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Nonlinearity and structural stability in the Phillips curve: Evidence from Turkey

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ABSTRACT

In this paper, we investigate possible nonlinearities in the inflation–output relationship in Turkey for the 1980–2008 period. We first estimate a linear bivariate model for the inflation rate and output gap, and test for linearity of the estimated model against nonlinear alternatives. Linearity test results suggest that the relationship between the inflation rate and output gap is highly nonlinear. We estimate a bivariate time-varying smooth transition regression model, and compute dynamic effects of one variable on the other by generalized impulse response functions. Computed impulse response functions indicate that inflation–output relationship in Turkey during the analyzed period was regime dependent and varied considerably across time.

1. Introduction

The relationship between real economic activity and inflation rate has been one of the most investigated yet controversial issues in macroeconomics both on theoretical and empirical grounds. Perhaps, the most popular model that describes the relationship between economic activity and inflation rate is the "Phillips curve" first introduced by Phillips (1958) and popularized by Samuelson and Solow (1960). According to the Phillips curve, inflation rate is negatively correlated with unemployment rate. Although the original Phillips curve has been severely criticized on theoretical grounds and a wide variety of Phillips curve has been developed, some variant of the Phillips curve is still being used to estimate the relationship between some measure of real economic activity and inflation rate as well as to forecast future inflation rate (e.g. Blinder, 1997; Mehra, 2004; Sbordone, 2006; Canova, 2007). Besides its theoretical importance, a reliable and robust estimate of Phillips curve has a clear implication for policy authorities. Because Phillips curve foresees a trade-off between inflation rate and unemployment rate, it might be important for policy authorities to estimate the so-called "sacrifice ratio", i.e., how much output will be lost as a result of reducing inflation rate by 1%. Therefore, the relationship between inflation rate and real economic activity has been extensively investigated in the empirical literature (Lucas and Rapping, 1969; Ball, 1993; Romer, 1993; Fischer, 1996; Temple, 2002; Daniels et al. 2005).

Even though Phillips (1958) noted that the relationship between nominal wages and unemployment rate is nonlinear, subsequent studies employing the Phillips curve assumed that the relationship between inflation rate and unemployment rate is linear due to relative simplicity of handling linear models (Gordon, 1970, 1977; Barro and Gordon, 1983). However, it has long been recognized that the dynamics and interrelationships of many economic variables might be inherently nonlinear. For example, Keynes (1936, p.314) noted that economic downturns happen to be sudden and violent whereas there is, as a rule, no such sharp turning points for economic recoveries. As pointed out by Morgan (1993), nonlinearity is a common feature of many widely accepted economic models, including standard Keynesian model with 'Keynesian' and 'Classical' regions of the aggregate supply curve, the liquidity trap theory, credit constraint models, and menu cost models. In addition, capacity constraints as well as wage and price stickiness may also cause to nonlinearity in the inflation–output relationship (see also Jackman and Sutton, 1982; Tsiddon, 1993; Ball and Mankiw, 1994; Clark et al., 2001).

Another disputable issue in the literature is the shape of the Phillips curve. For example, Debelle and Laxton (1997) reported results in favor of a convex Phillips curve, while Stiglitz (1997) suggested that the Phillips curve is concave in the case of USA. The results of Cover (1992), De Long and Summers (1988), Karras (1996) and Telatar and Hasanov (2006) indicate that positive demand shocks are more inflationary than negative shocks are disinflationary in various countries, and thus provide some evidence in favor of a convex Phillips curve. Nobay and Peel (2000) use convex Phillips curve model to examine optimal monetary policy under discretion.

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Schaling (2004) also uses a convex Phillips curve to analyze monetary policy rules. Laxton et al. (1999) argued that traditional econometric methods have a little power to show the convexity of the Phillips curve and suggested that policy authorities must take account of nonlinearities in making their decisions. On the other hand, Filardo (1998), based on the empirical evidence for the USA, argued that the Phillips curve is not purely convex or concave, but instead, it is convex when the output gap is positive and concave when the gap is negative.

