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Testing the Lucas Critique for the Turkish Money Demand Function

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Abstract. This paper aims to test the prevalence of the Lucas critique by use of an applied modelling approach. The Turkish narrow money demand is chosen for investigation purposes and an extensive statistical-based econometric application has been carried out to observe whether the model in question has been exposed to the content of such a critique. The results confirm the theory to explain the behavioral foundations of aggregate monetary economics approaches and reveal that no evidence can be found in favor of the non-rejection of the Lucas critique that leads us to infer that the modelling attempt can be considered by the researcher a feedback model which is able to encompass a whole class of expectation models.

Key Words: Money Demand; Lucas Critique; Turkish Economy;

JEL Codes: C32; E41; E61;

Özet. **Para talebi eşitlikleri için Lucas eleştirisinin sınanması: Türkiye ekonomisine uygulama.** Bu çalışma Lucas eleştirisinin geçerliliğini uygulamalı bir modelleme yaklaşımı kullanarak sınamayı amaçlamaktadır. Türk dar para talebi araştırma amacı ile seçilmekte ve geniş kapsamlı istatistiksel-temelli ekonometrik bir uygulama sorgulanın modelin bu tür bir eleştiriin içeriğine maruz kalmış olup olmadığını gözlemlemek için gerçekleştirılmıştır. Sonuçlar kuramı derneşik parasal iktisat yaklaşımının davranışsal bulgularını açıklamak için doğrulamakta ve Lucas eleştirisinin reddedilememesi yönünde bir bulguya ortaya koyamamaktadır. Bu sonuç ise bizi modelleme yaklaşımının araştırmacılar tarafından geniş kapsamlı olarak beklenen modellerini kapsayabilecek şekilde geribeslemeli bir model olarak dikkate alınabileceği çıkışmasına götürmektedir.

Anahtar Kelimeler: Para Talebi; Lucas Eleştiri; Türkiye Ekonomisi;

JEL Sınıflaması: C32; E41; E61;

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1. Introduction

Today's contemporaneous economic debates witness that one of the most important issues of interest for policy makers is to provide consistency of forecasts resulted from model evaluation process with the decision making of economic agents. Related to the well-known Lucas' critique (Lucas, 1981), since the optimal decision rules on which the structure of econometric models are based have been varied with changes in the structure of series that represent the behaviour of economic agents, the structure of econometric models used for estimation purposes will also have been altered by the systematic changes in the policy choices. That the critique holds gives rise to that comparisons of the effects of alternative policy rules using macroeconomic models will be invalid regardless of the performance of these models over the sample period or in short term forecasting.

We can easily observe that in the real world of economies, attempts to make simulations derived from empirical models coincide in many times with the application of alternative policy regimes. For applied studies, this means a structural shift in the parameters of estimated equations resulted from regime changes, and in this case, baseline regime conclusions should not be used to evaluate the effect of the control policy. That is, the Lucas critique applies (Favero, 2001). Hence, if conditional economic models have been found dependent upon specific policy actions and institutional structures of the economy though they have been estimated by using most recent or popular econometric estimation techniques, substantial changes in policies or the institutional structure may lead researchers to unwarranted estimation results and nullify the best econometric models even when the estimates seem to have desired statistical prerequisites (Stanley, 2000). In these cases subject to the regime changes and parameter instabilities, empirical studies used for estimation purposes will probably be undermined in a way leading to the invalidated policy proposals.

The concept of money demand can be considered a good candidate to examine these kind of criticisms in the economics literature. Briefly to say, money demand deals with what motives determine the economic agents' holding of monetary balances. Given the *ex-ante* designed purposes, knowledges extracted from money demand functional forms can provide insights as to the future course and success of policy implementations and help decision makers to decide about which monetary policies are better to implement under the current economic conditions. Therefore, these inferences are able to best serve to appreciate the analyses of policy choices. If policy makers lack of observing the true data structure of money balances, disorderly velocity shocks which lead to persistent deviations of growth rates of

monetary aggregates from estimated values will likely to dominate the markets in the economy (Kontolemis, 2002). Considering these cursory appreciations for modelling purposes, we can state that conclusions derived from money demand equations are capable of providing crucial knowledge of how economic agents determine their behaviors of monetary holdings.

In this sense, such researches can yield estimation results informative for the constancy of money demand models which are also invariant to regime changes. Following Cheong (2003), the main difference between constancy and invariance of the applied modelling approaches is that while the former is related to the time independence of parameters, the latter would mean constancy across policy interventions both of which are among our main purposes of testing in the paper. We hope that these methodological choices in constructing and testing economic model in question will be highly useful for researchers and policy makers in search for the appropriateness of monetary stabilization policies.

This study aims to construct a conventional aggregate money demand model by employing recent developments in the time series methodologies, and for this purpose, uses data taken from the Turkish economy. Of special emphasis, in so doing, however, has been given to the stability of empirical regularities found in the data and we try to elaborately investigate whether the criticisms directed to the applied studies as summarized above can be attributed to the data realizations examined in this paper. As is briefly expressed in the following sections, there exist many papers dealing with empirical money demand models constructed upon the Turkish economy. Our paper is aimed to contribute to these studies both by re-examining the money demand function for the recent post-1998 period up to the year 2010 and by further testing the concept of non-linearity in such a function. To the best of our knowledge, there exists no paper in the Turkish economics literature applying an estimation methodology outlined in this study. We must state that in addition to the estimation of a theoretically consistent money demand model from which both conditional and marginal models are derived to test for the issues of constancy and invariance, one of the main contributions to this strand of literature is to enable the reader to observe the testing procedure of a non-linear vector error correction model also permitting smooth adjustment toward equilibrium conditions for the Turkish economy.

Thus, we are motivated by the testing procedure of whether possible instabilities leading to a critique stemmed from, e.g. Lucas (1981), is of a special concern for us, and if so, as is plainly documented by Özmen (1996), “not only dynamic misspecifications but also an invalid conditioning and a change in the relevant variable space due to a policy regime change

and/or financial innovation should be taken as potentially complementary explanations of a money demand instability". Given the importance of a stable money demand relationship, many studies in recent years have been conducted on various country cases by researchers such as Sriram (1999), Kontolemis (2002), Ramachandran (2004) and Dreger et al. (2006). On the other side, Yavan (1993), Metin (1994; 1995), Üçdoğruk (1996) in this journal , Civcir (2003), Bahmani-Oskooee and Karacal (2006), Çatik (2007) in this journal and some papers by the CBRT researchers such as Mutluer and Barlas (2002), Akıncı (2003) and Altınkemer (2004) try to test the demand for money relationship for the Turkish economy.

In light of this introduction, the next section discusses some recent methodological developments to test non-linearities in time series analyses. Therefore, we extend our scope to smooth transition regression modelling to test for the existence of a non-linear equilibrium correction model for the research area in this paper. Benefited from these explanations, an application to the Turkish narrow money balances is carried out in section 3. The last section summarizes results to conclude the paper.

2. Non-linearity and Smooth Transition Regressions (STR): Conceptual Fundamentals

Since the multivariate co-integration and vector error correction methodologies as well as the related exogeneity issues are well-known in the economics literature, we omit a methodological discussion upon these subjects. But the interest readers are suggested to glance at the seminal papers of Engle et al. (1983), Engle and Granger (1987), Johansen (1988), and Johansen and Juselius (1990) or can apply to a well-mixture of these papers in Favero (2001) and Harris and Sollis (2003).

