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This paper is circulated for discussion purposes only and its contents should be considered preliminary.

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Roberto A. De Santis*

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ABSTRACT

This paper employs the multivariate cointegration technique of Johansen to construct a price equation for Turkey over the period 1950 to 1991. In addition, a monetary model in error correction form has been formulated to explain the inflationary process in the short run. Adaptive and rational expectation hypotheses are introduced and tested one against the other. Several test statistics suggest that the ECM under adaptive expectation hypothesis is to be preferred, matching all salient features of the data. The model indicates that the per-capita money supply, the per-capita real income and the difference between the opportunity cost of holding a unit of money and its return are the cardinal variables in explaining the price dynamics in the Turkish economy. The model has very good static forecasting properties, successfully predicting the peak of the inflation rate in 1980 and the subsequent trended increase from 1982 onwards.

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1. INTRODUCTION

Inflation in Turkey started its erratic and explosive phase during 1977-1980. Annual inflation measured by the consumer price index accelerated from 28 per cent in 1977 to 116 per cent in 1980. In the same period, a grave decline in the rate of economic growth was recorded. Real GNP grew by over 7 per cent a year between 1974 and 1976. The growth rate fell to 3 per cent in 1978 and was negative in both 1979 and 1980. The stabilisation and export-led recovery policies employed at the beginning of the '80s, following the liberalisation process of the Turkish economy, helped to reduce inflation and to promote a per-capita real growth. However, the government was not able to control the inflation rate, which has followed a rising trend since 1982.

In the last years, several papers have been published trying to explain the dynamics of prices in Turkey. Fry (1986) and Togan (1987) explained the inflationary process with a monetary model, in which the short run money demand has been specified by a partial adjustment model; Onis and Ozmucur (1990) tried to explain the high and volatile rates of inflation with a vector autoregressive model, in which high-powered money, price level, exports and the exchange rate were simultaneously considered; Uygun (1992) attempted to explain the behaviour of aggregate inflation and growth in the Turkish industry on the basis of firms' microeconomic decisions; Ozatay (1992b) employed a model emphasising the prices of goods produced by public enterprises in the dynamics of Turkish inflation; the same author (1992a) has also showed that excess money supply plays an important role in explaining the short run dynamics of price, interest rate and exchange rate variables.

The purpose of this paper is to explain the inflationary process in Turkey from 1950 to 1991 with a monetary model expressed in error correction form, using two different hypotheses of expectation formation: adaptive expectation (AE) and rational expectation (RE) hypotheses. The study consists of six sections. Sections 1 and 2 describe the monetary model expressed in error correction form and the models of expectations formation. Section 3 defines the variables and the data employed in the econometric study. Section 4 presents the unit root tests and the cointegrating analysis. Section 5 reports the estimates of the coefficients of the error correction model (ECM) under AE and RE hypotheses and several test statistics to examine the properties of the two models. Section 6 provides a summary and some conclusions.

2. THE MODEL

To understand the price dynamics in the context of a monetary model, the determinants of nominal money supply and real money demand must themselves be examined. Provided that the money market clears within the time period under consideration, the equilibrium condition in the money market can be expressed as

$$M^s = M^d$$

or

$$M^s = P \cdot n \cdot m^d \qquad (1)$$

where M^s is the nominal money supply, M^d the nominal money demand, P the price level, n the population, and m^d the per-capita demand for real money balances.

Equation (1), expressed in first difference logarithmic form, results in

$$\Delta \ln M^s = \Delta \ln P + \Delta \ln n + \Delta \ln m^d$$

which can be rewritten as

$$\pi = \Delta \ln(M^s/n) - \Delta \ln m^d \qquad (2)$$

where π is the continuously compounded rate of change in the price level.

Often, in theoretical money demand equations, researchers employ a measure of anticipated per-capita real activity (Y^e) and the discrepancy between the opportunity cost of holding a unit of money and its return. Calvo (1992), for example, employed the difference between the nominal interest rate on loans and the nominal interest rate on deposits. Not having sufficient data on the above variables, we use the expected inflation rate (π^e) as proxy of the opportunity cost of holding a unit of money (see Cagan, 1956), and the discount rate of the Central Bank (d) as proxy of its return. Hence, the long run or desired per-capita real money demand function takes the form:

$$m^* = a \cdot Y^{e^b} \cdot \exp\{c(d - \pi^e)\} \tag{3}$$

Calvo (1992) assumes that the demand for money is a positive function of the difference between its return and its opportunity cost. In addition, Cagan (1956), in a context of hyperinflation, argues that the demand for real money balances is negatively related with the expected rate of change in prices; which implies "c" to be positive. On the other hand, $b > 0$; that is, the higher the level of real activity anticipated for the near future, the more real money is demanded to accommodate the larger volume of generated transactions.

