Structural VAR identification of the Turkish business cycles

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Abstract

In this paper, we investigate some of the main properties of the Turkish business cycles. Our empirical findings indicate that domestic inflation is countercyclical with real output and lags the GDP cycle by one quarter. We then construct a structural VAR model upon the Turkish economy, and estimate that the courses of real variables are mainly determined by the supply shocks, while both real and nominal shocks affect significantly the dynamics of the nominal variables.

Keywords: Business Cycles, Turkish Economy, SVAR Models

JEL Classifications: E30, E31,E60

I. Introduction

Estimating whether the price level and inflation are procyclical or countercyclical will provide policy makers with a knowledge of how properties must the stabilization policies have and provided that price level and inflation turn out to be countercyclical, supply-driven models of business cycles including real business cycle models will be appropriate to analyze the implications of business cycles (Chadha and Prasad, 1994: 240). Especially for a country case such as Turkey which had undergone an unstable real income growth process with anomalies in the course of real income and a chronic inflationary framework in a thirty years period till the early 2000s, such an analysis would be of special concern for the policy makers. In this sense, that the prices move in the same direction with output will point out the importance of demand side disturbances which enables discretionary Keynesian “leaning against the wind” type fiscal and monetary policy interventions (Alper, 2002: 22-54), whereas following e.g., Kydland and Zarazaga (1997: 21-22), supply-driven models can be based on real or supply-side factors which account for the business cycles, such as the amount of resources used by the government, tax policies, technological changes, government regulations, modifications of financial intermediation rules, and even political shocks signaling possible changes in property rights, rather than nominal factors such as the money supply, interest rates, and price rigidities employing a crucial role in the policy design and implementation of Keynesian and Monetarist business cycles.[1]

In line with such issues and from a policy perspective, use of potentially inappropriate conclusions regarding the stylized facts or broad regularities of macroeconomic fluctuations in different country cases can adversely affect the efficacy of stabilization policies. As Cashin (2004) expresses, economic policy is often contingent on whether or not a country is experiencing a cyclical contraction or expansion, and so it is vital that appropriate tools be used to extract the country-specific business cycle facts from the data. These all, of course, would compel the researchers to take into consideration the stylized facts of various country

[1] Following Ahmed and Park (1994: 2), in other words, if external and domestic supply disturbances are found to be important in explaining macroeconomic fluctuations and domestic aggregate demand disturbances are not, this would imply that policy makers’ attempts to fine-tune the economy will prove ineffective. As for the Turkish economy, effects of short-term capital inflows (outflows) and following appreciation (depreciation) of the real effective exchange rate, as expressed below, can lead to supply shocks affecting the business cycles in the economy.

cases so as to see whether boom-bust cycles in the level of real output resemble each other, and if so, similar stabilization policies can be advised to different country cases, but if not, different stabilization policies would be required for eliminating the pattern of fluctuations in economic activity.

In our paper, we try to give an essay upon the business cycles of the Turkish economy by both investigating the procyclical (countercyclical) characteristics of real domestic income and annualized inflation based on GDP-deflator and then constructing a structural vector autoregression (SVAR) model employing identification restrictions *a priori* assumed through the economics theory. For this purpose, the next section is devoted to the decomposition issues of real domestic income and inflation into their stationary cyclical components thus enabling us to reveal the cyclical characteristics of these aggregates between each other, and the section three extends our analysis by constructing a SVAR model including the effects of exogenous portfolio flows and the following changes in the real effective exchange rate on the business cycle pattern of the Turkish economy. And the final section summarizes our results and concludes.

II. Is the Turkish Inflation Procyclical?

Lucas (1977: 9) refers to that prices are generally procyclical as one of the commonly held beliefs among business cycle regularities, which leads to using equilibrium models with monetary policy or price surprises in the policy implementation process as the main source of fluctuations so that monetary disturbances would appear to be the only possible source of fluctuations. But contemporaneous literature considering different country cases upon this issue are able to yield conflicting estimation results revealing the countercyclical role of prices and inflation as a fact of business cycles. To deal briefly with empirical literature upon this issue, many studies touch on similar subjects for both developed and developing countries. For instance, Chadha and Prasad (1994: 239-257) and Fiorita and Kollintzas (1994: 235-269) find that price level is countercyclical for G-7 countries, while the former also find that inflation rate is procyclical thus suggest that the cyclical behavior of price level and inflation do not provide conclusive grounds for rejecting either demand-determined or supply-determined models of the cycle. Similarly, Kydland and Prescott (1990: 3-18) for the US, Backus and Kehoe (1992: 864-888) for 10 developed countries, Serletis and Krause (1996:

49-54) for the US, and Cashin and Ouliaris (2001) for Australia reveal the importance of countercyclical prices with output suggestive of predominance of shocks to aggregate supply in the economy. Besides, Lopez et al. (1997) estimate that for the case of Spanish business cycles inflation is mainly supply-driven and in this line suggest that strong disinflationary demand policies could prove both inefficient and very painful for Spain which needs more active supply policies.