As an empirical contribution to these papers, let us briefly summarize STR framework so that we are able to test the non-linearity of error correction model derived from the co-integration relationship. Through the developments within the last two decades, we tend to follow the approaches mainly revealed by the studies of Granger and Teräsvirta (1993), and Teräsvirta (1994, 1998). The common form of STR modelling can be summarized as in Eq. 1:

$$y_t = \phi' z_t + \theta' z_t G(\gamma, c; s_t) + \varepsilon_t \quad (1)$$

where $z_t = (w_t', x_t')'$ is a $((m+1) \times 1)$ vector of explanatory variables with $w_t' = (1, y_{t-1}, \dots, y_{t-p})'$ and $x_t' = (x_{1t}, \dots, x_{kt})'$. ε_t is assumed to be subject to i.i.d. $(0, \sigma^2)$ process. ϕ represents the

parameter vector for the linear part of the model, while θ is used for the non-linear parameter vector. $G(\cdot)$ expresses applying to a non-decreasing and continuous transition function usually bounded between 0 and 1, and depends on the transition variable s_t , the slope parameter γ and the vector of location parameters c . The transition variable s_t can be part of z_t , as one of the explanatory variables, or of a time trend. The slope parameter $\gamma > 0$ measures the speed of transition between 0 and 1, and the location parameter c brings out where the transition takes place. Observe that the model enables the researcher to begin the analysis with a linear model nested in the formulation, and permits for a further investigation to proceed with a non-linear model specification. The most common forms of transition function of a K^{th} order STR model using a logistic function with $\gamma > 0$ and $c_1 \leq \dots \leq c_k$ can be written as follows:

$$\text{LSTR1 Model where } K=1: G_1(\gamma, c; s_t) = \frac{1}{1 + e^{-\gamma(s_t - c)}} \quad (2)$$

$$\text{LSTR2 Model where } K=2: G_1(\gamma, c_1, c_2; s_t) = \frac{1}{1 + e^{-\gamma(s_t - c_1)(s_t - c_2)}} \quad (3)$$

For $s_t > c_t$, the transition function will approach 1 as $\gamma \rightarrow \infty$. In the above equations, we test non-linearity against alternative LSTR(k) models by searching for whether $\gamma=0$. The model is then identified by applying to a Taylor series approximation of the transition function that was developed by Luukkonen et al. (1988). Teräsvirta (1994) derives LM-type tests of linearity against LSTR model. A transformed auxiliary regression is used if s_t is an element of z_t :

$$y_t = \beta_0 z_t + \sum_{j=1}^3 \beta_j z_t s_t^j + \varepsilon_t^* \quad (4)$$

where $z_t = (1, z_t')$. The null hypothesis of linearity would be:

$$H_0: \beta_1 = \beta_2 = \beta_3 = 0 \quad (5)$$

The test is carried out for each of the chosen variables and the variable with the strongest rejection against the H_0 hypothesis is suggested to select as a decision rule. If non-linearity cannot be rejected, the model proceeds to choosing the type of LSTR model based on a sequence of sub-hypotheses using same auxiliary regression Eq. 4:

$$\begin{aligned}
H_{04} : \beta_3 &= 0 \\
H_{03} : \beta_2 &= 0 | \beta_3 = 0 \\
H_{02} : \beta_1 &= 0 | \beta_2 = \beta_3 = 0
\end{aligned} \tag{6}$$

An F -test is used for this purpose, and the test statistics for the null hypothesis H_{04} , H_{03} , H_{02} are estimated as F_4 , F_3 and F_2 , respectively. In line with Teräsvirta and Eliasson (2001), our decision rule is based on the inference that if the p -value of the test of H_{03} is rejected more strongly than H_{02} and H_{04} , we tend to decide the use of an LSTR(2) model, otherwise an LSTR(1) model will be chosen. In case the test sequence cannot yield unambiguous results, the researcher can scrutinize the relevant information criteria or the residual sum of squares.

Note that in the LSTR formulation, the parameters will change monotonically with the transition function with asymmetric adjustment toward equilibrium, and as $\gamma \rightarrow \infty$, the model becomes a two-regime switching model. Whereas, in exponential STR (ESTR) models, the adjustment would be symmetric around long run equilibrium as is in Eq. 7:

$$G(\gamma, c; s_t) = 1 - e^{-\gamma(s_t - c)^2}, \gamma > 0 \tag{7}$$

3. Empirical Findings

3.1. Preliminary Data Issues: Definitions and Time Series Characteristics

Following the theoretical bases summarized in the former sections, a money demand model using the Turkish data is constructed. For this purpose, the sample period with quarterly observations lies between 1998Q1 and 2010Q4. The motivation of the starting point is the use of 1998: 100 based new income series to extract the price data derived from gross domestic product (GDP) deflator (p). The monetary variable (m) is the narrow money balances, which is the sum of currency in circulation and demand deposits in the banking system. Under the assumption of no money illusion, the demand for money is assumed demand for real money balances. We can state that narrowly defined money balances reflect rather the transaction necessities in the economy, whereas broadly defined balances would be better off for the portfolio approaches. As a next step, the alternative costs discouraging people to hold these balances must be specified. In a standard money demand relationship, these may be a selected interest rate, e.g. on securities and bonds, returns of equities, changes in the exchange rate representing currency substitution especially for a developing economy and also the inflation

rate representing the increase of prices of intangible assets. Given these choices, the scale income variable (y) for the maximum amount of money balances to be held in hand is the real GDP data at constant 1998 prices. The interest rate variable (i) is an average interest rate on securities in Treasury actions to represent returns on financial assets. For the effect of the course of equity prices on money demand equation, the Istanbul Stock Exchange (ISE) National-100 index (eq) is considered. Further, putting a variable representing currency substitution or dollarization phenomenon in a money demand functional form may be important especially in an inflationary environment. Such a variable will reflect the extent to which policy authority loses domestic monetary control when the demand for foreign exchange increases in the economy. For this purpose, the Turkish lira / US dollar exchange rate (e) is included into the analysis.

We must state that within the estimation process, we have observed that adding a price variable converted into inflationary developments as an alternative cost besides the interest rate has been resulted in a contradictory inverse sign due possibly to a multicollinearity problem between these aggregates, but putting one of them as a single variable has given us a theory-consistent functional relationship. Therefore, we decided to proceed with interest rate and dropped the inflation variable from the money demand equation. Finally, the own return of money balances is assumed to be zero for the narrow money demand. All the data except the interest rates are used in their natural logarithms and indicate seasonally unadjusted values. For the variables m , y , eq and e , the source is the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT), while data for interest rate variable i are compiled from the electronic data delivery system of the Turkish Republic Prime Ministry State Planning Organization. The underlying model can be expressed as follows:

$$(m - p)_t = c_t + \beta_1 y_t + \beta_2 i_t + \beta_3 eq_t + \beta_4 e_t + \varepsilon_t \quad (8)$$

where ε_t is *a priori* assumed to be a white noise error term.

Having defined the data, we will conduct an empirical analysis of aggregate money demand function. For this purpose, we will also take advantage of the knowledge that various exogeneity concepts provide us. As Hendry and Ericsson (1991a) well-specify, this distinction would be correspond to different notions of being determined outside the model according to the researcher's purpose to investigate, and in no case would it be legitimated to arbitrary relating the variables to exogeneity properties simply by not modelling them.