In the partial adjustment model, if long run desired equilibrium is disturbed by a change in the real activity or the interest rate, then adjustment costs cause individuals to adjust slowly towards the new long run equilibrium level of desired real money balances. In the short run, only a fraction of the difference between the desired and actual stock of real money balances is adjusted in each period. In a generalised version of the partial adjustment hypothesis, that is the error correction model (ECM), the adjustment of the individual's money demand is originated by two components: the change in the desired level (i.e. adjustment cost) and the past period disequilibrium (i.e. disequilibrium costs):

$$lrm_t^d - lrm_{t-1}^d = \theta\left(lrm_t^* - lrm_{t-1}^*\right) + \gamma\left(lrm_{t-1}^* - lrm_{t-1}^d\right) \tag{4}$$

where $0 \le \theta \le 1$ and $0 \le \gamma \le 1$. If $\theta = \gamma$, we obtain the partial adjustment model. The partial adjustment model is encompassed in the ECM.

Substituting equation (3) into equation (4), and assuming the equilibrium condition in the money market, we obtain

$$\Delta \ln m_t^d = b\theta\Delta \ln Y_t^e + c\theta\Delta\left(d_t - \pi_t^e\right) - \ln\left(M^s/n\right)_{t-1} + \ln P_{t-1} + \\ +\gamma\left[\ln a + b\ln Y_{t-1}^e + c\left(d_{t-1} - \pi_{t-1}^e\right)\right] \qquad (5)$$

Substituting equation (5) in the identity (2), we obtain

$$\pi_t = \Delta \ln\left(M^s/n\right)_t - b\theta\Delta \ln Y_t^e - c\theta\Delta\left(d_t - \pi_t^e\right) + \\ -\gamma\left[\ln p_{t-1} - \ln\left(M^s/n\right)_{t-1} + \ln a + b\ln Y_{t-1}^e + c\cdot\left(d_{t-1} - \pi_{t-1}^e\right)\right]$$

On the right and side, we can recognise the error correction term of the long run relationship between price, money supply and the determinants of the money demand. As suggested by the economic theory, we might assume that expected values are equal to the actual values in the long run. Thus, we can write the price equation for the ECM as:

$$\pi_t = \alpha_0 + \alpha_1\Delta \ln\left(M^s/n\right)_t + \alpha_2\Delta \ln Y_t^e + \alpha_3\Delta\left(d_t - \pi_t^e\right) + \alpha_4 Z_{t-1} + u_t \qquad (6)$$

where $\alpha_2 = -b\theta$, $\alpha_3 = -c\theta$, $\alpha_4 = -\gamma$, Z_{t-1} is the short run error correction term of the cointegrating relationship and u_t is a white noise error. α_1, α_2, α_3 and α_4 are the adjustment parameters in the ECM and correspond to the short run dynamic coefficients. α_1, α_2, and α_3 represent the dynamic price elasticities on per-capita money supply, anticipated per-capita real activity, and interest rate differential, respectively. α_4, however, represents the effect in the short run of a deviation from the long run relationship between price, per-capita money supply and the determinants of the per-capita real money demand. If the variables are all integrated of the same order and are cointegrated, we might expect α_4 to be negative, indicating that, if the price level is above the long run value, the overall effect is to slow down the short term inflation.

3. MODELS OF EXPECTATIONS FORMATION

We study the price dynamics in Turkey by considering two different models of expectation formation for both Y^e and π^e: adaptive expectations (AE) and rational expectations (RE).

AE hypothesis assumes that economic agents adapt their expectations in the light of past experience and that they can learn from their mistakes. More formally AE may be formulated as follows:

$$X_t^e - X_{t-1}^e = (1-\lambda)(X_t - X_{t-1}^e)$$

This equation shows that expectations are revised each period by a constant fraction $(0 \leq 1-\lambda \leq 1)$ of the discrepancy between the current observed value of the variable and the previously made expected value of that variable. In other words, expectations are revised by a fraction of the most recent error.

The introduction of AE leads equation (6) to be written in reduced-form as

$$\pi_t = \beta_0 + \beta_1 \Delta \ln(M^s/n)_t + \beta_2 \Delta \ln(M^s/nP)_{t-1} + \beta_3 \Delta \ln Y_t + \\ + \beta_4 \Delta(d_t - \pi_t) + \beta_5 \Delta(d_{t-1} - \pi_t) + \beta_6 Z_{t-1} + \beta_7 Z_{t-2} + v_t \qquad (7)$$

where the β's are the coefficients to be estimated and v_t is the error term. Although v_t is equal to $(u_t - \lambda u_{t-1})$, OLS is a consistent estimator of the parameters, because the equation (7) does not involve π_{t-1}. The structural coefficients can be calculated from the reduced-form coefficients as follows:

$$\beta_2 = -\lambda, \quad \beta_3 = -b\theta \cdot (1-\lambda), \quad \beta_4 = -c\theta, \quad \beta_5 = c\theta\lambda, \quad \beta_6 = -\gamma, \quad \beta_7 = \lambda\gamma$$