A related issue about the nonlinearity of the Phillips curve is possible structural breaks in the output–inflation relationship. In his seminal paper, Lucas (1976) argued that the parameters of the econometric models are not policy invariant. Therefore, the slope of the Phillips curve shall change whenever the macroeconomic policy has changed. Alogoskoufis and Smith (1991) found a one-time structural break in the UK and US Phillips curves from 1967 to 1968, which was attributed to demise of the Bretton Woods system. Bai and Perron (2003), on the other hand, provided strong evidence on two breaks in the Phillips curve, one in 1967, and another in 1975, providing evidence in support of the Lucas critique. Khalaf and Kichian (2005) found some evidence in favor of both linear and nonlinear breaks in the Phillips curve in the case of Canada.

The shape of the Phillips curve is also important for policy design. Linear Phillips curve assumes that the slope of the curve is constant everywhere and hence the sacrifice ratio, i.e., output cost of disinflation is constant irrespective of the level of economic activity as well as the speed of disinflation. On the other hand, a nonlinear Phillips curve suggests that the sacrifice ratio is dependent on both speed of disinflation and current level of economic activity. If Phillips curve is convex, then output cost of disinflation will be lower (higher) when the initial level of inflation is high (low). However, if the Phillips curve is concave, then disinflation might be costly for higher levels of inflation. Therefore, from a policy perspective, it is crucial to know whether the Phillips curve is linear or nonlinear, and whether it is concave or convex, if found to be nonlinear.

In this paper, we examine possible nonlinearities and structural changes in the inflation–output relationship in the case of Turkey. Starting from early 1980s Turkey has implemented a wide range of economic reforms. Turkey has abandoned import-substitution policies, liberalized foreign trade regime and financial markets, and allowed real exchange rate to depreciate in order to stimulate exports. In late 1980s Turkey started to let exchange rates to appreciate in real terms to fight inflation and attract huge capital inflows to cover domestic savings gap and finance economic growth. However, in 1994 Turkey underwent a severe economic crisis. Although Turkey has implemented tight fiscal and monetary policies to mitigate the effects of the crisis, it switched to expansionary policies soon after recovery from the crisis. Starting from December 1999 Turkey implemented exchange rate-based stabilization program under support of the IMF. However, after the economic crisis in February 2001, it abandoned exchange rate-based stabilization program, implemented massive structural reforms, including granting independence to central bank, and adopted inflation targeting regime. Therefore, taking account of the fact that it has implemented quite different macroeconomic policies during the last three decades, Turkey provides a good and natural example to examine possible structural changes and nonlinearities in the inflation–output relationship.

Several authors have examined the inflation–output relationship in the case of Turkey. Önder (2004) reports that forecasts of inflation based on Phillips curve relationship outperform forecasts from other models. Yazgan and Yilmazkuday (2005) provide some evidence in favor of forward-looking New Keynesian Phillips curve, and conclude that back-ward looking behavior is rejected in the case of Turkey. On the other hand, Kuştepeli (2005) fails to find any evidence of Phillips curve for any specification. Çatik et al. (2008) argue that inclusion of a measure of distribution of relative prices is crucial for discovering the Phillips curve trade-off. Önder (2009) uses structural change models and Markov regime-switching models to explore possible instabilities and nonlinearities in the Phillips curve. She finds some evidence of both instability and nonlinearity but fails to find any evidence for asymmetric response of inflation to output gap. Within a related framework, Telatar and Hasanov (2006) examine possible nonlinearities and asymmetries in the effects of monetary shocks on inflation rate and output. They concluded that positive and negative monetary shocks affect both inflation and output nonlinearly, providing some evidence in favor of convex Phillips curve.