Using non-stationary variables in estimation process will likely to produce the *so-called* spurious regression problems in conventional ordinary least squares (OLS) estimation techniques that strongly lead the variables to diverging from long run means with biased standard errors and result in unreliable correlations and unbounded variance processes. Also, the existence of structural breaks in time series would make it difficult to discern definitive properties in a variable form. Therefore, we first conduct some unit root tests suggested by Dickey and Fuller (1981) assuming no structural break, Zivot and Andrews (1992) assuming one endogenous break and Lumsdaine and Papell (1997) that allow for two endogenous breaks. To save space, we have omitted methodological discussion upon these tests. In Table 1, Both the ADF, ZA and LP tests, for which the latter two tests can be referred to as C and CC models, are not able to reject the unit root null hypothesis in the level form variables:

Table 1. Univariate unit root tests

Variable	Augmented Dickey-Fuller (ADF) Test	Zivot-Andrews (ZA) Test	Lumsdaine-Papell (LP) Test
$(m - p)$	-3.02 ^{c&t} (lag 4)	-2.83 (lag 2)	-6.05 (lag 6)
$\Delta(m - p)$	-7.16 ^c (lag 1) [*]	TB: 03Q3	TB1: 04Q4 & TB2: 10Q2
(y)	-0.30 ^c (lag 5)	-3.28 (lag 2)	-5.20 (lag 8)
$\Delta(y)$	-3.10 ^c (lag 4) ^{**}	TB: 08Q4	TB1: 05Q4 & TB2: 09Q2
(i)	-2.56 ^{c&t} (lag 1)	-3.68 (lag 1)	-5.11 (lag 1)
$\Delta(i)$	-4.67 ^c (lag 3) [*]	TB: 01Q1	TB1: 03Q1 & TB2: 06Q4
(eq)	-2.96 ^{c&t} (lag 1)	-3.80 (lag 1)	-6.29 (lag 1)
$\Delta(eq)$	-5.30 ^c (lag 1) [*]	TB: 00Q4	TB1: 00Q4 & TB2: 04Q4
(e)	-2.11 ^{c&t} (lag 1)	-4.684 (lag1)	-5.63 (lag 7)
$\Delta(e)$	-4.53 ^c (lag 1) [*]	TB: 01Q2	TB1: 03Q1 & TB2: 06Q2

Notes: In the ADF test * and ** indicate the rejection of the null hypothesis for the 1% and 5% levels, respectively. ^c and ^{c,t} represent the allowance for only constant and constant&trend terms as deterministic components, respectively. Δ is the difference operator and $\Delta y_t \equiv y_t - y_{t-1} \forall y$; k is the number of lags determined for each possible breakpoint by minimizing the Schwarz Bayesian information criteria and ε_t is assumed to be i.i.d. error term. The critical vales for the ZA test are -5.57 (1%) and -5.08 (5%). The critical value for the LP test equal -7.19 (1%) and -6.75 (5%), which are taken from the paper of Ben-David et al. (2003).

3.2. Rank Analyses and Evidence for VAR-Based Multivariate Co-integration

The previous analysis indicates that all variables seem to have I(1) process, thus, we now try to estimate an unrestricted VAR model. For the dynamic structure of the model, we consider Schwarz information criterion that suggests the use of 1 lag order. To account for seasonality

in the model estimation process, a set of centered seasonal dummies which sum to zero over a year are used. Further, an impulse dummy variable (*dummy01*) which takes on values of unity from 2001Q1 to 2001Q4 is attributed to the macroeconomic crisis conditions experienced by the Turkish economy. These dummies in the model enter the system in an unrestricted way, because, we do not expect them to have long-run effects endogenous to the specification of the variable vector. An intercept and trend factor are restricted into the long-run variable space as exogenous variable for the deterministic part of the model, but we do not assume a quadratic deterministic trend lying in both the co-integrating model and the dynamic VEC model since allowing for linear trends in the short-run VEC model possibly leads the researcher to be obliged to rationalize why an implausible ever-increasing or decreasing rate of change dominates the data in an economic sense. We will see below that such a choice of restricting trend factor into the system yields significant estimation results so that attempts to drop this deterministic term from the model will be a lack of researcher to catch the true data generation process derived from our applied modelling. The long-run co-integrating relationships between the variables are estimated by using two likelihood test statistics known as maximum eigenvalue for the null hypothesis of r versus the alternative of $r+1$ co-integrating relationships and trace for the null hypothesis of r against the alternative of n co-integrating relations, for $r = 0, 1, \dots, n-1$ where n is the number of endogenous variables.

Based on the critical values taken from Osterwald and Lenum (1992) and on newer p -values from the study of MacKinnon et al. (1999), in Table 2 in Appendix, we first try to determine the rank order of the model for the existence of possible statistically significant co-integrating vectors. In the upper part of the table can easily be observed that against the null hypothesis of no co-integration in the multivariate model we are unable to reject that there exists a unique co-integrating relationship lying in the long-run variable space estimated by both maximum likelihood statistics, thus, some combinations of the I(1) variables of our interest tend to yield a stationary relationship. The unrestricted co-integrating coefficients indicate that the first row with the largest eigenvalue statistic seems to satisfy the *a priori* theory-based expectations for a conventional money demand model, since all the variables carry quite plausible coefficients to which economic interpretations in Eq. (8) can be imposed and are found with statistically significant normalized signs with regard to the real money balances:

$$(m - p)_t = -18.40 + 2.3672 y_t - 1.7680 i_t - 0.6091 eq_t - 0.2175 e_t + 0.0132 trend + \varepsilon_t \quad (9)$$

We estimate that the real income elasticity of money balances under the money market equilibrium conditions is 2.37, which is obviously larger than a unitary coefficient. Following Sarno (1999), on this point, we can suppose that the disequilibria between real income and real money balances will affect the current demand conditions through the inverse monetary velocity measure ($m-p-y$). If real income elasticity equals unity in a long-term stationary relationship, which can also be tested by employing homogeneity restrictions, acceptance of this assumption will give support to the quantity theoretical approaches that assume a strong proportional relationship between real income and real balances to provide a stationary income velocity of monetary aggregates. If real income elasticity takes values between one-half and unity, such a finding will be consistent with the economies of scale argument put forward in the context of the inventory-theoretic transactions models. On this issue, see e.g. Ozmen (1998) estimating a currency seigniorage model for the Turkish economy. On the other hand, if real income elasticity is significantly above unity as is found in our study, which will indicate an increasing ongoing monetization process in the economy, demand for real money balances can be considered like a demand for luxury goods, which will be resulted in declining monetary velocity. Such a finding should not be surprising for the Turkish economy, since inside the sample period of 1998-2010 an obvious downward tendency in domestic inflation dominates the economy. Let us note here that in a consistent way to the above findings, the unit real income homogeneity restriction cannot be accepted for the 5% significance level using the estimation results of $\chi^2(1)=3.0185$ against the table-value $\chi^2(1)=3.8415$.