If, instead, we assume that economic agents are rational, in the sense they do not make systematic mistakes in their predictions, we create contemporaneous correlation between regressors and the error term. Namely, if we define

$$\Delta \ln Y_t = \Delta \ln Y_t^e + \varepsilon_t$$

and

$$\pi_t = \pi_t^e + \eta_t$$

we can substitute them into equation (6), so that

$$\pi_t = \alpha_0 + \alpha_1 \Delta \ln\left(M^s/n\right)_t + \alpha_2 \Delta \ln Y_t + \alpha_3 \Delta\left(d_t - \pi_t\right) + \alpha_4 Z_{t-1} + u_t - \alpha_2 \varepsilon_t + \alpha_3 \Delta \eta_t \qquad (8)$$

The assumption of orthogonality between regressors and disturbance term is now violated. To solve this problem, we need to instrument $\Delta \ln Y_t$ and π_t, and use their predictors to estimate the coefficients by OLS. The standard errors of all coefficients are incorrect due to the problem of generated regressors (Pagan, 1984). To construct correct standard errors for the coefficients, we use the Instrumental Variable (IV) estimation based on actual values, with the variables used to construct the predictors acting as instruments. The main characteristic of the instruments is that they must be strongly correlated with the instrumented variables, and independent of the error term.

4. THE DATA

In order to estimate the inflationary process in Turkey from 1950 to 1991, the chosen index of the price level is the Consumer Price Index (CPI). The money supply is proxied by the sum of the narrow money (M1), which includes demand deposits and currency in circulation, plus the quasi money, which includes time savings and foreign currency deposits. M1, plus quasi money, gives a broader measure of money similar to that which is frequently called M2. We prefer to use M2, instead of narrow money as proxy of the money supply, because economic agents in Turkey hold foreign currency deposits to protect their savings from the strong domestic inflation, especially when the liberalisation of the Turkish economy took place after 1980. Not having data on the interest rate on deposits from 1950 to 1972, and in order not to loose the 55 per cent of observations, we preferred to employ the discount rate of the Central Bank as its proxy, in view of the fact that there is a strong positive relation between the discount rate and the interest rate on deposits[1]. With regard to the per-capita real activity in the money demand function, the difficulties in measuring wealth or the transactions

[1]The coefficient of determination between the Central Bank's discount rate and the interest rate on deposits is 0.88, in the period 1973-1991.

demand for money lead economists to use some measure of income as the activity variable in empirical money demand function. The activity variable used in this study is the per-capita real GNP, calculated as the ratio of nominal GNP and the product between population and GNP-deflator.

5. UNIT ROOT TESTS AND COINTEGRATING ANALYSIS

In order to estimate the price equation in error correction form, we must test that the variables are integrated of the same order and that the cointegrating residuals are stationary. Namely, we need the error correction formulation to be I(0). Thus, it is fundamental to verify if the variables are generated by a difference-stationary (DS) or a trend-stationary (TS) process. Therefore, we estimate the theoretical unrestricted model, including the time trend "t",

$$\Delta \ln X_t = \alpha + \rho \ln X_{t-1} + \xi \cdot t + \phi \Delta \ln X_{t-1} + \varepsilon_t \qquad (9)$$

and test the null hypothesis $\rho = \xi = 0$ (DS), against the alternative $\rho < 0$ (TS), by using the Augmented Dickey-Fuller test (Dickey and Fuller, 1981) and the Perron test (Perron, 1989). For the price level and the per-capita money supply, ADF cannot be rejected at 10% level of significance. The above tests applied to income and interest rate variables cannot be rejected even at 1% level of significance. In addition, the residuals of all estimated equations appear to be white noises. Thus, we are unable to reject the hypothesis of a unit root for any of the variables under analysis at 10% level of significance.

Two main approaches are currently used to test for cointegration. One is the residual-based ADF method proposed by Engle and Granger and the other is the Johansen's Full Information Maximum Likelihood (FIML) approach. Johansen (1988) argues that the residual-based cointegration tests are inefficient and can lead to contradictory results, especially when there are more than two I(1) variables under consideration. He suggests employing the FIML procedure, which provides a unified framework for estimation and testing of cointegrating relations in the context of vector autoregressive (VAR) error correction models. A VAR model can be written in error correction form as

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \cdots\cdots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + BW_t + \varepsilon_t \qquad (10)$$

where X_t is an $p \times 1$ vector of I(1) variables, W_t is an $n \times 1$ vector of I(0) variables and $\varepsilon_t \rightarrow N(0, \Sigma)$. The Johansen procedure estimates (10), subject to the hypothesis that the long run parameter vector Π has a reduced rank, $r < p$, where "p" represents the order of the VAR model. The rank of Π gives the number of cointegrating relationships. Π can be decomposed as $\Pi = \alpha\beta'$, where α are the parameters on the error correction terms, and β are the coefficients of the cointegrating relations.

