmatrix} \chi_{11} & \chi_{12} \\ \chi_{21} & \chi_{22} \end{pmatrix} \begin{pmatrix} \theta_{11} & \theta_{12} \\ \theta_{21} & \theta_{22} \end{pmatrix}' \begin{pmatrix} r_{t-1} \\ \Delta^2 p_{t-1} \end{pmatrix} + \begin{pmatrix} \eta_{1t} \\ \eta_{2t} \end{pmatrix}.$$

Estimation of this system produces eigenvalues of $\lambda_1 = 0.109$ and $\lambda_2 = 0.055$; the λ-max and λ-trace statistics are 10.31 and 15.38, respectively, which are well below their 95 percent critical values of 19 and 25.3.[8] These results imply that the interest rate and the inflation rate are not cointegrated. The long-run Fisher equation does not hold.

The above approaches to testing the Fisher equation have the problem of the integration order of interest rates and inflation variables,

[7] The ADF t-value was -2.637 with seven lags included in the regression, which is above the 1 percent critical value of -2.591 for the regression without trend or constant.

[8] A linear trend was included in the cointegrating space.

because the Croatian inflation is I(0).[9] To avoid the integration order problems and consistently estimate $\hat{\beta}_1$ from the Fisher equation, $\ln(r_t) = \hat{\beta}_0 + \hat{\beta}_1 \Delta \ln(p_t) + \hat{u}_t$, consider the OLS estimator[10]

$$\tilde{\beta}_1 = \frac{\sum_{t=1}^{T} \Delta^2 \ln(p_t) \Delta \ln(r_t)}{\sum_{t=1}^{T} [\Delta^2 \ln(p_t)]^2}.$$

It can be shown that $\tilde{\beta}_1$ is asymptotically normally distributed, because $\ln(r_t) \sim I(1) \Rightarrow \Delta\ln(r_t) \sim I(0)$, while $\Delta\ln(p_t) \sim I(0) \Rightarrow \Delta^2\ln(p_t) \sim I(0)$; this estimator uses only I(0) variables and the standard distribution theory applies.[11] Estimation produces the following results:

$$\Delta \ln(r_t) = \underset{(2.69)}{3.56}\, \Delta^2 \ln(p_t),$$

where $R^2 = 0.018$, $\sigma = 0.190$ and DW $= 2.04$. These results allow correct statistical inference on the estimated coefficients to be drawn, and also the Durbin–Watson statistic is indicative of no remaining autocorrelation in the residuals. However, the standard error of the $\tilde{\beta}_1$ coefficient is 2.69, which gives a t-ratio of 1.33. The null hypothesis $H_0\colon \tilde{\beta}_1 = 0$ cannot be rejected. This result again implies that the Fisher equation does not hold in Croatia. Thus, it may be that the inflation rate enters the long-run money-demand relation as a separate variable along with the interest rate.

IV.3 Money demand estimation

Following Baba *et al.* (1992), the baseline real money demand, or $(m - p)_t$, is specified so as to include real income y_t, the nominal interest rate r_t and the inflation rate Δp_t. Within a multivariate cointegration framework, the order of the estimated VECM needs to be properly specified in terms of the lag-length selection before commencing with the cointegration analysis. Formal tests of the system's reduction validity, progressively reducing the number of lags in the system, reject all reductions beyond VAR(12), making the model a VECM with

[9] However, Sargent's (1972) extension that includes levels of money will yield valid inference, given that money is I(1) and cointegrated with interest rates; hence, the I(0) inflation would enter merely as an additional stationary regressor.

[10] We assume the variables are measured as deviations from the means.

[11] To see that $\tilde{\beta}_1$ is a consistent estimator of $\hat{\beta}_1$, observe that

$$\ln(r_t) - \ln(r_{t-1}) = \hat{\beta}_0 + \hat{\beta}_1 \Delta \ln(p_t) + \hat{u}_t - [\hat{\beta}_0 + \hat{\beta}_1 \Delta \ln(p_{t-1}) + \hat{u}_{t-1}]$$
$$\Rightarrow \Delta \ln(r_t) = \hat{\beta}_1 \Delta^2 \ln(p_t) + \breve{u}_t$$

where $\Delta^2 \ln(p_t) \equiv \Delta\ln(p_t) - \Delta \ln(p_{t-1}) = \ln(p_t) - 2\ln(p_{t-1}) + \ln(p_{t-2})$ and $\breve{u}_t \equiv \hat{u}_t - \hat{u}_{t-1}$. However, the $\hat{\beta}_0$ coefficient from $\ln(r_t) = \hat{\beta}_0 + \hat{\beta}_1 \Delta \ln(p_t) + \hat{u}_t$, i.e., the long-run equilibrium real rate of interest, cannot be estimated.