Besides the real income scale variable, the most significant alternative cost variable is found as the semi-elasticity of the Treasury interest rate with a normalized coefficient -1.7680 followed by the coefficients of equity price and exchange rate variables which take the estimation values -0.6091 and -0.2175, respectively. These results confirm the empirical success of the theory to explain the behavioral foundations of aggregate data-supported monetary economics approaches. We find a significant positive trend which implies that agents' demand for real money balances has an increasing tendency in the sample period.

When we look at the unrestricted adjustment coefficients, we can observe that real money balances take a value highly different from zero. Indeed, in a statistically significant way, nearly 7.3% of the adjustment in the money demand disequilibrium conditions to the long-run equilibrium is realized within one period. Among the other variables, the adjustment coefficients of real income and Treasury interest rate, of the magnitudes 0.0679 and -0.2857

respectively, are statistically significant. For the adjustment coefficients of equity prices and exchange rate, we are unable to reject the null hypothesis of being weakly exogenous, thus, these variables cannot be warranted in a dynamic VEC equation to include the relevant error correction term constructed upon them. Such a case means explicitly that the main factors leading to the weakly exogenous variables are determined out of the money demand variable space. Note also that the joint weak exogeneity restriction for the variables m , y and i cannot be accepted by using a $\chi^2(3)=19.373$ (prob. 0.0002) statistic, but the joint weak exogeneity of the variables eq and e are valid by using a $\chi^2(2)=0.1567$ (prob. 0.9247) statistic.

3.3. Sensitivity to Endogenous Break in the Co-integration Space: Gregory-Hansen Evidence

In the former section, we explore that it is possible to construct a theory consistent co-integrating relationship of the same order integrated variables. However, in a similar way to the unit root analysis, there may be a question of structural shifts in the estimated relationship. For this purpose, we follow the methodology proposed by Gregory and Hansen (henceforth GH) (1996a, 1996b) which enable the researcher to test an endogenously determined break point chosen by the data structure of the model. GH test is an extension of the Engle and Granger (1987) co-integration methodology and proposes a residual based approach to statistically examine the presence of one unknown shift for the null hypothesis of no co-integration against the alternative of co-integration with a break. We will consider three alternative models which are Model C with a level shift, Models C/T with a level shift with trend, and Model C/S with a regime shift that allows the slope vector to shift:

$$\begin{aligned} \text{Model C: Level shift} & \quad y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha X_t + \varepsilon_t \\ \text{Model C/T: Level shift with trend} & \quad y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha X_t + \varepsilon_t \quad t = 1, \dots, n \\ \text{Model C/S: Regime shift} & \quad y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1 X_t + \alpha_2 X_t \varphi_{t\tau} + \varepsilon_t \end{aligned} \quad (10)$$

where:

$$\varphi_{t\tau} = \begin{cases} 0 & \text{if } t \leq [n\tau] \\ 1 & \text{if } t > [n\tau] \end{cases} \quad (11)$$

The unknown parameter $\tau \in (0,1)$ represents the relative timing of the change point, and equals T_B/n where T_B is the break point. $[]$ denotes the integer part. μ_1 is the intercept before the

shift, μ_2 the change in the intercept at the time of the shift, β the coefficient of the time trend, α_1 the co-integrating slope coefficient before the regime shift and α_2 the change in the slope coefficient. Following GH (1996a), the test statistics in Eq. 12 is computed for each break point ($\tau \in T$) in the interval ([0.15n], [0.85n]) recursively, and the smallest value is chosen:

$$\begin{aligned} Z_{\alpha}^* &= \inf_{\tau \in T} Z_{\alpha}(\tau) \\ Z_t^* &= \inf_{\tau \in T} Z_t(\tau) \\ ADF^* &= \inf_{\tau \in T} ADF(\tau) \end{aligned} \quad (12)$$

Above, Z_{α}^* and Z_t^* give the minimum values of the relevant Phillips test statistics, and ADF^* is the minimum value of the ADF test. The results are given in Table 3:

Table 3. GH Test for Structural Shift in the Co-integrating Relationship

Test =>	ADF^*	TB	cv	Z_t^*	TB	cv	Z_{α}^*	TB	cv
<u>Model</u>									
C	-4.28 (lag 4)	07Q2	-5.56	-4.17	08Q2	-5.56	-27.93	08Q2	-59.40
C/T	-4.39 (lag 2)	04Q2	-5.83	-4.00	06Q4	-5.83	-25.87	06Q4	-65.44
C/S	-4.44 (lag 1)	04Q1	-6.41	-4.38	08Q4	-6.41	-29.06	08Q4	-78.52

Notes:

¹ For the autoregressive lag structure of the ADF model, the Bayesian information criterion (BIC) is used.

² 5% critical values (cv) assume four regressor case ($m = 4$), and TB indicates the estimated break date.

The results reveal that we cannot reject the null hypothesis of no co-integration relationship at the 5% significance level, and there exists no endogenous structural break subject to the long-run money demand variable space. Had there not been found evidence of stability in the co-integration analysis, we would have been obliged to estimate the long-run money demand equation for each sub-period considering pre- and post-break sample dates inclusive of related dummy variables. Fortunately, on no account do we have to interest in such an instability issue. We will examine the single equation conditional error correction model.

3.4. The General-to-Specific Single Equation Conditional Error Correction Modelling

We now jump to the single equation unrestricted error correction model for the real money balances using ordinary least squares (OLS) method with a general-to-specific procedure. This model is based on the autoregressive distributed lag structure of each variable in addition

to the one-period lagged knowledge of the co-integrating relationship, as the error correction term, and centered seasonal dummies:

$$D(m-p)_t = c_t + \zeta_1[(m-p)_{t-1} + y_{t-1} + i_{t-1} + eq_{t-1} + e_{t-1})] + \sum_{i=1} \zeta_2 D(m-p)_{t-i} + \sum_{i=1} \zeta_3 D(y)_{t-i} + \sum_{i=1} \zeta_4 D(i)_{t-i} + \sum_{i=1} \zeta_5 D(eq)_{t-i} + \sum_{i=1} \zeta_6 D(e)_{t-i} + \zeta_7 D_Q2 + \zeta_8 D_Q3 + \zeta_9 D_Q4 + \varepsilon_t \quad (13)$$

where the long term co-integrating relationship is indicated in [], and ‘ D ’ represents the first difference operator. For the dynamic lag structure, we follow the long-run model with the use of lag order 1. ε_t is again assumed to be a white noise disturbing term. To eliminate the over-parametrization, we present the reduced form model below for which statistically insignificant variables are sequentially dropped by applying to redundant variables F -test, and parsimonious error correction model only with the econometrically meaningful variable results are obtained:

Redundant Variables $F(4,38)$ -statistic=0.2006 Prob. 0.9364

White Heteroskedasticity-Consistent SEs and Covariance in()

$$\begin{aligned} D(m-p)_t &= 0.0230 - 0.0791 EC_{t-1} - 0.1010 D(i)_{t-1} + 0.0741 D(eq)_{t-1} + 0.0963 D_Q2 + 0.1446 D_Q3 + 0.0715 D_Q4 \\ &\quad (0.0052) (0.0341) \quad (0.0439) \quad (0.0272) \quad (0.0128) \quad (0.0216) \quad (0.0155) \\ L_i \text{ statistics} & [0.0936] [0.0569] \quad [0.4163] \quad [0.1882] \end{aligned}$$

$Adj.R^2 = 0.6339$, Std. Error of Regression = 0.0341, $SSR = 0.0489$, F -stat.=14.8529 (Prob. 0.00), D-W stat.=1.9859,
 B-G AR(1) F -stat.=0.0355 (Prob. 0.8514), B-G AR(4) F -stat.=0.0879 (Prob. 0.9857), J-B=0.1649 (Prob. 0.9208),
 White F -stat.=1.3413 (Prob. 0.2338), ARCH(1) F -stat.=2.5566 (Prob. 0.1167), ARCH(4) F -stat.=1.3127 (Prob. 0.2818),
 Ramsey RESET F -stat.=1.9254 (Prob. 0.1728)

Above, L_i is the Hansen (1992) individual stability test statistics, while L_c is the joint stability test statistic of the coefficients and error variance. Under the null hypothesis of parameter stability, 5% asymptotic critical value for the L_i test with 1 degree of freedom (d.o.f.) is 0.47, and 5% asymptotic critical value for the joint test with $m+1=8$ d.o.f. is 2.11 where m represents the regression parameters. Consider that the dummies are dropped from the equation when the breakpoint tests are carried out.