Our approach in this paper is different from previous researches. Önder (2009) examined possible instabilities and nonlinearities in the Phillips curve separately. However, our methodology in this paper allows examining both features simultaneously. Second, the methodology adopted in Önder (2009) restricts both structural breaks and nonlinearities to be instantaneous. But we use smooth transition models to examine these features, which are more appealing than abrupt change models. Telatar and Hasanov (2006) also used smooth transition methodology to examine possible asymmetries in the effects of monetary shocks. Third, in order to account for possible endogeneity problem, we use vector autoregressive (VAR) model to study inflation–output relationship. Only Önder (2004) and Telatar and Hasanov (2006) considered multivariate models. Using quarterly observations on industrial production index we first generate series of output gap using the Hodrick and Prescott (1997) filter. Then, we estimate a linear VAR model between output gap and inflation rate, proxied by the first difference of logarithms of consumer price index. Using the linearity tests of Lundbergh et al. (2003) we find that the relationship between output gap and inflation rate in the case of Turkey can be best prescribed by a time-varying smooth transition regression model. Then we use generalized impulse response functions proposed by Koop et al. (1996) to investigate the dynamic effects of one variable on the other. Our results suggest that the dynamic relationship between inflation rate and output gap has changed significantly across time that can be attributed to macroeconomic environment, thus providing support for the Lucas critique in the case of Turkey.

The remaining of the paper is organized as follows. In the next section we outline and discuss econometric methodology employed in this paper to examine the instability and nonlinearity of the inflation–output relationship. In Section 3 we give and discuss empirical results, and Section 4 concludes.

2. Econometric methodology: specification, testing, and estimation of time-varying smooth transition regression models

Nonlinearities and structural change in dynamics and interrelationships of economic variables have attracted a great interest of researchers in recent years. While some authors have examined only nonlinearities (e.g., Teräsvirta and Anderson, 1992; Granger and Teräsvirta, 1993; Weise, 1999; Rothman et al., 2001; Arghyrou et al., 2005), others studied structural changes in economic variables (for example, Perron, 1989; Lin and Teräsvirta, 1994; Stock and Watson, 1996; Bai and Perron, 2003). Lundbergh et al. (2003), on the other hand, argued that time series might be described more appropriately by simultaneous structural change and nonlinearities, and developed a time-varying smooth transition autoregressive (TV-STAR) model that can be used to study such features of series. Sollis (2004, 2008), Huang and Chang (2005), Holt and Craig (2006), and Telatar and Hasanov (2009) have provided some evidence that both structural change and nonlinear models might be appropriate for many economic variables.

The TV-STAR model proposed by Lundbergh et al. (2003) is based on the principle of smooth transition, which allows for nonlinear dynamics in conjunction with time-varying parameters. Smooth transition regression (STR) models have several advantages over competing structural break and nonlinear models. First, STR models are theoretically more appealing over simple threshold and Markov regime-switching models, which impose an abrupt change in coefficients. Instantaneous changes in regimes are possible only if all agents act simultaneously and in the same direction. For the market of many traders acting at slightly different times, however, a smooth transition model is more appropriate. In addition, instantaneous change models can be obtained as a special case of the STR models by appropriate parameter restrictions. Second, STR models allow for modeling different types of nonlinear and asymmetric dynamics depending on the type of the transition function (see, for example, Teräsvirta and Anderson, 1992). Third, STR modeling approach allows one to choose both the appropriate switching variable and the type of the transition function unlike other regime-switching models. Finally, the modeling procedure of Lundbergh et al. (2003) can be used to examine both nonlinearity and structural change simultaneously and to distinguish between these two features. Considering these merits of STR models as well as the fact that both structural change and nonlinearities might govern dynamic properties of time series, in this paper we generalize the methodology of Lundbergh et al. (2003) to multivariate setting in order to study possible structural changes and nonlinearities in the interrelationships between inflation rate and output gap in Turkey.