Before testing for cointegration, we make inference on all variables to specify the order of the underlying VAR model. The diagnostic tests of the regressions of each variable in level (i.e. endogenous variable) on the others (i.e. exogenous lagged variables) suggest the order of the VAR model being one.

Table 1 provides log-likelihood ratio test statistics for determining the number of cointegrating vectors "r". Starting with the null hypothesis of r = 0, the maximum eigenvalue statistic is 121.9 which is well above the 90% critical value of the test and hence strongly rejects the null of no cointegration among the four variables. Moving to the second row of the table, the maximum eigenvalue statistic is computed to be 13.2, which is below the 90% critical values. Hence, the hypothesis that there is one cointegrating vector against the alternative that there are two cointegrating vectors cannot be rejected, and it is reasonable to conclude that there is only one cointegrating relationship linking together $\ln P_t$, $\ln(M^s/n)_t$, $\ln Y_t$, and $(d_t - \pi_t)$. The test results based on Johansen's trace statistic confirm the conclusion that there is only one cointegrating relationship amongst the four variables.

Figure (1) shows the cointegrating residuals. They are stationary in the period 1951-76, but trended in the next period, 1977-91. This is essentially due to the huge disequilibrium, which hit the Turkish economy at the end of 1970's, caused by the oil

Table 1 Testing the rank of Π

λ_{max}				Trace				
H_0	H_1	Stat.	90%	H_0	H_1	Stat.	90%	λ
r=0	r=1	121.90	24.73	r=0	$r \geq 1$	143.56	43.95	0.95
$r \leq 1$	r=2	13.17	18.60	$r \leq 1$	$r \geq 2$	21.65	26.78	0.27
$r \leq 2$	r=3	7.70	12.07	$r \leq 2$	$r \geq 3$	8.48	13.32	0.17
$r \leq 3$	r=4	0.78	2.69	$r \leq 3$	$r \geq 4$	0.78	2.69	0.02

shock, the lack of foreign exchange, the state coup and the liberalisation of the Turkish economy. We tried to resolve the non-stationarity problem of the cointegrating residuals by introducing a linear trend in the long run relationship (Johansen and Yuselius, 1990; Yuselius, 1991). It improves the stationarity of the cointegrating residuals; leaves the number of the cointegrating vectors unchanged, as one; and slightly effects the magnitude of the long run coefficients. However, the estimated ECMs with the detrended cointegrating residuals have poor forecasting properties. Thus, as the disequilibrium is likely to be temporary, we employed the residuals of the above cointegrating relation as the error correction term in the ECM.

Table (2) reports the estimation of α and β vectors, and the corresponding t-statistics. The parameters of the long run relationship are significantly different from zero. The log-likelihood ratio statistic on α's indicates that the real per-capita income and the real interest rate can be treated as weakly exogenous to the rest of the system of equations at 10 and 2.5 per cent level of significance, respectively. However, although we have two different ECMs, we focus our attention only on the ECM corresponding to the price equation.

All the variables have a significant effect on the price level, in the long run. The simultaneous increase of the per-capita money supply and real income by 10% induces the price level to decrease by 10%. The interest rate differential has a big impact on the price level. Its increase by 1% leads to a decrease by almost 7% of the price level in the long run. The coefficient of the error correction term computed in the Johansen estimation is equal to -0.107, suggesting that when the price level is above the long run equilibrium by 10%, inflation decreases by 1% in the short run.

Table 2 α and β vectors

Variables	α	β
LP	- 0.107 (6.939)	- 1.000 (1.989)
LM2P	-0.091 (7.454)	1.685 (3.194)
LRYP	0.008 (1.054)	- 2.677 (3.692)
RI	0.024 (1.987)	- 6.871 (8.151)