Further, we have applied to the Quandt and Andrew breakpoint tests (Andrews, 1993) which give a set of statistics for one or more unknown structural breakpoints in the sample period of the error correction equation and estimate a Chow-type break point test which is performed at every observation under the null hypothesis of no breakpoints. For this purpose,

15% trimming is chosen where the first and last 7.5% of the observations are excluded for estimation purposes. The results are given in Table 4 below:

Table 4. Quandt-Andrews Breakpoint Tests

H_0 : No breakpoints within trimmed data

Equation Sample: 1998Q4 2010Q4

Test Sample: 2000Q1 2008Q1

Number of breaks compared: 33

Max. LR F -stat.(2003Q2)=2.6631 (Prob. 0.9999), Max. Wald F -stat.(2005Q1)=3.5229 (Prob. 0.9943),

Exp LR F -stat.=0.6240 (Prob. 0.9999), Exp Wald F -stat.=1.0347 (Prob. 0.9516),

Ave LR F -stat.=1.1172 (Prob. 0.9999), Ave Wald F -stat.=1.7948 (Prob. 0.9468)

Hansen's Instability Tests

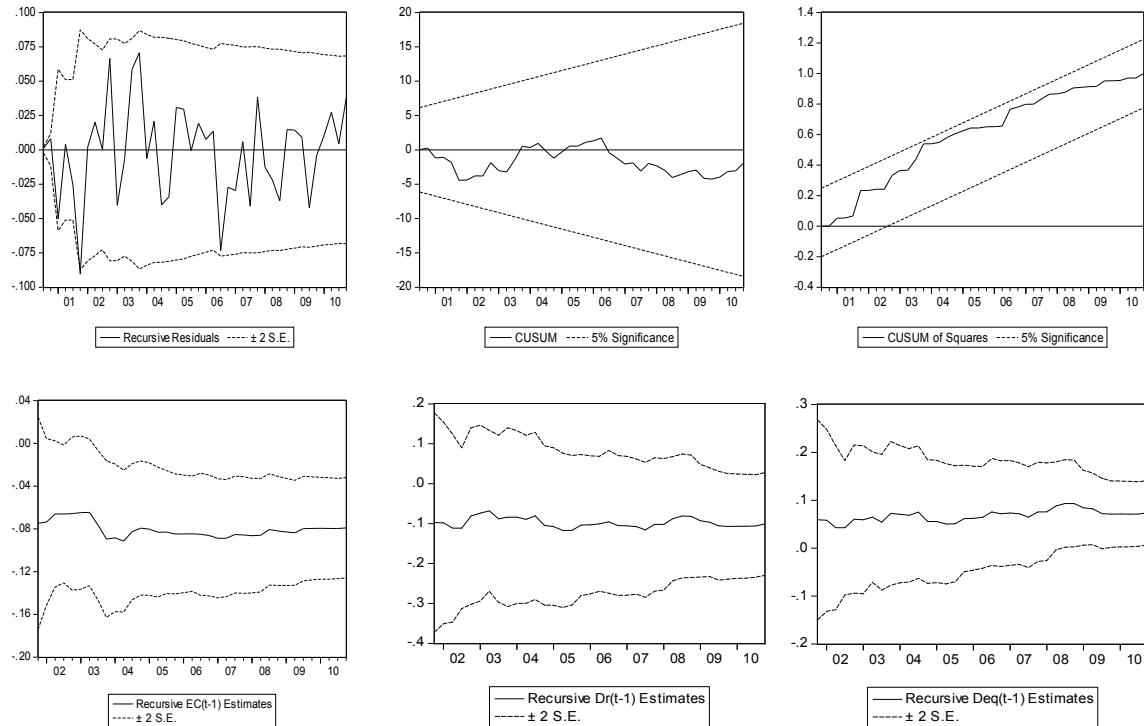
Variance: 0.2920 Joint Statistic L_c : 1.4843

In a supporting way the co-integration results in Table 2, we find that nearly 7.9% of the adjustment in money demand disequilibrium conditions to long-run static equilibrium is realized within one period. We obtain significant coefficient estimates for one-period lagged values of the differenced interest rate and equity price variables. The model is highly well-behaved as for the diagnostics and whitens satisfactorily the residual structure of the parsimonious error correction equation. No unknown breakpoint can be seen in the equation. Further, Hansen (1992) coefficient stability L test results given above reveal both the stability of the individual coefficients and the joint stability of the coefficients and the estimated variance.

To illustrate the stability of single equation OLS findings, we present the recursive estimates of the model. For this purpose, we have initially computed residuals of the recursive least squares estimation as one-step ahead forecast errors by recursively re-estimating the equation over the period examined. Then, we present the plot of recursive residuals about a zero line for the error correction model derived from the co-integrating relationship. We have also applied the CUSUM test and the CUSUM of Squares tests, and the recursive estimates of the only-error correction coefficient as more and more of the sample data are used in estimation. Fig. 1 indicates that recursive residuals estimates do not reject parameter constancy, maybe except the 2001Q4 period. Neither CUSUM and CUSUM of Squares test

results nor recursive coefficient estimates support any evidence in favor of parameter or variance instability:

Figure 1. Recursive Estimates for the Linear Conditional Error Correction Model



3.5. Testing for Super Exogeneity and Marginal Modelling

In the system co-integration analysis carried out in the former sections, we find that there exist two weakly exogenous variables, i.e., equity prices (eq) and exchange rate (e), lying in the long-run variable space. Under the assumption of a single co-integrating vector derived from a constrained maximum likelihood problem, in Table 2, we are unable to reject the joint hypothesis that the two unrestricted adjustment or loading coefficients for these variables are zero, which is inferred from a $\chi^2=0.1567$ (prob. 0.9247) statistic. Further, no evidence against the strong version of exogeneity can be revealed for these variables. All these results enable us to take them as conditioning variables for statistical purposes in the money demand model, and serve to obtain additional inferences for the parameter vector of the variables without, otherwise, loss of information. However, such findings do not guarantee that these parameters upon which exogeneity or policy variable characteristic in an economics sense is imposed are

invariant to the class of interventions occurring during the sample period or that they can be reliable estimators against instabilities resulted from regime changes. This means that models based on backward looking behaviors would not be resulted in constant conditional models if they have been exposed to violent variations in the parameter structures. Considering these prerequisites for a complete modelling, Hendry and Ericsson (1991b) using a general-to-specific approach try to obtain constant error correction framework against the instability of money demand models and also apply to the super exogeneity tests to examine the invariance property in the sense of the Lucas critique.