The procedure of Lundbergh et al. (2003) has been used widely in the empirical literature to analyze possible structural breaks and nonlinearities in economic variables. For example, Franses and van Dijk (2005) compare short- and long-horizon (up to 12 periods ahead) forecast performance of different linear and nonlinear models. They find that more sophisticated nonlinear models, such as TV-STR models, outperform linear and other simple nonlinear models only in long horizons. Holt and Craig (2006) examine the US pig–maize cycle (the cyclic behavior of pig and maize prices) and find strong evidence of nonlinearity, regime-dependent behavior, and time-varying parameter change. Sollis (2008) uses TV-STAR models to examine dollar-based exchange rates of 17 developed countries. Although he finds that in some cases there is support for both nonlinearity and structural change, in other cases there appears to be stronger support for structural change than for nonlinearity. Sollis (2008) concludes that structural change is an important feature of the data. Balagtas and Holt (2009) use TV-STR models to investigate the long-run decline in the relative prices of primary commodities, and find that most of the series are characterized by simultaneous structural change and nonlinearity. Musso et al. (2009) analyze structural change and nonlinearity in the Euro-area Phillips curve. They find strong evidence of time variation in the Phillips curve. However, they fail to find significant evidence of nonlinearity, in particular, in relation to the output gap.

Although Musso et al. (2009) also use a TV-STR model to examine both instabilities and nonlinearities in the Phillips curve, our approach here is different from their approach in that we generalize TV-STR modeling approach to multivariate framework in order to deal with possible endogeneity problem.[3] Then we use generalized impulse response functions to investigate the dynamic relationships between output gap and inflation rate. Now, we briefly discuss specification and estimation procedure of bivariate TV-STR models.

2.1. Specification of time-varying smooth transition regression models

The TV-STAR model and specification procedure of Lundbergh et al. (2003) can easily be generalized to a multivariate case following Weise (1999) and Rothman et al. (2001), who generalized STAR model of Teräsvirta (1994) to vector autoregressive (VAR) models. Let π_t and y_t denote the inflation rate and output gap, respectively. Then, a bivariate time-varying smooth transition regression (bivariate TV-STR) model for inflation rate and output gap can be written as follows:

$$\begin{aligned} x_t = {} & \Phi_{1,0} + \sum_{i=1}^{p} \Phi_{1,i}x_{t-i} + \left(\Phi_{2,0} + \sum_{i=1}^{p} \Phi_{2,i}x_{t-i}\right)^* F(s_t;\gamma_1,c_1) \\ & + \left(\Phi_{3,0} + \sum_{i=1}^{p} \Phi_{3,i}x_{t-i}\right)^* G(t;\gamma_2,c_2) \\ & + \left(\Phi_{4,0} + \sum_{i=1}^{p} \Phi_{4,i}x_{t-i}\right)^* F(s_t;\gamma_1,c_1)^* G(t;\gamma_2,c_2) + \varepsilon_t \end{aligned} \tag{2.1}$$

where $\mathbf{x}_t$ is a (2×1) column vector given by $\mathbf{x}_t=(\pi_t,\ y_t)'$, $\mathbf{\Phi}_{j,0}$, $j=1,2,3,4$ are (2×1) vector of constants, $\mathbf{\Phi}_{j,i}$, $j=1,2,3,4$, $i=1,2,\ldots,p$ are (2×2) matrices of parameters, and $\boldsymbol{\varepsilon}_t=(\varepsilon_{\pi t}, \varepsilon_{yt})'$ is a (2×1) vector of residuals with mean zero and (2×2) covariance matrix $\mathbf{\Sigma}$.

Lundbergh et al. (2003) consider following logistic transition functions:

$$F(s_t;\gamma_1,c_1) = \left[1+\exp\{-\gamma_1[s_t-c_1]\}\right]^{-1}, \quad \gamma_1>0, \tag{2.2}$$

and

$$G(t;\gamma_2,c_2) = \left[1+\exp\{-\gamma_2[t-c_2]\}\right]^{-1}, \quad \gamma_2>0. \tag{2.3a}$$