6. ECM UNDER AE AND RE HYPOTHESIS

Equation (7) is estimated introducing a dummy variable to take into account the structural breaks in 1968 and in 1980. The variable deletion test on the variables not significant cannot be rejected at 5% level of significance. The re-estimation of the model presented in table (3) leads to very good results, with R^2 equal to 0.97. The diagnostic tests are accepted and the coefficients are significant at 5% level of significance. The coefficients of the per-capita money supply and per-capita real income are almost equal, obviously with the opposite sign. The main conclusion is that, if the Turkish government expands money supply to increase the real GNP, given the Keynesian multiplier less than one, the monetary policy will have an inflationary effect in the short run. The coefficient of the interest rate differential has a negative effect on inflation. Its size is more than double that of $\Delta \ln Y_t$'s coefficient, as also in the long run relationship. The model is just-identified, if the long run values "b" and "c" are taken into account. $\gamma = -\beta_6 = 0.09$, which is close to the value of the error correction term computed with the Johansen estimation. θ and λ can be calculated as follows:

$$\theta = -\beta_4/c = 0.11, \qquad \lambda = 1 + \beta_3/b\theta = -0.16$$

γ's and θ's values are close to zero, but significantly different from zero, suggesting that if the long run desired equilibrium of desired real money balances is disturbed by a unit change in the per-capita real GNP or the interest rate differential, then adjustment and disequilibrium costs cause individuals to adjust slowly towards the new long run equilibrium level. λ presents a negative sign, which contradicts the economic theory. However, the Wald test restriction $\lambda = 0$ is strongly accepted, indicating, therefore, that economic agents revise their expectations rapidly to recent data. This also implies that the residuals in equation (7) are serially incorrelated.

Table 3 OLS estimates of the ECM under AE hypothesis

Variables	Coefficients	Stand. Errors	Statistics	
Constant	1.3776	0.1465	R^2	0.9669
DM2P	0.3171	0.0826	σ	0.0349
DRYP	-0.3317	0.0958	LM(1)	2.6464
DRI	-0.7371	0.0579	FF(1)	0.0137
DU6880	-0.0768	0.0291	N(2)	0.2430
$-\beta' X_{t-1}$	-0.0910	0.0094	H(1)	1.6264

Note: LM(1) is the Lagrange multiplier of residual serial correlation of order 1 and is asymptotically distributed as χ^2_1; FF(1) is the RESET test and is asymptotically distributed as χ^2_1; N(1) is the Jarque-Bera test for normality and is asymptotically distributed as χ^2_2; H(1) is a test for heteroscedasticity and is asymptotically distributed as χ^2_1.

With regard to testing of parameter variation, we test for structural change within the sample (i.e. Chow test) and for adequacy of prediction (i.e. Chow's second test) in two different years 1968 and 1984. As it was verified that the null hypothesis of homoscedasticity cannot be rejected in the two different sub-periods, the Chow test is strongly accepted, while the predictive failure test cannot be rejected at 5% level of significance. In addition, the plot of actual and static forecasts indicate that the model is stable and useful for predictions (Fig. 2-3). The simulated inflation rate predicts successfully the peak of the inflation rate in 1980, the subsequent reduction in 1981, and the trended increase from 1982 onwards. It is likely that the extraordinarily good forecasts from 1969 to 1991 are based on the large negative value in the cointegrating residuals in 1980. However, the accurate forecasts from 1985 to 1991 indicate that the model is stable and powerful in making predictions.

In estimating equation (8), we need to define appropriate instruments of the variables $\Delta \ln Y_t$ and π_t, in order to avoid the contemporaneous correlation problem between regressors and error term. We choose $\Delta \ln Y_{t-1}$ and a dummy variable which is equal to 1 in 1968 to offset the impact of an outlier for $\Delta \ln Y_t$; and the change of per-capita narrow money (M1) supply, the change of the effective exchange rate and a dummy variable which is equal to one in 1980, 1984 and 1988 for π_t[2]. The predictors of these variables are introduced in the equation (8), in order to estimate the coefficients of the variables and to test for serial correlation and heteroscedasticity by OLS (Pagan, 1984). The null hypothesis of the mentioned test is strongly accepted, by introducing a dummy variable which is equal to one in 1968 and 1980, years of instability in the

[2]Not having data on the effective exchange rate, we preferred to use the Special Drawing Right (SDR) as its proxy. As it is known, the SDR was equal to the US dollar until 1974; from 1974 to 1981, the SDR valuation basket consisted of the currencies of the countries whose share in world exports on goods and services averaged more than 1% in the period 1968-72; from 1981 to 1991, the currencies included in the basket were those of the five members having the largest exports of goods and services: US dollar, Deutsche mark, French franc, Japanese yen, and Pound sterling. 50% of Turkish imports comes from these 5 countries.

Table 4: OLS estimates of the ECM under RE hypothesis

Variables	Coefficients	Stand. Errors	Statistics	
Constant	1.1571	0.2032	R^2	0.8594
DM2P	0.3757	0.1095	σ	0.0720
DRYP	-0.6859	0.2167	LM(1)	0.2577
DRI	-0.4040	0.1170	FF(1)	2.9321
DU6880	0.1636	0.0441	N(2)	0.5927
$-\beta' X_{t-1}$	-0.0750	0.0132	H(1)	0.1036

Note: LM(1), FF(1), N(2) and H(1) are defined in Table 3; the Sargan's test is distributed as χ^2_4 and is equal to 4.5933.