Dealing with the developing country cases, Rand and Tarp (2002: 2071-2088) confirm the negative relationship between the price level and real income for a set of developing countries, providing support for a supply-driven interpretation of the business cycles including real business cycle models. Agénor et al. (1999) also find countercyclical variation of prices/inflation and cyclical component of output in many of the developing countries they examine, including Turkey such as Kydland and Zarazaga (1997: 21-36) for the cases of Latin American business cycles. Kim (1996: 69-82) estimates countercyclical relationships between the detrended price level and cyclical output for Korea and Taiwan but finds a positive correlation between inflation and cyclical component of output in line with Chadha and Prasad (1994: 239-257) considering G-7 countries. For the Turkish case, Alper (1998: 233-244), Metin-Özcan et al. (2001) and Alper (2002: 22-54) confirm the countercyclical pattern of fluctuations of the price level and inflation *vis-à-vis* real GDP.[2]

Various estimation methods have been come into use in contemporaneous economics literature to reveal the interactions between macroeconomic time series, such as structural vector autoregression models and decomposing the macroeconomic time series into their trend and cyclical components after linearizing them and using various filtering approaches as the mostly popular filter proposed by Hodrick and Prescott (1997: 1-16) trying to estimate the correlations between stationary cyclical series. We employ initially the latter type decomposing techniques to the Turkish data and so aim at extracting the cyclical characteristics of real domestic income and inflation between each other. For this purpose, we deseasonalize the annualized domestic inflation based on GDP deflator and the real gross domestic product (GDP) using U.S. Census Bureau's X12 seasonal adjustment program also available within EViews 5.1 and use multiplicative (ratio to moving average) method to

[2] Altuğ and Yılmaz (1998: 81-103) also estimate in their dynamic vector autoregression (VAR) modelling framework that shocks to inflation in Turkey would lead to a significant negative response in real activity proxied by industrial production.

extract the seasonal component, considering the period of 1988Q1-2006Q3 with quarterly data. Having deseasonalized the time series, we linearize them by taking natural logarithms to smoothen the changes in those since the business cycle literature is concerned with percentage deviations from trend in growing series (Kydland and Zarazaga, 1997: 33). The sample period has not been divided into sub-periods since, as Fiorito and Kollintzas (1994: 241) express, the smoothed trend should be able to capture the most important structural breaks. All the data are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT).

Following QMS (2004: 344-349), we apply in our paper to the widely-used Hodrick-Prescott (henceforth, HP) filter to obtain a smooth estimate of the long-term trend component of a series. We can define HP filter as a two-sided linear filter that computes the smoothed series s of y by minimizing the variance of y around s, subject to a penalty that constrains the second difference of s. That is, the HP filter chooses s to minimize,

$$\sum_{t=1}^{T}(y_t - s_t)^2 + \lambda\sum_{t=2}^{T-1}((s_{t+1} + s_t) - (s_t - s_{t-1}))^2 \qquad (1)$$

where T is the sample size and λ is a parameter that penalizes the variability of trend. Thus the penalty parameter λ would control the smoothness of the series. The larger the λ, the smoother is the trend path of the series. If λ = 0, an extreme real business cycle model is taken into consideration where all of the fluctuations in real output are caused by technology shocks, and in this case the HP trend would be the same as the historical time series itself (Metin-Özcan et al. 2001: 217-253). As λ = ∞, *s* approaches a linear deterministic trend. Following Canova (1998: 484) the optimal value of λ is $\lambda = \sigma_x^2/\sigma_c^2$ where σ_x and σ_c are the standard deviation of the innovations in the trend and in the cycle, respectively. Hodrick and Prescott (1997: 4) assume that a 5 percent cyclical component is moderately large, as is a one-eighth of 1 percent change in the growth rate in a quarter, which lead us to select $\sqrt{\lambda} = 5/(1/8) = 40$ or λ = 1600 as a value for the smoothing parameter. Thus we set λ = 1600 in our paper, as well.

When examining the cyclical characteristics of real income and inflation based on HP-filtered data, we report the highest degree of comovement of each variable with real output in bold if the correlation coefficient is significant such as in Alper (2002: 22-54). If the cross correlation

$\rho(j)$, $j\varepsilon\{0,\pm1, \pm2, \dots\}$, between Y_t and X_{t+j} up to four quarters reaches the maximum for a negative j, the series leads the reference cycle, i.e. reaches its turning points j units of time earlier than the GDP. In the other case, if the cross-correlation is maximum for a positive j, the series' cycle lags behind the GDP cycle by j units of time. If the cross correlation between Y_t and X_{t+j} is maximum for $j = 0$, the cycle of X is synchronous. Also if contemporaneous correlation coefficient $\rho(0)$ is positive, zero, or negative, the series X would be considered as procyclical, acyclical, or countercyclical, respectively (Kydland and Prescott, 1990: 10; Fiorito and Kollintzas, 1994: 240). In our sample of 75 observations of the period 1988:Q1-2006:Q3 with quarterly data, the unknown population contemporaneous correlation coefficient is taken to be significant when $0.23<|\rho(t)|<1.00$ leading us to reject at the 5% level of significance the null hypothesis that the population correlation coefficient is zero in a two sided test for bivariate normal random variables.[3]

We now try to extract the cross correlations between HP-detrended cyclical component of real output and annualized domestic inflation in Table 1 below, in which we report both the ratio of standard deviation of inflation with that of the real output (σ/σ_{GDP}) and comovement with real output as correlation of the inflation series with real output in natural logarithms. Our estimation results indicate that deflator-based domestic inflation has a countercyclical characteristic with real output supporting what the supply-driven business cycle models bring out and also that the GDP-deflator based inflation lags the cycle by one quarter.[4] The HP-trended stationary component of inflation is about four times much more volatile than that of real income. Bivariate Granger causality tests not reported here to save space have been given unidirectional causality from cyclical stationary real output to the domestic inflation supporting the cross correlation results in the sense that the stationary real output component precedes the cyclical component of domestic inflation.[5]

[3] As Agénor et al. (1999) emphasize, estimation results obtained in such a way are based on unconditional correlations between filtered real output and inflation, and these correlations do not necessarily imply causal relationships and thus may require at least bivariate exogeneity tests. Nevetheless, these results will provide *a priori* knowledge for the cyclical characteristics of the business cycles.

[4] As Fiorito and Kollintzas (1994: 251-253) emphasize, a benchmark real business cycle (RBC) model can easily account for a negative correlation between output and prices, as technology shocks shift the aggregate supply of output upward.

[5] These estimation results are available from the authors upon request.