$\Delta z_t = [\Delta(m-p)_t, \Delta y_t, \Delta^2 p_t, \Delta r_t]$, and using 12 lags. The four-variable system is specified as

$$\begin{pmatrix} \Delta(m-p)_t \\ \Delta y_t \\ \Delta^2 p_t \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} \omega_1 \\ \omega_2 \\ \omega_3 \\ \omega_4 \end{pmatrix} + \sum_{i=1}^{12} \begin{pmatrix} \phi_{11}^{(i)} & \phi_{12}^{(i)} & \phi_{13}^{(i)} & \phi_{14}^{(i)} \\ \phi_{21}^{(i)} & \phi_{22}^{(i)} & \phi_{23}^{(i)} & \phi_{24}^{(i)} \\ \phi_{31}^{(i)} & \phi_{32}^{(i)} & \phi_{33}^{(i)} & \phi_{34}^{(i)} \\ \phi_{41}^{(i)} & \phi_{42}^{(i)} & \phi_{43}^{(i)} & \phi_{44}^{(i)} \end{pmatrix} \begin{pmatrix} \Delta(m-p)_{t-i} \\ \Delta y_{t-i} \\ \Delta^2 p_{t-i} \\ \Delta r_{t-i} \end{pmatrix}$$

$$+ \begin{pmatrix} \psi_{11} & \psi_{12} & \psi_{13} & \psi_{14} \\ \psi_{21} & \psi_{22} & \psi_{23} & \psi_{24} \\ \psi_{31} & \psi_{32} & \psi_{33} & \psi_{34} \\ \psi_{41} & \psi_{42} & \psi_{43} & \psi_{44} \end{pmatrix} \begin{pmatrix} \xi_{11} & \xi_{12} & \xi_{13} & \xi_{14} \\ \xi_{21} & \xi_{22} & \xi_{23} & \xi_{24} \\ \xi_{31} & \xi_{32} & \xi_{33} & \xi_{34} \\ \xi_{41} & \xi_{42} & \xi_{43} & \xi_{44} \end{pmatrix} \begin{pmatrix} (m-p)_{t-1} \\ y_{t-1} \\ \Delta p_{t-1} \\ r_{t-1} \\ t \end{pmatrix}$$

$$+ \begin{pmatrix} \iota_{1t} \\ \iota_{2t} \\ \iota_{3t} \\ \iota_{4t} \end{pmatrix}.$$

Estimation using the Johansen maximum likelihood technique indicates two stationary combinations among (real) money, output, the interest rate and inflation rate variables (Table 5).[12] In particular, the restricted estimation where the rank condition ($r = 2$) and weak exogeneity of inflation were jointly imposed produced an LR $\chi^2_{(2)}$ test of 3.733 ($p = 0.155$). Thus, the joint hypothesis that $r = 2$ and that inflation is weakly exogenous with respect to the long-run parameters cannot be rejected.

The $\boldsymbol{\xi}'$ and $\boldsymbol{\psi}$ are estimated as

$$\boldsymbol{\xi}' = \begin{pmatrix} 1.00 & -2.66 & 17.00 & 0.36 & 0.0079 \\ -0.02 & 1.00 & -3.40 & 0.09 & -0.0002 \\ 0.00 & -0.01 & 1.00 & -0.00 & -0.000 \\ 18.45 & -15.30 & -147.46 & 1.00 & -0.1223 \end{pmatrix}$$

$$\boldsymbol{\psi} = \begin{pmatrix} 0.09 & -0.58 & -4.93 & -0.0023 \\ 0.20 & -0.99 & 0.23 & 0.0025 \\ -0.02 & 0.02 & -1.18 & 0.0005 \\ -2.42 & -3.34 & 15.53 & 0.0027 \end{pmatrix}.$$

[12] See also Cziráky (2002).

TABLE 5

*Johansen cointegration tests: $z = [(m - p)_t, y_t, \Delta p_t, r_t]$**

H_0: $r = p$	λ-*max*	95% *CV*	λ-*trace*	95% *CV*
$p = 0$	66.11	31.5	124.50	63.0
$p \leq 1$	36.03	25.5	58.40	42.4
$p \leq 2$	15.48	19.0	22.37	25.3
$p \leq 3$	6.89	12.3	6.89	12.3

*Eigenvalues: $\lambda_1 = 0.528$, $\lambda_2 = 0.336$, $\lambda_3 = 0.161$, $\lambda_4 = 0.075$.