Turkish economy. To construct correct standard errors for the coefficients and test for correct functional form and normal distribution of the error term, we employ the IV estimation based on the actual values (Pagan, 1984), with the variables used to construct the predictors acting as instruments. Table (4) shows the OLS estimates of the short run price equation under RE hypothesis and the correct standard errors of the IV estimation. All coefficients cannot be rejected at 5% level of significance. The Sargan's test of the goodness of the instruments and the diagnostic tests (serial correlation and heteroscedasticity by OLS, functional form and normality by IV) cannot be rejected. Given the long run values of "b" and "c", the model is over-identified. γ is equal to 0.075, and θ can be equal to 0.059 or 0.256. Testing for parameter variation within the sample cannot be rejected at 5% level of significance for the sub-periods 1969-91 and 1985-91. But, the predictive failure test for the same sub-periods does not strengthen the results obtained with this model, as confirmed by Figure 4. In conclusion the model is stable, but not useful for making static forecasts. As suggested by Pasaran (1982), for choosing the best model having the best properties, it is important to make non-nested tests between different models. In table

5, M1 represents the ECM under AE hypothesis, while M2 corresponds to the ECM under RE hypothesis. All test statistics suggest that the ECM based on adaptive expectation is to be preferred.

Table 5: Alternative tests for Non-Nested Regression Models

Test Statistic	M1 against M2	M2 against M1
N-Test	1.496 [.135]	-17.115 [.000]
NT-Test	1.419 [.156]	-14.458 [.000]
W-Test	1.540 [.124]	-7.069 [.000]
J-Test	-1.513 [.130]	11.275 [.000]
JA-Test	-1.734 [.083]	8.125 [.000]
Encompassing F(2,32)	2.836 [.073]	64.038 [.000]
Akaike's Information Criterion of M1 versus M2 = 28.934 favours M1		
Schwarz's Bayesian Information Criterion of M1 versus M2 = 28.934 favours M1		

Note: M1 and M2 represent the ECMs under AE and RE hypothesis, respectively. In brackets, we report the probability of type I error.

7. CONCLUSIONS

We tried to explain the inflationary process in Turkey with a monetary model expressed in error correction form. It has been used to explain the relationship between prices, per-capita money supply, per-capita real income and the difference between the opportunity cost of holding a unit of money and its return. Assumptions on expectations of individuals were made and tested. Several test statistics suggest that the ECM with adaptive expectation hypothesis is to be preferred, matching all salient features of the data. The chosen dynamic model produces an acceptable set of past sample predictions using annual data. The data indicate that narrow money, plus quasi money (M2), is the relevant monetary variable in the inflationary process.

Considering the dynamic short and the static long run coefficients, we might maintain that the difference between the interest rate on money and the interest rate on loans has a fundamental role in controlling the inflationary process in Turkey. A rise in the interest rate on money has the same effect of a contractionary monetary policy in the short run. However, in the medium or long run, this policy may jeopardise the success of an inflation stabilisation program (see Calvo, 1992). Furthermore, the per-capita money supply effects the price level in the short as well as in the long run. Hence, the main economic conclusion is that, in considering the trade off between inflation and income-employment-poverty in Turkey, inflation should be controlled by limiting the growth rate of money and stimulating higher profitable investments. The Phillips curve proved a positive relation between higher employment-output and higher inflation. The negative relation between inflation and real activity seems more consistent with a money demand-quantity theory world, in which the increase of the anticipated real activity causes the increase of the real money demanded, which is not satisfied by nominal money growth and so must be accommodated by opposite movements in prices. We also found that the per-capita real income is a weakly exogenous variable to the rest of the system of equations. This implies that the Turkish government does not stimulate real growth by inflating. The theory and tests explain the phenomenon of stagflation, that is, the negative relation between inflation and real activity observed during the post-1950 period. In conclusion, the model presents results conforming with the economic theory and has good forecasting properties.

LEGEND

DM2P	- Per-capita Money Supply Change
DRI	- Interest Rate Differential Change
DRYP	- Per-capita real income change
DU6880	- Dummy Variable (1 in 1968 and 1980, 0 elsewhere)
LM2P	- Per-capita Money Supply Level
LP	- Price Level
LRYP	- Per-capita Real Income Level
RI	- Real Interest Rate
P	- Inflation

Note: The variables are in logarithmic form.