In order to conduct a policy analysis, as Favero and Hendry (1992) point out, only through the researcher satisfies the property that the parameters of the model are invariant to changes in the distribution of the weakly exogenous conditioning variables can we arrive at the inference that these variables have also been of a super exogeneity form. This requires constructing marginal models using these variables on which the money demand model has been conditioned, then, the super exogeneity assumption can only be rejected if the constructed variables derived from the marginal models have no statistical significance on the conditional model. Provided that the super exogeneity property can be obtained the researcher will gain a more robust degree of autonomy in policy conducting analyses by changing the processes driving these policy variables.

Considering the encompassing implications of feedback versus feedforward models used in econometrics, Hendry (1988) provides an alternative assessment of the concept of super exogeneity. Briefly to say, feedforward (expectational) models that are based on the fact that much economic theories are constructed through intertemporal optimization problems emphasize the role of both expectations for the future outcomes of policies and feedbacks as determinants of current plans, while feedback (conditional) models assume that economic agents adopt to the complexity of real world by using past information. Hurn and Muscatelli (1992) attribute assessing the validity of these models to the obstacle of observational equivalence that deals with the problem that the reduced forms of a forward-looking model and a more conventional backward-looking error correction model coincide. This implies that different policy prescriptions can be suggested from an error correction framework. In this line of thought, Hendry (1988) and Favero and Hendry (1992) reveal that under regime changes, if super exogeneity is satisfied for the conditioning variables, the conditional model can be considered a feedback model. Such a model will encompass a whole class of rational expectations models and invalidate the Lucas' critique.

Following the explanations in the studies of Hendry (1988) and Engle and Hendry (1993), in this section, the super exogeneity tests are tried to be applied. To the best of my knowledge, Metin (1995) chosen to be followed in our estimation process is an extensive study that deals with the empirical findings of super exogeneity and the Lucas critique for the Turkish economy. To examine whether non-constancies in the marginal models for conditioning variables are able to affect the conditional model in a significant way, we first construct univariate fifth order autoregressive marginal models for the weakly exogenous contemporaneous variables eq_t and e_t . Then, we calculate the recursive residuals followed by some Chow instability tests and assign dummies for the periods that reflect structural breaks and instabilities included into the data. The calculated dummies and the residuals extracted from the marginal models are added into the conditional error correction model so as to test invariance and super-exogeneity, respectively. The parsimonious marginal models are given below (White HCSEs in parentheses),

Model1: Marginal $D(eq)_t$ equation

Redundant Variables $F(3, 40)$ -statistic= 0.2744 Prob. 0.8435

$$D(eq)_t = 0.0575 + 0.2693D(eq)_{t-1} - 0.3298D(eq)_{t-5} \\ (0.0245) (0.1208) \quad (0.1608)$$

Adj.R² = 0.2038, Std. Error of Regression = 0.1616, SSR=1.1228, F – stat.=6.7579 (Prob. 0.00), D-W stat.=1.9425,

B-G AR(1) F – stat.=0.1659 (Prob. 0.6860), B-G AR(4) F – stat.=0.5352 (Prob. 0.7107), J-B=0.0124 (Prob. 0.9938),

White F – stat.=3.0606 (Prob. 0.0045), ARCH(1) F – stat.=0.1738 (Prob. 0.6788), ARCH(4) F – stat.=0.9293 (Prob. 0.4576),

Ramsey RESET F – stat.=6.6164 (Prob. 0.0140)

Model2: Marginal $D(e)_t$ equation

Redundant Variables $F(3, 40)$ -statistic= 0.0781 Prob. 0.9715

$$D(e)_t = 0.0064 + 0.3340D(e)_{t-1} + 0.2714D(e)_{t-5} \\ (0.0136) (0.1118) \quad (0.0989)$$

Adj.R² = 0.1821, Std. Error of Regression = 0.0872, SSR=0.3270, F – stat.=6.0105 (Prob. 0.00), D-W stat.=2.0084,

B-G AR(1) F – stat.=0.0054 (Prob. 0.9417), B-G AR(4) F – stat.=0.1632 (Prob. 0.9557), J-B=60.2750 (Prob. 0.0000),

White F – stat.=0.6234 (Prob. 0.6827), ARCH(1) F – stat.=0.2475 (Prob. 0.6214), ARCH(4) F – stat.=0.3190 (Prob. 0.8634),

Ramsey RESET F – stat.=1.4047 (Prob. 0.2426)

The one-step residuals and one-step Chow test results using recursive least squares estimators are figured out to determine possible instabilities. In Figure 2, the recursive residuals lie outside the ± 2 S.E. bands for the 2008Q4 period with a coefficient -0.3716 (S.E. 0.1720), recursive one-step Chow tests yield significant F -stat. (1,4) = 66.640 (prob. 0.0012) for the 2001Q2 period and F -stat (1,34) = 5.4751 (prob. 0.0253) for the 2008Q4 period and

recursive breakpoint (N-down) tests result in respective F -stat. $(41,2) = 24.535$ (prob. 0.0399) for the 2000Q4 period and F -stat. $(39,4) = 7.3621$ (prob. 0.0319) for the 2001Q2 period. For the marginal $D(e)$ model, the recursive residuals again lie outside the ± 2 S.E. bands for the 2008Q4 period with a coefficient 0.2548 (S.E. 0.0943), recursive one-step Chow tests yield significant F -stat. $(1,3) = 31.377$ (prob. 0.0112) for the 2001Q1 period, F -stat. $(1,4) = 48.231$ (prob. 0.0023) for the 2001Q2 period and F -stat. $(1,34) = 9.6289$ (prob. 0.0038) for the 2008Q4 period, and finally recursive breakpoint (N-down) test result in respective F -stat. $(42,1) = 570.84$ (prob. 0.0332) for the 2000Q3 period, F -stat. $(41,2) = 131.28$ (prob. 0.0076) for the 2000Q4 period and F -stat. 47.637 (prob. 0.0042) for the 2001Q1 period:

Figure 1. Recursive Tests for $D(eq)$ equation

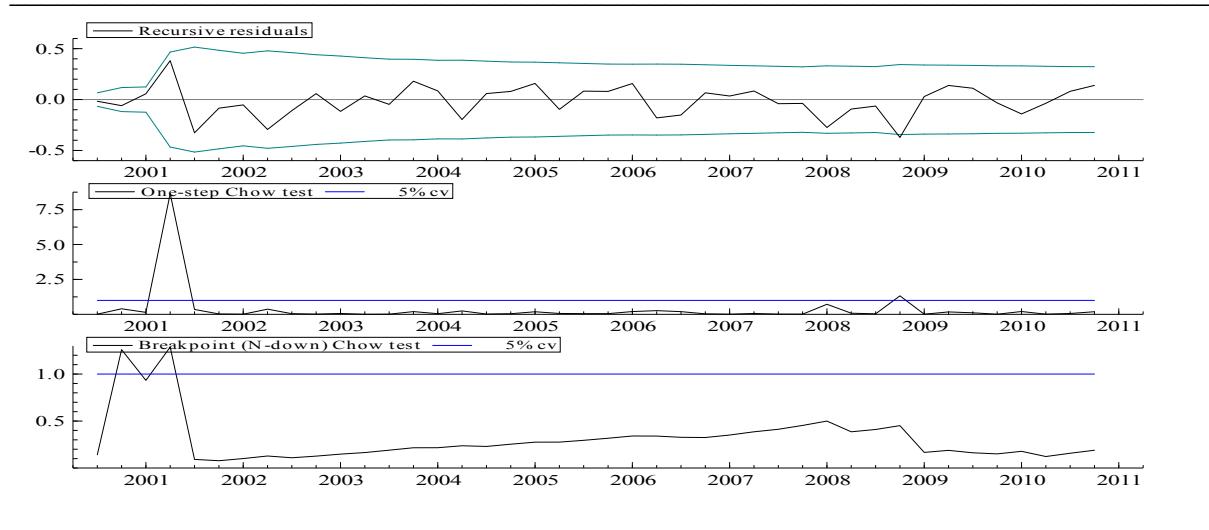
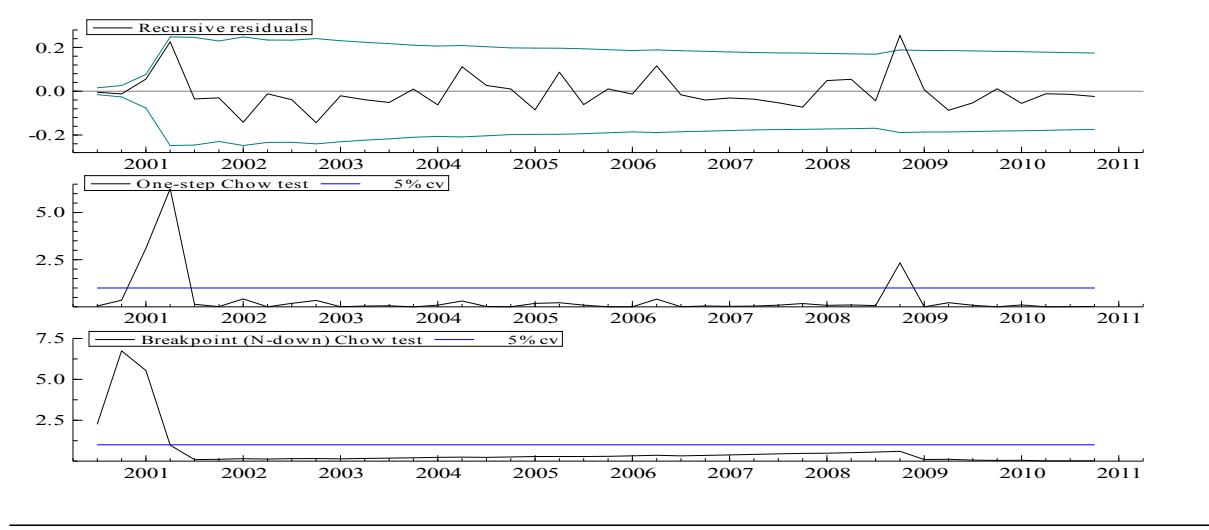


Figure 2. Recursive Tests for $D(e)$ equation



Then, we have re-estimated the marginal models by including zero / one shift dummies to account for instabilities:

Redundant Variables $F(4,38)$ -statistic= 0.2500 Prob. 0.9079

$$D(eq)_t = 0.0675 + 0.2397 D(eq)_{t-1} - 0.3229 D(eq)_{t-5} - 0.3947 mardummy2 \\ (0.0231) (0.1184) \quad (0.1592) \quad (0.0274)$$

Redundant Variables $F(3,38)$ -statistic= 0.2836 Prob. 0.8369

$$D(e)_t = -0.0081 + 0.3326 D(e)_{t-1} + 0.2329 D(e)_{t-5} + 0.1172 mardummy1 + 0.2730 mardummy2 \\ (0.0122) (0.1120) \quad (0.1123) \quad (0.0392) \quad (0.0733)$$

where *mardummy1* takes unity for the 2000Q3, 2000Q4, 2001Q1 and 2001Q2 periods and is zero otherwise, and *mardummy2* takes unity for the 2008Q4 period and is zero otherwise. As can be clearly noticed, the both dummies capture the upward tendency in exchange rate for the instability periods. Also, the negative impact on the stock exchange is highly evident in the 2008Q4 period. Having estimated these generated regressors, we follow Hurn and Muscatelli (1992) and carry out the super exogeneity tests by additionally including residuals extracted from the marginal models (res_{eq} and res_e , respectively) plus the dummies that account for the concept of invariance into the conditional error correction model in Table 5:

Table 5. Superexogeneity tests for the generated regressors in the conditional EC model

<u>Single Variable Wald Coefficient Tests</u>	<u>F-statistic</u>
(Restrictions are linear in coefficients)	
res_{eq}	$F(1,35) = 0.1230$ (Prob. 0.7279)
res_e	$F(1,35) = 0.9809$ (Prob. 0.3288)
<i>mardummy1</i>	$F(1,35) = 0.0007$ (Prob. 0.9791)
<i>mardummy2</i>	$F(1,35) = 0.7230$ (Prob. 0.4010)
<u>Joint Wald Coefficient Tests of Significance</u>	
$res_{eq} \wedge res_e$	$F(2,35) = 0.7261$ (Prob. 0.4910)
$mardummy1 \wedge mardummy2$	$F(2,35) = 0.4251$ (Prob. 0.6570)
<u>Joint Wald Coefficient Tests (for all constructed variables)</u>	
$res_{eq} \wedge res_e \wedge mardummy1 \wedge mardummy2$	$F(4,35) = 0.5327$ (Prob. 0.7125)

In the table, no evidence can be found in favor of the rejection of the super exogeneity. The results are highly robust to adding each constructed variable separately and jointly. We can infer that the conditional money demand error correction model is not affected in a statistically significant way by the residuals of the marginal models and the created dummies

for marginal model instability points. These results mean that the conditional error correction model has not been exposed to parameter non-constancies that yield evidence for the Lucas critique and that it can be considered a feedback model which is able to encompass a whole class of expectation models.

3.6. Testing for Non-Linearity in the EC Equation

As a final stage in our applied paper of the Turkish narrow money demand, we will try to test the linearity of the baseline parsimonious error correction equation against the relevant non-linear STR version. In the original equation above, we find that the RESET test cannot reject the null hypothesis of no mis-specification in the parsimonious linear EC model. If not so, the model would be able to require a higher order polynomial functional form inclusive of, e.g., quadratic and/or cubic terms. As Granger and Teräsvirta (1993) emphasize, this test can be considered an LM test of linearity against a very special modelling approach of the STR type. Thus, we further aim to conduct some tests of the logistic STR model to control the data consistency, so we can infer in a more robust way whether the estimation procedure can be improved by additionally modelling non-linearity.