Table 1: Cross Correlations Between Detrended Domestic Real Income and Annualized Inflation

obs	σ/σ_{GDP}	X_{t-4}	X_{t-3}	X_{t-2}	X_{t-1}	X_t	X_{t+1}	X_{t+2}	X_{t+3}	X_{t+4}
75	4.37	-0.06	-0.03	-0.03	-0.17	-0.41	**-0.43**	-0.35	-0.18	0.22

Thus, as Chadha and Prasad (1994: 240) express, even though it is widely perceived that temporary movements in output are associated with shocks to demand while longer-term movements are associated with movements in supply, the countercyclical variation of prices suggests that even temporary movements in output may be due to supply shocks.

III. Svar Model of The Turkish Business Cycles

III.1. Model

We now construct an economic model through *a priori* restrictions by way of using contemporaneous economics theory. Considering the Turkish economy conditions, we try to separate real or supply side shocks and nominal or demand side shocks from each other.[6] Let us assume a small open economy highly subject to the effects of capital flows rendering the real exchange rate to be one of the main determinants of the domestic business cycles, and consider the short-term net portfolio investments as exogenously given to the domestic economy and transmitting their effects through the changes in the course of the real exchange rate. These all effects in turn would lead to the supply-side shocks affecting also the course of general economic activity. If we indicate the real effective exchange rate as ER and the relevant structural innovation as φ_{ER} and follow a similar notation to the Ahmed and Park (1994: 1-36), we can write down the process identifying the course of the real exchange rate as,

[6] Of special emphasis here is upon the decomposition of real and nominal factors in our model specification. By assuming in such a way, we think of that for the Turkish economy conditions real shocks represented below by the shocks on real output and real effective exchange rate can easily be coincided with supply shocks. But that the nominal shocks represented by the shocks on inflation and the Treasury interest rate can be considered as demand shocks are subject to be questionable, and in this line that would in our opinion be more appropriate to assume shocks on expenditure-sided proxy variables as the demand shocks. Thus, for instance, shocks on the Treasury interest rate can be assumed a demand side innovation resulted from fiscal shocks. Thus, we here prefer to use the phrase 'real and nominal shocks' to identify our economic model.

$$ER = C_{11}(L)\ \varphi_{tER} \qquad (2)$$

where $C_{11}(L)$ is a finite-order polynomial in the lag operator and φ_{tER} is a white noise independent and identically distributed (i.i.d.) disturbance term.[7] In Equation (2), real exchange rate is assumed as a function of innovations upon itself led mainly by the course of exogenous portfolio flows, and we expect that C_{11}>0, that is, real exchange rate would be appreciated through the positive shocks upon itself.

We assume that in the long run there exists no effect of the innovations on the nominal shocks upon the real variables. So real output is mainly affected by the real exchange rate, while the couse of the latter is assumed to be driven by the exogenously given portfolio flows as expressed in Equation (2). If we indicate the real domestic output as Y and the relevant structural innovation as φ_Y,

$$Y = C_{21}(L)\ \varphi_{tER} + C_{22}(L)\ \varphi_{tY} \qquad (3)$$

Many empirical papers upon the Turkish economy, such as Kirmanoğlu and Özçiçek (1999: 27-34), Berument and Paşaoğulları (2003: 401-435), Berument and Dinçer (2004: 20-32) and Saatçioğlu and Korap (2006) reveal either contractionary impact of real exchange rate depreciations on output or that appreciations of the real exchange rate through the capital flows would stimulate the domestic real income growth process and lower the domestic inflationary pressures and the interest rates. Following these papers, we expect in Equation (3) that C_{21}>0. Saatçioğlu and Korap (2006) attribute this to the relieving effect on the cost pressure settled in the domestic economy resulted from real appreciations. The effect of positive domestic supply shock on real output is also assumed in a positive way and is expected to persist over time as was emphasized in Ahmed and Park (1994: 1-36) such that C_{22}>0.

As a third identifying assumption, we assume that domestic inflation would be contingent upon the structural innovations on real output and real exchange rate as well as innovations on

[7] Cij means that the response of the i-th variable to the j-th structural shock.

itself reflecting the price inertia phenomenon.[8] We expect that aggregate supply shocks would lower the domestic inflation structure supporting a Real Besiness Cycle (RBC) perspective and that a similar effect would occur through the real appreciations as expressed above. If we indicate the domestic inflation as P and the relevant structural innovation as μ_P,

$$P = C_{31}(L)\ \varphi_{tER} + C_{32}(L)\ \varphi_{tY} + C_{33}(L)\ \mu_{tP} \qquad (4)$$

Thus, $C_{31}<0$, $C_{32}<0$ and $C_{33}>0$. The last identifying assumption in our SVAR model is constructed upon the domestic interest structure, in which domestic interest rate is considered a function of all three structural shocks assumed so far and of the structural shock upon itself. If we indicate domestic interest rate, represented by the short-term Treasury bond rates, as R and the relevant structural innovation as δ_R,

$$R = C_{41}(L)\ \varphi_{tER} + C_{42}(L)\ \varphi_{tY} + C_{43}(L)\ \mu_{tP} + C_{44}(L)\ \delta_{tR} \qquad (5)$$

In Equation (5), we expect that $C_{42}>0$ if excess aggregate demand and high domestic absorption levels lead to inflationary pressures resulted in an increase in domestic interest structure through the *so-called* Fisher interest parity effect, which states in the long run the nominal interest rate moves one for one with inflation and thus one for one with nominal money growth, leaving the real interest rate unchanged (Blanchard, 1997: 383). But if supply shocks rather than the demand shocks are the main determinant of the domestic business cycles, we can assume that $C_{42}<0$ due to the relieving effect on domestic inflation structure, as expressed above, resulted from positive supply shocks. Following this specification issues, if the business cycle is supply-driven, $C_{41}<0$ since such an effect can easily be attributed to the course of nominal interest rates occured downwards because of the relevant effect on domestic borrowing possibilities resulted from capital inflows which leads to the real appreciation of the domestic currency, by increasing the price of domestic borrowing assets thus pulling down the nominal interest rates. We also expect that $C_{43}>0$ due to the Fisher effect expressed above.