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Fig. 1 Plot of the Cointegrating Residuals

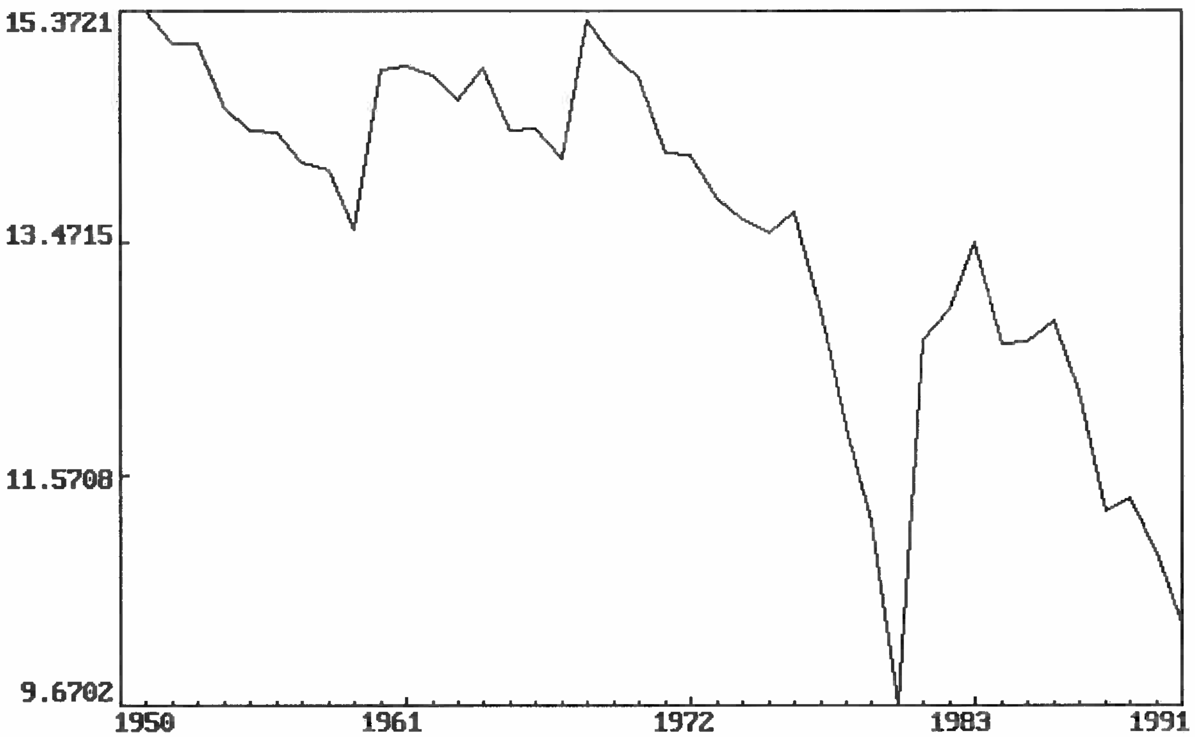

Fig. 2

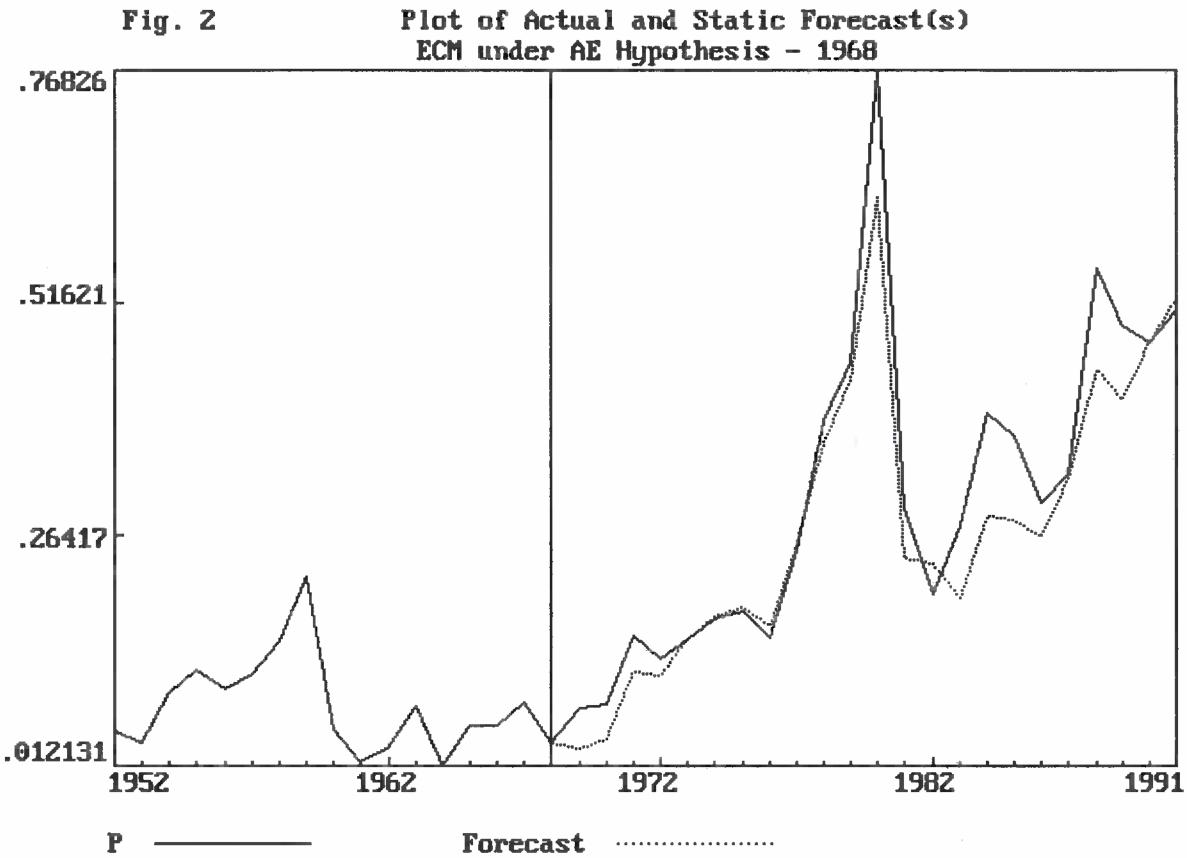
Plot of Actual and Static Forecast(s)
ECM under AE Hypothesis - 1968
.76826
.51621
.26417
.012131
1952
1962
1972
1982
1991
P
Forecast