For this purpose, we first examine the null hypothesis of linearity for the parsimonious conditional error correction model and test that $H_0 : \beta_1 = \beta_2 = \beta_3 = 0$ against the alternative non-linearity hypotheses. Note that the constant term and centered seasonal dummies are allowed to enter the linear part of the model as is in the original error correction representation. Each of the variables EC_{t-1} , $D(i)_{t-1}$, and $D(eq)_{t-1}$ are used as transition variables in turn. The results are reported in Table 6 below:

Table 6. *p*-values of the Linearity Tests of the Conditional EC Equation

Testing Linearity against STR					
Variables in AR part: constant EC_{t-1} $D(i)_{t-1}$ $D(eq)_{t-1}$ D_Q2 D_Q3 D_Q4					
Sample range: [1998Q2 2010 Q4], $T=51$					
<i>p</i> -values of F-tests					
Transition Variable	H_0	H_{04}	H_{03}	H_{02}	Suggested Model
EC_{t-1}	2.5198e-01	5.6440e-01	1.1946e-01	3.1053e-01	Linear
$D(i)_{t-1}$	3.6151e-01	6.3549e-01	4.7655e-01	9.5739e-02	Linear
$D(eq)_{t-1}$	8.7555e-01	6.6726e-01	8.4998e-01	5.8017e-01	Linear

In the table, we do not succeed in rejecting the H_0 hypothesis since the relevant p -values for all the possible transition variables cannot reach the acceptable significance levels to keep on an analysis of non-linearity in our applied modelling approach. Thus, we choose to retain the specification of the linear error correction model of the money demand as the warranted time series estimation process against the alternative STR modeling attempts. However, it must be stressed out that due to the preliminary investigation of the time series in the former sections, we do not include inflation data realizations into the estimation process, but various other model specifications different than the application in this study can make a model formulation inclusive of the inflation rate possible as the alternative cost variable of money demand.

4. Conclusions, Policy Discussion and Suggestions for Future Researches

One of the most important issues of interest for today's macroeconomic debates is to reveal the stability of functional relationships upon which economic theories are constructed and tested by using popular estimation techniques. If constancy of the models cannot be provided, the *ex-ante* model evaluation process will not possibly reflect the true data generation process to test the motives used for decision making of economic agents and to infer what policies are appropriate for stabilization purposes under the whole periods examined. In this paper, we have tried to examine such a policy issue mostly highlighted by the Lucas critique in the economics literature for the narrowly defined money demand relationship in the Turkish economy. We have thus aimed at testing whether the money demand relationship has been exposed to the structural breaks and parameter instabilities that give rise to regime changes.

Using multivariate co-integration estimation methodology, the findings confirm the empirical success of the theory to explain the behavioral foundations of aggregate data-supported monetary economics approaches. In this sense, under the money market equilibrium conditions, the real income elasticity which is larger than unity indicates that the demand for real money balances can be considered like a demand for luxury goods, which will be resulted in declining monetary velocity in a long run perspective. All the alternative costs variables which are chosen as the Treasury interest rate, equity prices and exchange rate have statistical significance with expected signs. The model is found to have a highly robust characteristic against the possible endogenous breaks. These inferences are also verified by the conventional general-to-specific single equation modelling of the short run dynamics of the money demand relationship, which enable us, to a greater extent, to learn from the data so

as to improve our empirical model. Of all the variables, the weak exogeneity property in a system co-integration relationship cannot be rejected for the equity prices and exchange rate, thus, we have applied marginal modelling procedure constructed on these variables to test whether they are really of the conditioning form in the VEC estimation. The results reveal that no evidence can be found in favor of the rejection of the super exogeneity property of the equity prices and exchange rate variables which means that the conditional EC model has not been exposed to in-sample parameter non-constancies that yield evidence for the Lucas critique and that it can be considered a feedback model which is able to encompass a whole class of expectation models. The linearity in the conditional error correction model cannot be rejected as the warranted time series estimation against the alternative non-linear models.

If theory-backed empirical models are able to be estimated by the researchers, these findings will give a change to the policy makers in order to have a foresight for the possible outcomes of *ex-ante* designed policies. Such an inference has been of a special importance especially in monetary theory analyses. In this line of thought, since we are succeed in estimating a stable narrow money demand model which is not fortunately subject to the contemporaneous Lucas critique mainly directed to the econometric applications of the researchers, we can state that our findings tend to increase the autonomy of policy makers at first to control the course of the monetary aggregates and then to use them for various stabilization purposes, e.g. in fighting inflation. But, the readers do consider the issue that our findings in the paper reflect a general tendency of the data restricted for our sample period. However, and of course if possible, the extension of the time series for the earlier and later periods can yield results violating super exogeneity and invariance property of the estimated model due to possible policy regime changes. We think that this is the critical issue to understand the main theme of this paper and to derive various policy outcomes from such empirical studies.

In consequence of our findings, complementary papers should be constructed also for broadly defined money demand relationships so as to further verify whether monetary-based theory of economics are able to succeed in tracking down the real data generation processes of the estimation techniques. In this sense, as is briefly expressed above, these studies are suggested to be interested in yielding the necessary knowledge of inflation and money demand relationship. If this task can be implemented, the monetary business cycles of the economy will be more explicitly revealed and the data consistency of the models will enable the policy makers to enforce their hands in applying discretionary monetary stabilization policies.

In this empirical economics paper, a set of econometric procedures available in the software programs EViews 6.0., STATA 9.0., Gauss 10.0., PcGive 10.40. and JMulTi 4.24 are tried to be used for econometric modeling purposes.

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APPENDIX

Table 2. Multivariate co-integration analysis (restricted linear deterministic trend)

Null hypothesis	$r=0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r \leq 4$	
Eigenvalue	0.5705	0.3588	0.3118	0.2060	0.1312	
λ trace	101.72 *	59.467	37.246	18.562	7.0299	
5% cv	88.804	63.876	42.915	25.872	12.518	
λ max	42.253 *	22.221	18.684	11.533	7.0299	
5% cv	38.331	32.118	25.823	19.387	12.518	
<u>Unrestricted Co-integrating Coefficients</u>						
m	y	i	eq	e	$trend$	
-6.7997	16.096	-12.022	-4.1417	-1.4793	0.0895	
-1.3550	13.807	-2.9249	-2.2768	2.8921	-0.0544	
20.534	-11.430	0.4590	-3.6021	0.6989	-0.3287	
-1.7006	26.281	7.2356	0.2651	1.9717	-0.1553	
-6.8030	6.5815	0.5580	-2.0099	0.9278	0.1997	
<u>Unrestricted Adjustment Coefficients</u>						
D(m)	0.0108	-0.0001	-0.0090	-0.0099		
D(y)	-0.0100	0.0016	0.0012	-0.0111		
D(r)	0.0420	-0.0068	0.0118	-0.0005		
D(eq)	-0.0063	0.0311	0.0447	0.0005		
D(e)	-0.0030	0.0311	0.0447	0.0005		
<u>1 Co-integrating Equation (standard errors are in parentheses)</u>						
m	y	i	eq	e	$trend$	c
1.0000	-2.3672 (0.6581)	1.7680 (0.2864)	0.6091 (0.1255)	0.2175 (0.0932)	-0.0132 (0.0067)	18.40
<u>Adjustment Coefficients (Standard errors are in parentheses. 'D' is the difference operator)</u>						
D(m)	D(y)	D(i)	D(eq)	D(e)		
-0.0731 (0.0339)	0.0679 (0.0304)	-0.2857 (0.0869)	0.0428 (0.1859)	0.0202 (0.0755)		

Notes: * denotes rejection of the null hypothesis of no co-integration at 0.05 level

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