[8] We here omit any relationship between domestic inflation and money supply measures considered by the Quantity Theoretical and Monetarist explanations of the Turkish inflation in line with the empirical findings and policy suggestions of the CBRT (2002) and Saatçioğlu and Korap (2006) to ease the identification of our structural VAR system.

III.2. Preliminary Data Issues

We now move to the multivariate analysis of the Turkish business cycles in an empirical way, employing a variety of econometric procedures available in the program EViews 5.1. From a point of view identifying the long run effects of structural shocks on output fluctuations through *a priori* restrictions on economic theory using SVAR methodology of Blanchard and Quah (1989: 655-673), we try to construct a SVAR model of the Turkish economy for the investigation period of 1992Q1-2006Q3 with quarterly observations.[9] For this purpose, we assume a VAR system consisted of four endogenous and one exogenous variables, that is, real gross domestic product (Y), real effective exchange rate (ER) based on producer price index published by the CBRT using the IMF weights for 17 countries[10], annualized inflation (P) and the Treasury interest rate (R), which is the maximum rate of interest on the Treasury bills whose maturity are at most twelve months or less. We also consider the sum of portfolio investments net of assets and liabilities as equity securities and debt securities in millions of US$s (PORTNET) as an exogenous variable in our multivariate system specification.[11]

The aggregates representing domestic income can normally be expected to indicate seasonality, thus for estimation purposes they are used in a de-seasonalized form. We use U.S. Census Bureau's X12 seasonal adjustment program also available within EViews 5.1 to adjust real income variable against seasonality. We consider real income and real effective exchange rate data in natural logarithms, while the latter variables in our VAR system, that is, INFLATION and BONOFAIZ, are used in the linear form, not in natural logarithms. All the data are taken from the electronic data delivery system of the Central Bank of Republic of Turkey (CBRT). Since the availability of quarterly capital flows data is possible as of the

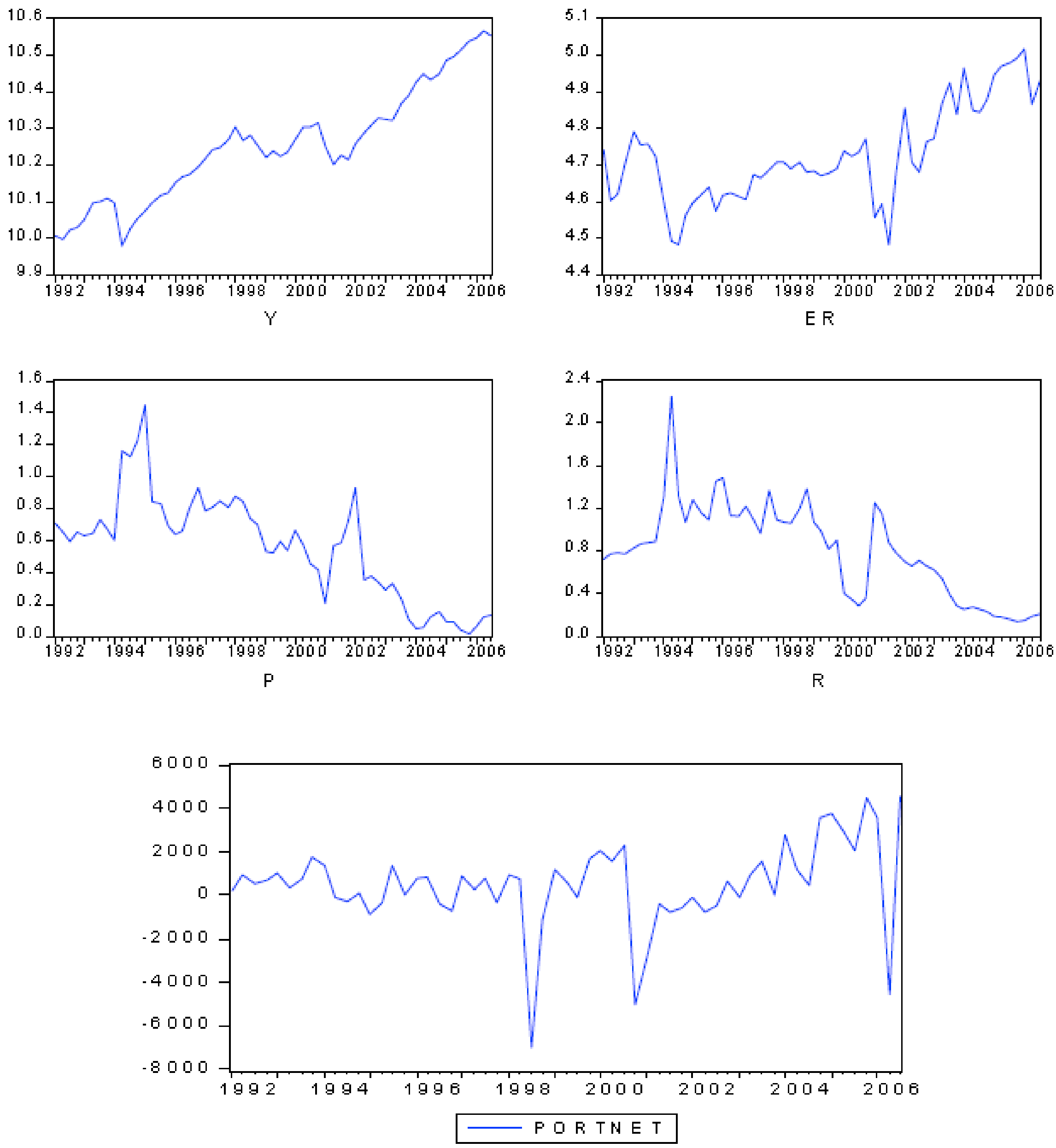
[9] Ahmed et al. (1993: 335-359), Ahmed and Park (1994: 1-36), Hoffmaister and Roldós (1997), Lopez et al. (1997), McCoy (1997), Buckle et al. (2002) and Meuers (2006) can be considered some empirical papers employing structural identification of the business cycles for various country cases. Besides, Çulha (2006) investigating the determinants of capital flows and Çatık (2006) revealing the contractionary impact of devaluations leading also to high inflation are some recent papers employing SVAR approach upon the Turkish economy.