Fig. 3 Plot of Actual and Static Forecast(s)
ECM under AE Hypothesis - 1984

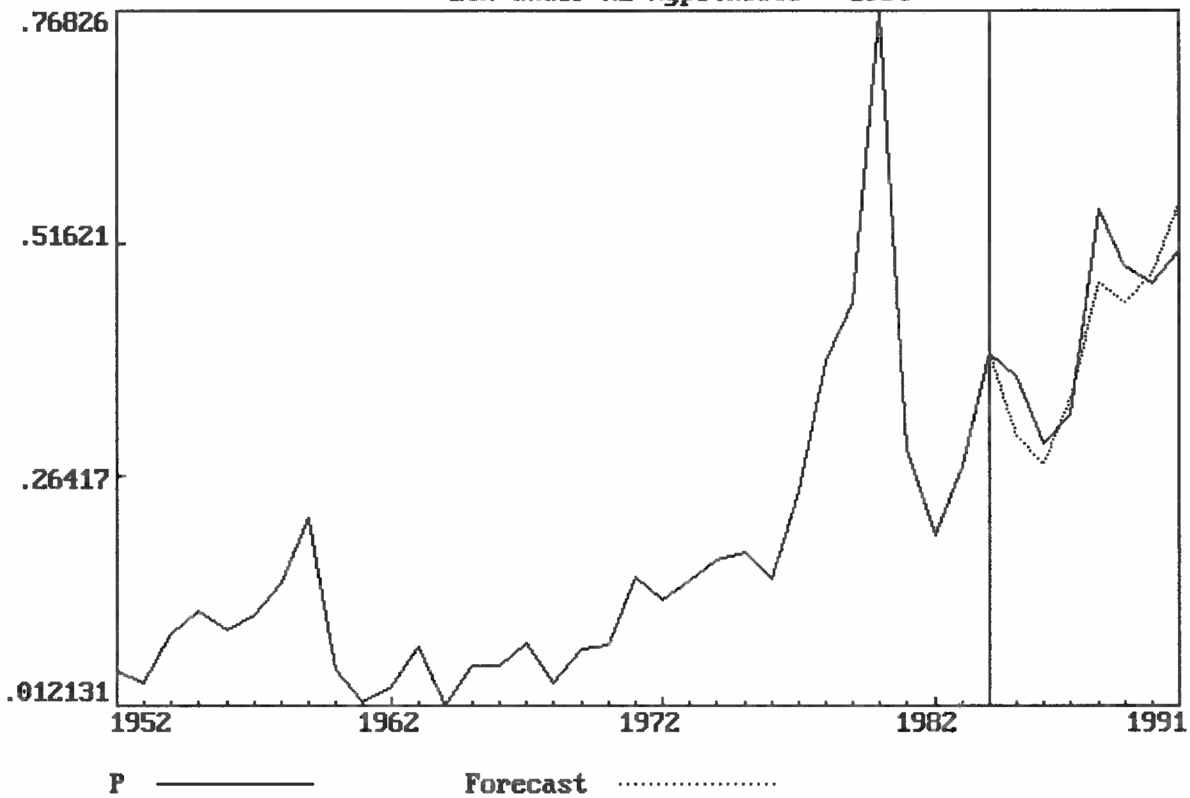

Fig. 4 Plot of Actual and Static Forecast(s)
ECM under RE Hypothesis - 1968