The transition functions $F(s_t;\gamma_1,c_1)$ and $G(t;\gamma_2,c_2)$ are continuous functions that are bounded between 0 and 1. In Eq. (2.2), s_t is the transition variable, and γ_1 and c_1 are slope and location parameters, respectively. The restriction $\gamma_1>0$ is an identifying restriction. The transition function $F(s_t;\gamma_1,c_1)$ allows parameters of the model (2.1) to change smoothly between two regimes that are associated with extreme values of the transition function $F(s_t;\gamma_1,c_1)=0$ and $F(s_t;\gamma_1,c_1)=1$. As s_t increases, the logistic function $F(s_t;\gamma_1,c_1)$ changes monotonically from 0 to 1, with the change being symmetrically around c_1. The slope parameter γ_1 determines the smoothness of transition from one regime to another. This function is appropriate for modeling dynamic behavior of macroeconomic variables that depends nonlinearly on the phase of business cycles (Teräsvirta and Anderson, 1992).

Transition variable in the transition function $G(t;\gamma_2,c_2)$ in Eq. (2.3a) is time trend (t), which allows parameters of the model (2.1) to change smoothly over time. The parameter γ_2 is a slope parameter that determines the speed of change and reciprocal of c_2 represents the (average) location of the parameter change (Lin and Teräsvirta, 1994). The change is monotonic as $G(t;\gamma_2,c_2)$ is a monotonic function of t. In order to allow for non-monotonic structural changes, Lin and Teräsvirta (1994) proposed following quadratic and cubic logistic functions:

$$G(t) = \left[1+\exp\left\{-\gamma_2[t-\alpha]^2\right\}\right]^{-1} \tag{2.3b}$$

or

$$G(t) = \left[1+\exp\left\{-\gamma_2\left[t^3+\alpha_1t^2+\alpha_2t+\alpha_3\right]\right\}\right]^{-1}. \tag{2.3c}$$

[3] The problem of endogeneity is a disputable issue in output growth–inflation literature (see, e.g., Fischer, 1993; Khan and Senhadji, 2001). Therefore, after estimating linear bivariate model for inflation rate and output gap, we carried out Granger-causality tests. The null hypothesis that inflation rate does not Granger-cause output gap was rejected with a p-value 0.051. On the other hand, the null hypothesis that output gap does not Granger-cause inflation rate could not be rejected with a p-value 0.512. This result complies with the earlier empirical findings that causality runs only from inflation rate to output (e.g., Fischer, 1993; Andres and Hernando, 1997). However, bearing in the mind that the dynamic relationship of variables may be nonlinear in nature, we proceeded to estimate a nonlinear bivariate model for the variables under investigation. In fact, after estimating the nonlinear model, we performed regime-dependent Granger-causality tests, results of which suggest that the causality between the two variables is indeed regime dependent and that there exists a bi-directional causality between these variables. Our estimated nonlinear model suggests that there are four distinct regimes (depending on the extreme values of the transition functions) that govern the relationship between output gap and inflation rate. The null hypothesis that output gap does not Granger-cause inflation rate was rejected at 1% significance level for all four regimes. Similarly, the null hypothesis that inflation rate does not Granger-cause output gap was rejected at 1% significance level in two regimes, at 5% significance level in one regime and at 10% significance level in the remaining regime. The test statistics and their details are available upon request.

The parameters $\alpha_1, \alpha_2, \alpha_3$ have no clear interpretation and one can look at the graph of the estimated transition function $G(t;\, \gamma_2,\, c_2)$ to understand the nature of the parameter change. Since economic theory does not suggest any choice for the transition function $G(t;\, \gamma_2,\, c_2)$, Lin and Teräsvirta (1994) proposed a statistical selection technique. Lin and Teräsvirta (1994) also suggested test procedures to test against structural change and provided evidence that their tests are superior to CUSUM and Fluctuation tests in small samples.