[10] Namely Germany, USA, Italy, France, United Kingdom, Japan, Netherlands, Belgium, Switzerland, Austria, Spain, Canada, Korea, Sweden, Iran, Brazil and Greece. An increase in this index would denote a real appreciation of domestic currency, whereas a decrease would denote a real depreciation.

[11] We used the portfolio flows data as an exogenous variable for the ease of identification issues expressed below while taking the real effective exchange rate data as an endogenous variable in our VAR framework, since we *a priori* assume that short-term portfolio flows depend mostly on exogenous expectations of economic agents determined out of our system specification but in turn affect the real exchange rate data to a large extent and by this way transmit their effects onto the Turkish economy so that makes the real exchange rate an endogenous variable.

beginning of 1992 through using this source, our estimation sample begins as of the beginning of 1992, as well. The time series representation of the variables can be seen in Figure 1 below,

Figure 1: Time Series Used In The Paper



All the endogenous variables in Figure 1 seem to be non-stationary which drift together in the sample period. Even for the real exchange rate in the free float period of post-2001, there seems to be a trend drifting the real exchange rate upward. For this purpose, we also apply to the augmented Dickey- Fuller (ADF) and Phillips-Perron (PP) unit root tests assuming both constant and constant&trend terms in Table 2 below to confirm what we see in Figure 1,[12]

Table 2: Unit Root Tests

	ADF test (with only constant)		ADF test (with constant&trend)	
	(levels)	(first differences)	(levels)	(first differences)
Variable				
Y	-0.40 (0)	-7.13 (0)*	-2.07(0)	-7.08 (0)*
ER	-1.97 (0)	-8.75 (0)*	-3.31 (0)	-9.17 (0)*
P	-0.51 (4)	-6.29 (3)*	-2.26 (4)	-6.39 (3)*
R	-2.16 (0)	-7.74 (1)*	-3.39 (2)	-11.37 (11)*
	PP test (with only constant)		PP test (with constant&trend)	
	(levels)	(first differences)	(levels)	(first differences)
Variable				
Y	-0.43 (1)	-7.13(1)*	-2.35 (3)	-7.08 (1)*
ER	-1.86 (2)	-10.29 (7)*	-3.31 (0)	-11.65 (10)*
P	-1.56 (4)	-9.27 (1)*	-3.42 (3)	-9.22 (0)*
R	-2.16 (0)	-10.17 (9)*	-3.39 (0)	-7.77 (1)*

Critical Values	with only constant	with constant&trend
% 1 level	-3.55	-4.12
% 5 level	-2.91	-3.49

When we examine the results of the unit root tests, we see that the null hypothesis that there is a unit root cannot be rejected for all the variables in the level form supporting our cursory examination of Figure 1 above. From now on, therefore, we will carry on our empirical research by using the stationary form data.

[12] For the case of stationarity, we expect that these statistics are larger than the MacKinnon critical values in absolute value and that they have a minus sign. The numbers in parantheses are the lags used for the ADF stationary test and augmented up to a maximum of 10 lags due to using quarterly observations, and we add a number of lags sufficient to remove serial correlation in the residuals, while the Newey-West bandwidths are used for the PP test. '*'and '**' indicate the rejection of a unit root for the %1 and %5 levels, respectively.

For the lag length of our unrestricted VAR model of which the maximum lag number selected is 5 due to using quarterly frequency data considering five lag order selection criterions, that is, sequential modified LR statistics employing Sims' (1980) small sample modification, final predicton error criterion (FPE), Akaike information criterion (AIC), Schwarz information criterion (SC) and Hannan-Quinn information criterion (HQ), LR, FPE and AIC statistics suggest 4, while SC and HQ statistics suggest 0 lag orders. We choose thus the lag order 4 selected by sequential modified LR and minimized AIC statistics for our dynamic VAR specification in order to check our econometric model.

III.3. Identification and Estimation

Following Hoffmaister and Roldós (1997), a similar methodology is applied to the Turkish data using SVAR analysis of Blanchard and Quah (1989: 655-673) so as to recover the structural innovations in line with the economic model construction in section III.1. Since the unit root tests performed above reveal that the null hypothesis of a unit root cannot be rejected against the alternative hypothesis of stationarity around a constant or a constant and deterministic trend, we consider all four endogenous variables in first differenced-stationary form. But at the end of the paper, we also perform some cointegration tests to reveal any stochastic common trend between the variables in the level form.