Imposing the rank restrictions, the estimates of $\boldsymbol{\xi}'$ and $\boldsymbol{\psi}$ are

$$\boldsymbol{\xi}' = \begin{pmatrix} -0.23 & 0.59 & -3.04 & -0.08 & -0.0018 \\ -0.02 & 0.67 & -2.06 & 0.06 & -0.0002 \end{pmatrix}$$

$$\boldsymbol{\psi} = \begin{pmatrix} -0.37 & -0.86 \\ -0.99 & -1.51 \\ - & - \\ 10.71 & -4.97 \end{pmatrix}.$$

The adjustment coefficients for the money-demand relation are large and negative (−0.37 and −0.86), which indicates fast adjustment to the long run. Normalizing the first cointegrating relation to $(m - p)_t$ and the second one to y_t and writing the long-run relationships in equation format, the long-run money demand and income determination equations are

$$(m-p)_t = 2.57y_t - 13.22\Delta p_t - 0.35r_t - 0.01t,$$
$$y_t = 0.03(m-p)_t + 3.07\Delta p_t - 0.09r_t + 0.003t.$$

The latter relation can be interpreted as a small real balance effect on output (see, for example, Ireland, 2005, on this effect) or as indicating a Phillips curve relation.

One-step and breakpoint Chow tests were conducted for the individual equations and for the entire system. Stability of the system is indicated by the fact that the recursive breakpoint Chow tests generally fall below the 95 percent critical value. The one-step Chow tests detect an outlier in March 2000.

IV.3.1 Money demand without the inflation rate

As part of the robustness check of the baseline model, the money demand is also estimated with the assumption that the Fisher equation holds, and hence the inclusion of the inflation rate is not necessary. The estimation of the system without the inflation rate term requires a three-variable VECM instead of the four-variable one for the baseline. Experiments here find three cointegrating vectors with two of the three eigenvalues significant on the basis of both λ-max and λ-trace statistics (Table 6). This suggests that the third vector is apparently non-stationary, or I(1), while the estimates of the cointegrating vectors and their adjustment coefficients are similar in both the models. The money-demand cointegrating vector is $(m_t - p_t) = 2.25y_t - 0.44r_t - 0.01t$.

Additional tests are made for the reduced rank $r = 2$ and (jointly) for the exclusion of the deterministic trend from the cointegrating space. The exclusion of the trend is strongly rejected by the LR test statistic: $\chi^2_{(2)} = 25.36$. A significant problem with the reduced rank model emerges from one-step and breakpoint Chow tests. These tests are failed, which indicates a lack of parameter stability (or constancy) that may be causing instability of the entire system.

IV.3.2 Money demand with exchange rates

As another alternative that checks the robustness of the baseline specification, we consider Petrovic and Mladenovic's (2000) model of money demand which includes exchange rates in lieu of a nominal interest rate or an inflation rate. This approach is based on 'dollarization' or 'fear of floating' arguments (Calvo and Reinhart, 2002), although note that Taylor (2001) is more circumspect about what role exchange rates might play during an inflation targeting regime. To test the exchange rate approach, money demand is re-estimated with the exchange rate (ex_t) replacing the inflation rate; the nominal interest rate is kept in the system. The VECM system is then $\Delta\hat{z}_t = [\Delta(m - p)_t,\ \Delta y_t,\ \Delta ex_t,\ \Delta r_t]$, and the results of the cointegration tests are presented in Table 7. They indicate as many as three cointegrating vectors.

TABLE 6

*Johansen cointegration tests VAR(11): $\mathbf{z} = [(m - p)_t, y_t, r_t]$**

H_0: $r = p$	λ-*max*	*95% CV*	λ-*trace*	*95% CV*
$p = 0$	48.90	25.5	79.80	42.4
$p \leq 1$	26.25	19.0	30.90	25.3
$p \leq 2$	4.65	12.3	4.65	12.3

*Eigenvalues: $\lambda_1 = 0.419$, $\lambda_2 = 0.253$, $\lambda_3 = 0.050$.

TABLE 7
*Johansen cointegration tests: $\hat{z} = [(m-p)_t, y_t, ex_t, r_t]$**

H_0: $r = p$	λ-*max*	*95% CV*	λ-*trace*	*95% CV*
$p = 0$	60.62	31.5	149.00	63.0
$p \leq 1$	49.96	25.5	88.37	42.4
$p \leq 2$	26.32	19.0	38.41	25.3
$p \leq 3$	12.09	12.3	12.09	12.3

*Eigenvalues: $\lambda_1 = 0.498$, $\lambda_2 = 0.430$, $\lambda_3 = 0.256$, $\lambda_4 = 0.127$.