The bivariate TV-STR model in Eq. (2.1) can be regarded as a special case of the multiple regime STR model of van Dijk and Franses (1999). TV-STR model nests time-varying and nonlinear smooth transition models as special cases. For example, if $\Phi_{j,i}=0$, $j=3,4$, $i=0,1,2,\ldots,p$ in Eq. (2.1), the model reduces to a two regime smooth transition vector autoregressive model. With $\Phi_{j,i}=0$, $j=2,4$, $i=0,1,2,\ldots,p$ one obtains a time-varying vector autoregressive model. One can follow the specification procedure of Lundbergh et al. (2003) in order to decide whether or not the full bivariate TV-STR as given in Eq. (2.1) is required, or the time series under consideration may be adequately characterized by a nested sub-model, such as a TV-VAR or a STR model.

Lundbergh et al. (2003) suggest two approaches to choose among alternative models. The “specific-to-general” approach consists of several specification, estimation, and evaluation stages, starting with a linear model and proceeding toward the full TV-STR model via a STR or TV model. The “specific-to-general-to-specific” approach, on the other hand, begins with testing linearity directly against the TV-STR model (specific-to-general). If linearity is rejected, then sub-hypotheses are tested to determine whether a STR model or a TV model provides an adequate characterization of the time series at hand (general-to-specific).

The specification tests, however, are complicated by the presence of unidentified nuisance parameters. In particular, under the null hypothesis of linearity, $\Phi_{j,i}$, $j=2,3,4$, $i=0,1,2,\ldots,p$ are unidentified nuisance parameters, which renders standard asymptotic inference invalid. Following Luukkonen et al. (1988), Lundbergh et al. (2003) circumvent this problem by approximating the two transition functions in Eq. (2.1) by a first-order Taylor expansion around the null hypothesis. After replacing the transition functions in Eq. (2.1) by their first-order Taylor expansion, one obtains the following auxiliary regression model:

$$x_t = \Psi_{1,0} + \sum_{i=1}^{p} \Psi_{1,i}x_{t-i} + \sum_{i=1}^{p} \Psi_{2,i}x_{t-i}s_t + \sum_{i=1}^{p} \Psi_{3,i}x_{t-i}t + \sum_{i=1}^{p} \Psi_{4,i}x_{t-i}s_t t + e_t \tag{2.4}$$

where the vector $\mathbf{e}_t$ comprises the original shocks $\boldsymbol{\varepsilon}_t$ as well as the error arising from the Taylor approximation. In Eq. (2.4) it is assumed that the transition variable s_t is one of the elements of $\mathbf{x}_t$. If this is not the case, then one must add $\boldsymbol{\Psi}_5 s_t$ to the auxiliary regression model (2.4). The parameters in $\boldsymbol{\Psi}_{j,i}$, $j=1,2,3,4$, $i=0,1,2,\ldots,p$ are functions of the parameters $\boldsymbol{\Phi}_{j,i}$, $j=1,2,3,4$, $i=0,1,2,\ldots,p$ in the bivariate TV-STR model in Eq. (2.1).

In Eq. (2.4), it is clear that $\boldsymbol{\Psi}_1=\boldsymbol{\Phi}_1$ and $\boldsymbol{\Psi}_{2,i}=\boldsymbol{\Psi}_{3,i}=\boldsymbol{\Psi}_{4,i}=0$ if and only if $\gamma_1=\gamma_2=0$ in Eq. (2.1). Therefore, the null hypothesis of linearity in the auxiliary regression model (2.4) can be written as H_0: $\boldsymbol{\Psi}_{2,i}=\boldsymbol{\Psi}_{3,i}=\boldsymbol{\Psi}_{4,i}=\mathbf{0}$, which can be tested directly by a likelihood ratio (LR) test. Let $\boldsymbol{\Omega}_0=\sum \hat{\mathbf{e}}_t\hat{\mathbf{e}}_t'/T$ and $\boldsymbol{\Omega}_1=\sum \hat{\mathbf{e}}_t\hat{\mathbf{e}}_t'/T$ be the estimated variance–covariance matrices of residuals from the restricted and unrestricted regressions, respectively. Then the LR test statistic for linearity of a k variable VAR model with p lags is given by $LR=T\{\log|\boldsymbol{\Omega}_0|-\log|\boldsymbol{\Omega}_1|\}$, which is asymptotically distributed $\chi^2(3pk^2)$.