By making use of McCoy (1997), Buckle et al. (2002) and QMS (2004: 717-723), let X_t be a vector of the endogenous variables. Ignoring the constant term, assume first the structural form equation below,

$$B(L)X_t = u_t \tag{6}$$

where B(L) is the p^{th} degree matrix polynomial in the lag operator L, where p is the number of lagged periods used in the model, such that $B(L) = B_0 - B_1L - B_2L^2 - \ldots - B_PL^P$. B_0 is a nonsingular matrix normalised to have ones on the diagonal and summarizes the contemporaneous relationships between the variables in the vector X_t. The variance of u_t, Λ, is a diagonal matrix where diagonal elements are the variances of structural disturbances, therefore the structural disturbances are assumed to be mutually uncorrelated. The reduced form VAR with this structural model is,

$$A(L)y_t = \varepsilon_t \tag{7}$$

where $A(L)$ is a matrix polynomial in the lag operator L, ε_t a vector of serially uncorrelated reduced form disturbances, and $var(\varepsilon_t) = \Sigma$. The relationship between Equations (6) and (7) is,

$$A(L) = B_0^{-1}B(L) = I - A_1L - A_2L^2 - \dots - A_PL^P \tag{8}$$

and

$$\varepsilon_t = B_0^{-1}u_t \tag{9}$$

The parameters in the structural form equation and those in the reduced form equation are related by,

$$A(L) = I - B_0^{-1}[B_1L - B_2L^2 - \dots - B_PL^P] \tag{10}$$

and

$$\Sigma = B_0^{-1}\Lambda B_0^{-1\prime} \tag{11}$$

The restrictions with the long-run pattern matrix for the reduced form SVAR model leaving the short run dynamics unconstraint from Equation (2) to Equation (5) can be seen as follows,

Table 3: SVAR Long-Run Response Pattern

C_{11}	0	0	0
C_{21}	C_{22}	0	0
C_{31}	C_{32}	C_{33}	0
C_{41}	C_{42}	C_{43}	C_{44}

Below we plot the accumulated impulse response functions of the endogenous variables using 1000 Monte Carlo repititions to one standard deviation (S.D.) innovations

Figure 2: Accumulated Responses Of D(ER)

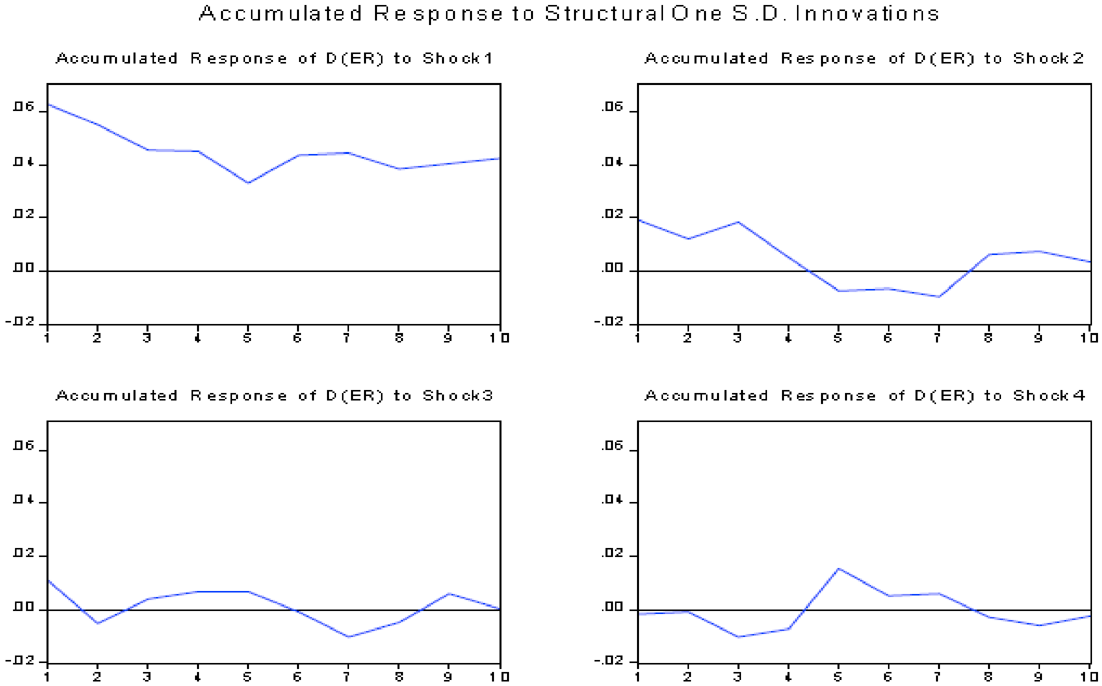


Figure 3: Accumulated Responses Of D(Y)

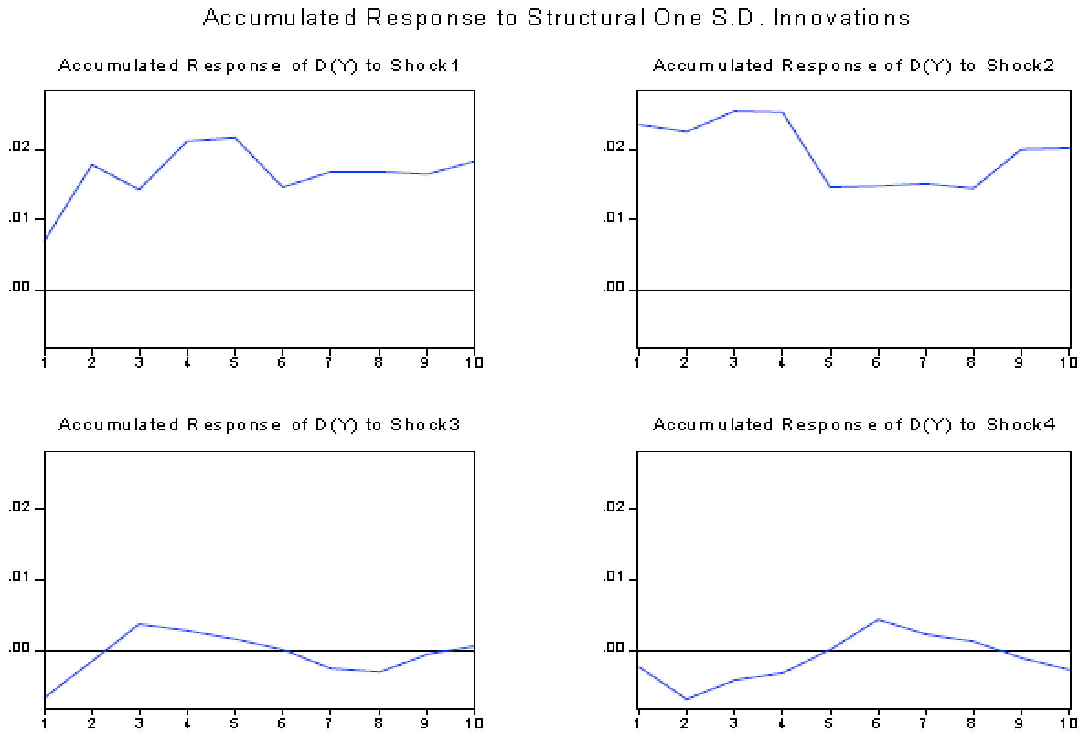


Figure 4: Accumulated Responses Of D(P)