The unrestricted estimates of the $\boldsymbol{\xi}'$ and $\hat{\boldsymbol{\psi}}$ matrices are similar to the baseline model. Restricting the cointegrating rank to $r = 2$ and imposing weak exogeneity of the exchange rate gives the following estimates:

$$\hat{\boldsymbol{\xi}}' = \begin{pmatrix} -0.48 & 1.69 & 0.10 & -0.17 & -0.005 \\ -0.16 & 1.62 & -0.30 & 0.20 & -0.004 \end{pmatrix}$$

$$\hat{\boldsymbol{\psi}} = \begin{pmatrix} -0.03 & -0.02 \\ -0.30 & -0.72 \\ - & - \\ 3.10 & -2.65 \end{pmatrix}.$$

The LR test for the imposed restrictions has a $\chi^2_{(2)}$ of 2.226 ($p = 0.329$), which does not reject the joint restriction that $r = 2$ and that the exchange rate is weakly exogenous for the long-run parameters. A notable difference, however, is in the near-zero values for the adjustment parameters in the money-demand equation (−0.03 and −0.02). Including the exchange rate in place of the inflation rate causes the model to lose completely the fast short-run adjustment property of the baseline model. The adjustment would take place almost never, making the exchange rate model unable to explain a stable money demand in the face of shocks.

V. CONCLUSION

The paper presents a rigorous model of money demand for an EU accession country, Croatia, during its transition years. First, it examines whether the classical Fisher equation of interest rates holds, whereby the nominal interest rate should move together with the inflation rate. Transition/EU-accession countries such as Croatia are perhaps especially likely to be undergoing changes in inflation rate policy that produce unexpected inflation rates. This can lead to a failure of the nominal interest rate and the inflation rate to move together. Finding no evidence in support of the Fisher interest equation for Croatia, using a battery of tests, the paper then specifies the baseline model as a classical money-demand function extended to include the inflation rate. With vector error correction methods, a cointegrated money-demand function results with both parameter stability and timely dynamic re-equilibration to shocks.

For robustness, the baseline model specification is compared to likely alternative specifications. First examined is the baseline without the inflation rate, this being the standard, classic, money-demand function. This alternative exhibits parameter instability. Second, the exchange rate is substituted into the baseline model in place of the inflation rate. This reflects a theme of the transition money-demand literature that the exchange rate acts as the inflation rate in the money-demand function, because the inflation rate is fully 'passed through' to exchange rate changes. This specification shows long-run cointegration but no timely dynamic adjustment to shocks. The lack of short-run adjustment makes it an inferior alternative. The robustness of the baseline model relative to the main alternatives allows for some confidence in the results.

Interpretation requires caution because of the data limitations that characterize all transition country studies. Starting the data series for Croatia only in 1994 avoids a hyperinflation that peeked at around a 1500 percent annual rate in 1993; after this, a new currency was introduced. Given the data qualification, the results can be interpreted first as showing that a stable money demand exists despite a less than calm period economically and politically.

Second, the analysis suggests that a policy that causes gradual changes in the inflation rate is unlikely to disrupt the baseline money-demand function because it includes the inflation rate as a variable. This means that a policy of maintaining a low inflation rate, or even gradually reducing the inflation rate if it were at a higher level as in Hungary, is not likely to induce an apparent instability in the estimated money-demand function. In turn, the inflation rate should be able to be more easily forecasted using variables that enter the money-demand function. Then the forecasts can be used by the central bank to continue to act to stabilize the inflation rate, a type of self-reinforcing interaction of policy with the behaviour of consumers.[13]

In contrast, a policy for example that targets the exchange rate without regard to the inflation rate could induce unexpected jumps in the inflation rate that cause apparent 'shifts' in the money-demand function. This can lead to the belief that money demand is unstable, and justify further discretion from the central bank to offset the apparently unstable money-demand function. This circle of interaction between

[13] In discussing inflation forecasting, Balfoussia and Wickens (2005) note that 'Although there is no necessary reason for a good forecasting model to have theoretical underpinnings, theory may still be able to help in the choice of the model to use' (p. 1).

policy and the consumer is less appealing in that the ultimate policy would probably be less efficacious and could lead to an 'inflation bias'. This is not to argue that exchange rate targeting is necessarily worse than inflation rate targeting. It does suggest that the use of exchange rate instruments in Croatia may *de facto* be part of a policy of inflation rate targeting.

Policy–consumer interaction is an important factor in the ultimate efficaciousness of policy, as emphasized by Lucas's (1976) 'critique'. The nature of such interaction in general is not regime or consumer-behaviour dependent, given the usual assumption of rationality of the agents. The specifics of the policy 'function' that incorporates the consumer behavioural reactions will certainly change with the particular policy employed. Some policies will be less wasteful of societal resources than others. Arguably, a stable money-demand function combined with inflation rate goals results in a rather efficacious interaction. And it is of some interest to see such a stable money-demand function arising in a dynamic economy like Croatia that has an explicit price stability goal set out in its central bank act.

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