Special cases of the auxiliary regression (2.4) yield other linearity tests. It follows that (a) $\boldsymbol{\Psi}_2=0$ if and only if $\gamma_1=0$, (b) $\boldsymbol{\Psi}_3=0$ if and only if $\gamma_2=0$, and (c) $\boldsymbol{\Psi}_4=0$ if $\gamma_1=0$ or $\gamma_2=0$ (see, Lundbergh et al., 2003). Therefore, assuming that $\boldsymbol{\Psi}_3=\boldsymbol{\Psi}_4=0$ in Eq. (2.4), linearity can be tested against STAR type nonlinearity by testing H_0: $\boldsymbol{\Psi}_2=0$. Similarly, linearity can be tested against smoothly varying parameters by assuming $\boldsymbol{\Psi}_2=\boldsymbol{\Psi}_4=0$ and testing H_0: $\boldsymbol{\Psi}_3=0$.

2.2. Testing against TV-STR

Lundbergh et al. (2003) suggest two approaches for specifying and testing linearity against TV-STR-type nonlinearity, namely, “specific-to-general” approach and “specific-to-general-to-specific” approach. The “specific-to-general” approach can be summarized as follows:

(1) Specify an appropriate linear model for vector time series $\mathbf{x}_t$ under investigation.
(2) Test the null hypothesis of linearity separately against smoothly changing parameters [H_0: $\boldsymbol{\Psi}_3=0$ by assuming that $\boldsymbol{\Psi}_2=\boldsymbol{\Psi}_4=0$ in Eq. (2.4)] and against STR-type nonlinearity [H_0: $\boldsymbol{\Psi}_2=0$ by assuming that $\boldsymbol{\Psi}_3=\boldsymbol{\Psi}_4=0$ in Eq. (2.4)]. To identify the appropriate transition variable s_t, the LR test can be computed for several candidates, and the one for which the p-value of the test statistics is smallest can be selected. If the null of linearity is not rejected against either alternative, then retain the linear model.
(3) Estimate the model under the alternative for which the null hypothesis is rejected most convincingly (in terms of p-value), and test for against additional nonlinear or time-varying structure. For example, if linearity is rejected most strongly against STR-type nonlinearity, then estimate a two-regime STR model and test its parameter constancy against the alternative of smoothly changing parameters using the methodology of Eitrheim and Teräsvirta (1996).
(4) Estimate a bivariate TV-STR model if the null of no remaining nonlinearity or parameter constancy is rejected in step 3. Evaluate the model by computing misspecification tests of no remaining nonlinearity and parameter constancy, using generalizations of the tests of Eitrheim and Teräsvirta (1996) for evaluating STR models. Otherwise, tentatively accept the null model.

As Lundbergh et al. (2003) argue, the “specific-to-general” approach has some drawbacks. First, it may involve the estimation of several nonlinear models. Another complication that may arise in the specific-to-general approach is caused by the fact that the tests of linearity against STR and TV-VAR used in step 2 are not robust to structural change and to nonlinearity. Hence these tests cannot discriminate perfectly between these competing alternatives. In addition, the form of the final model is conditional on the path selected. One may reach to a TV-STR model both through a STR and through a TV-VAR model, and it is likely that these two paths lead to different final models.