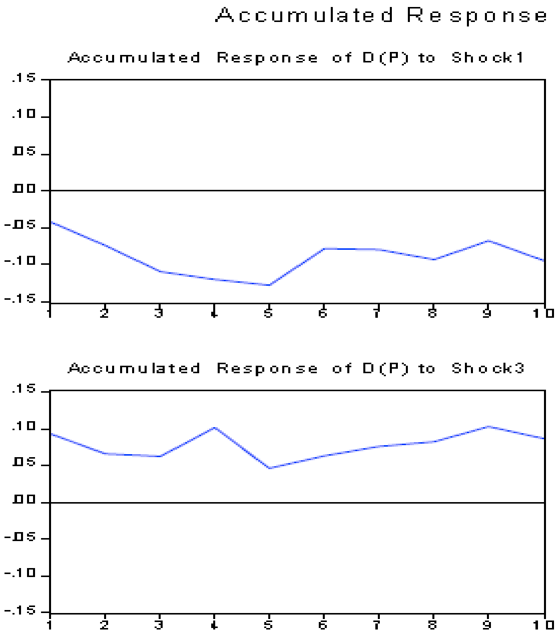
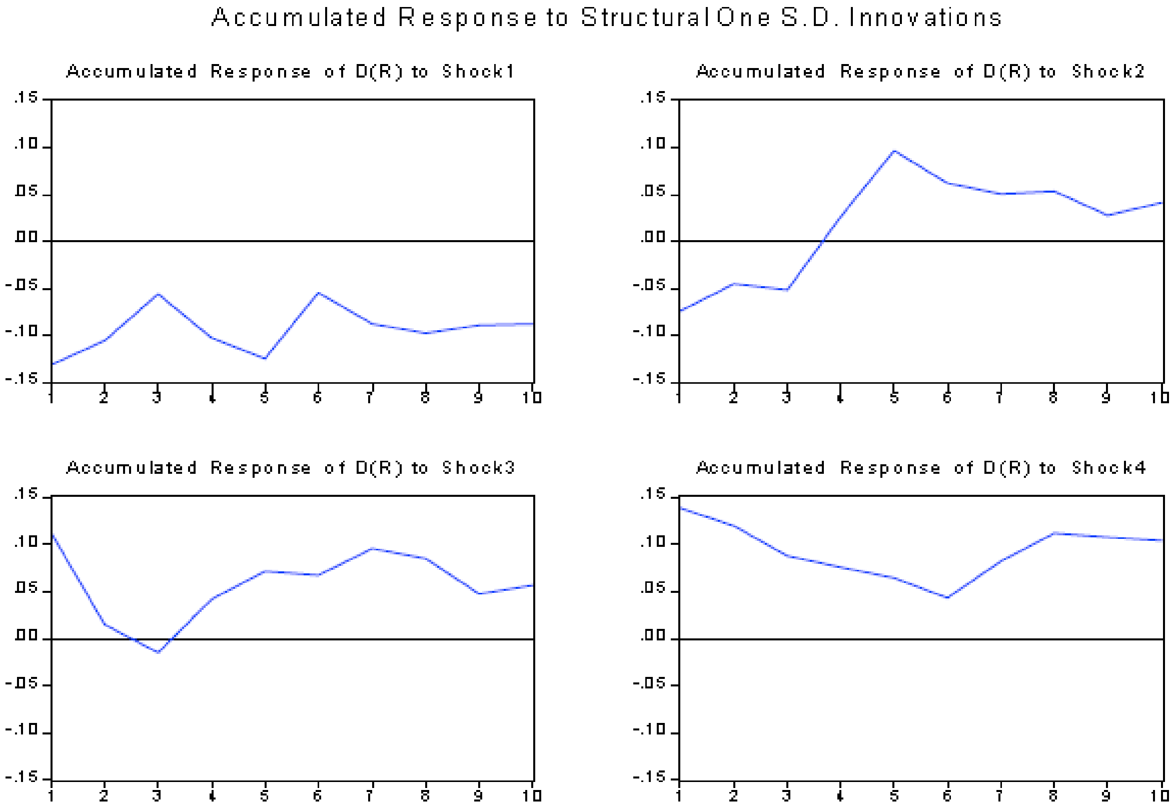


Figure 5: Accumulated Responses of D(R)



When examining Figure 2 can be seen that, as hypothesized above, the course of the real exchange rate is mainly responsive to the own shocks. Besides, some positive effects of real income innovations can be noticed, while the nominal shocks have no significant effect on the real exchange rate. In Figure 3, the impulse response of the real income is brought out. The expected persistence effect is estimated so that positive shocks on the real income growth process lead to higher growth rates for the subsequent periods. Appreciating real exchange rate can affect the real income growth process mainly in two alternative ways in the sense that positive innovations on the real exchange rate would lead to either a depreciating effect on the real income growth process, due to the possible deterioration in the international competitiveness of the domestic goods leading also to the trade balance deterioration, or an improvement on the production capacities, due to the relieving effect on the cost pressure settled in the domestic economy. Our findings reveal that real exchange rate appreciations significantly improve the domestic income growth process. A structural one S.D. positive shock on real exchange rate increases real income about %2 in a cumulative way inside the whole period. Although some negative effects occur on the variable D(Y) in response to the positive shocks on domestic inflation and interest structure, which can be attributed to that the larger the cost-pressure reflected to the price structure of the economy the lower the real income growth path, such a result seems not to have a cumulative persistence to the structural innovations. Thus in line with the estimation results here, we think of that this interpretation is open to be questionable at least through the estimation results obtained so far in this paper.

In Figure 4 and Figure 5, we obtain some supportive findings to such conclusions. The real appreciations lower to a large extent the domestic inflationary pressures. A structural one S.D. positive shock on real exchange rate lowers the domestic inflation in a cumulative way within a range of %5 and %10. Although this may be hypthesized as a stylized fact of the Turkish economy, cumulative responses of inflation to structural innovations on real income cast some doubt on this subject. We see that although responses of inflation to structural positive innovations on real income are about zero for the first five periods following the shock, the subsequent periods indicate a positive effect on inflation. We can assume that the larger the time period the larger the upward pressure on inflation of the real income growth. The price inertia phenomenon is also verified by our findings.

Finally we give support to an RBC-based policy issue that real appreciations would lower the domestic interest structure led by the public sector borrowing requirement. Positive innovations on inflation would increase the domestic interest structure as was hypothesized in Equation 5 above. Having examined impulse responses, the variance decomposition analysis using structural factorization is conducted below,

Table 4: Variance Decomposiition Of The Variables

Percentage of Real Exchange Rate Variance due to

Variance Period	S.E.	φ_{tER}	$\square_{tY}$	μ_{tP}	δ_{tR}
1	0.066277	88.96468	8.252908	2.733226	0.049188
4	0.072580	77.11365	11.90493	9.082128	1.899299
8	0.082644	63.68588	15.13845	9.511550	11.66412
12	0.084611	60.94921	15.25729	11.69523	12.09827
20	0.086244	59.68512	15.43601	12.28465	12.59422

Percentage of Real Domestic Income Variance due to

Variance Period	S.E.	φ_{tER}	$\square_{tY}$	μ_{tP}	δ_{tR}
1	0.025440	7.687580	84.75903	6.749880	0.803509
4	0.030226	24.71331	60.94312	10.75106	3.592513
8	0.033514	24.91036	59.43806	9.749454	5.902126
12	0.034361	24.21527	59.28681	9.971676	6.526240
20	0.034810	23.75198	58.69845	10.40492	7.144651

Percentage of Domestic Inflation Variance due to

Variance Period	S.E.	φ_{tER}	$\square_{tY}$	μ_{tP}	δ_{tR}
1	0.115422	13.31877	1.599837	65.84074	19.24066
4	0.143159	20.47498	3.364793	53.83978	22.32045
8	0.179308	21.45127	10.69683	45.76370	21.49275
12	0.194443	22.25820	13.33605	40.67643	23.72932
20	0.203201	21.95362	14.99981	39.16438	23.88219

Percentage of Domestic Interest Structure Variance due to

Variance Period	S.E.	φ_{tER}	$\square_{tY}$	μ_{tP}	δ_{tR}
1	0.234018	31.62172	10.31832	22.85862	35.20134
4	0.286826	27.63381	15.44270	31.63033	25.29316
8	0.316256	29.26318	19.00537	27.86718	23.86428
12	0.322460	28.37328	20.04037	28.36095	23.22540
20	0.325236	28.11969	20.15370	28.30495	23.42166

Examining Table 4 reveals that over a period of 20 quarters, nearly %60 of the forecast error variance of the real exchange rate can be accounted by the shocks over itself, while similar results are estimated for the real income growth process. Real exchange rate also explains one fourth of the forecast error variance on the variable Y. Nominal factors cannot be found to have significant effects on the forecast error variances of the real variables supporting an RBC perspective of the business cycles. On the other side, over a period of 20 quarters, nominal variables seem to have an endogenous characteristic for the period under investigation. Inflation explains %40 of the forecast error variance over itself, while one-fourth of the forecast error variance of inflation can be explained by domestic interest rate and one-fifth by real exchange rate. More interesting is that all the factors considered significantly affect the forecast error variance of interest rate, and about one-fourth or one-fifth of the error variance of the Treasury interest rate is explained by all these factors considered supporting the endogenous characteristic of interest rate, and such a conclusion may impose an accomodative role to the discretionary policy instruments used by the policy makers.

Our estimation results are based on the differenced stationary data. We finally estimate some cointegration tests of the Johansen-Juselius type using the data in the level form, and find at the %5 level one cointegrating vector lying in the variable space assuming both intercept and intercept&trend factors restricted in the cointegrating vector. We give below the normalized vector on the real income with standard errors in parentheses in Equation (12). Thus the variables seem to be driven by one comman trend, and following Ahmed and Park (1994: 1-36) our empirical model above may be somewhat overdifferenced. But we find that estimation results using I(1) variables yield quite similar results to those estimated above, and all these findings not reported here are available upon request.

$$Y = \underset{(0.44)}{1.22}*ER - \underset{(0.28)}{1.81}*P + \underset{(0.18)}{0.55}*R + 4.50 \qquad (12)$$

IV. Concluding Remarks

In our paper, we try to reveal some properties of the Turkish business cycles. At first, considering a business cycle perspective we find that domestic inflation has a countercyclical characteristic with real output and that inflation lags the cycle by one quarter. We then employ a structural VAR (SVAR) model upon the Turkish economy. Our *ex-post* findings reveal that the courses of real variables are mainly determined by the supply shocks, while both real and nominal shocks affect significantly the dynamics of the nominal variables.

In line with our empirical model construction, papers using larger models which account for the Turkish business cycles and also investigating the course of the Turkish trade balance identified by structural economic relationships, not considered in this paper, should be elaborately dealt with in future papers.

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