Taking into account these drawbacks of the “specific-to-general” procedure, one may follow the so-called “specific-to-general-to-specific” approach, which consists of the following steps:

(1) Specify an appropriate linear model for vector time series $\mathbf{x}_t$ under investigation.
(2) Use the LR test discussed above to test the null hypothesis of linearity directly against the TV-STR alternative [$H_0^{\text{TV-STR}}$: $\boldsymbol{\Psi}_2=\boldsymbol{\Psi}_3=\boldsymbol{\Psi}_4=0$ in Eq. (2.4)]. To identify the appropriate transition variable s_t, the LR test can be computed for several candidates, and the one for which the p-value of the test statistics is smallest can be selected. If the null of linearity is not rejected, then retain the linear model.
(3) If the null hypothesis is rejected, test the sub-hypotheses which are nested in $H_0^{\text{TV-STR}}$ to assess whether a TV-STR model is really necessary to characterize time series $\mathbf{x}_t$ or whether either a STR

model or a model with smoothly changing parameters is sufficient. In particular, one can test

$$H_0^{\mathrm{STR}}: \Psi_2 = \Psi_4 = 0$$

and

$$H_0^{\mathrm{TV-VAR}}: \Psi_3 = \Psi_4 = 0.$$

It follows that under H_0^{STR} the model reduces to a model with smoothly changing parameters (i.e., to TV-VAR model) whereas under $H_0^{\mathrm{TV\text{-}VAR}}$ a STR model results. These results lead to the following decision rule:

- If both H_0^{STR} and $H_0^{\mathrm{TV\text{-}VAR}}$ are rejected, select a TV-STR model;
- If H_0^{STR} is rejected and $H_0^{\mathrm{TV\text{-}VAR}}$ is not, then select a STR model;
- If H_0^{STR} is not rejected but $H_0^{\mathrm{TV\text{-}VAR}}$ is, select a model with smoothly changing parameters (i.e., TV-VAR model).

The only combination of test results which does not lead to a clear-cut model choice is when neither H_0^{STR} nor $H_0^{\mathrm{TV\text{-}VAR}}$ is rejected but the general null hypothesis $H_0^{\mathrm{TV\text{-}STR}}$ is rejected. In this case, one may resort to the LR tests, which test linearity against STR and against TV-VAR separately as in step 2 of the specific-to-general approach. In this way one can find out which (if any) of the sub-models of the TV-STR model is the most suitable for the time series at hand.

(4) Estimate the selected model and evaluate it by misspecification tests.

As Lundbergh et al. (2003) argue, “specific-to-general-to-specific” modeling strategy has two drawbacks. First, testing directly against the full TV-STR model may imply weak power, because the dimension of the null hypothesis may be large in many cases. Second, testing sub-hypotheses within the auxiliary regression (2.4) may not be appropriate, because the original null hypothesis $H_0^{\mathrm{TV\text{-}STR}}$ has already been rejected. In other words, the remainder in Eq. (2.4) may not be equal to 0 under the null hypotheses H_0^{STR} and $H_0^{\mathrm{TV\text{-}VAR}}$, which may affect the size of the corresponding tests. Thus the additional tests may be seen just as additional model selection devices that are likely to be helpful if the difference in p-values of the tests is reasonably large.

In passing we note that if the linearity tests described above suggest that the appropriate model is a bivariate STR model, then one may follow the procedure of Teräsvirta (1994) to decide whether a logistic or an exponential function is more convenient transition function. Similarly, one may apply the test procedure of Lin and Teräsvirta (1994) to choose a suitable transition function. Once appropriate model is chosen, the model can be estimated by nonlinear least squares. After estimating model, we use generalized impulse response functions to examine the dynamic relationships between inflation rate and output gap.

3. Data, estimation results and discussion

In this paper we use quarterly data spanning the period 1980:Q1–2008:Q3. This time period was dictated by data availability. The inflation rate is defined as $\pi_t = \ln(\mathrm{CPI}_t) - \ln(\mathrm{CPI}_{t-1})$ where CPI is consumer price index. The output gap here is proxied by (logarithm of the) industrial production index[4] detrended by HP filter due to Hodrick and Prescott (1997).[5] All data are obtained from International Financial Statistics database of the International Monetary Fund. Below we provide and discuss some preliminary data statistics and stationarity test results.

3.1. Preliminary data analysis

Table 1 presents preliminary data statistics. Although we use output gap in this paper, we also added output growth rate to Table 1 in order to provide a better picture of the dynamics of the Turkish economy during the analyzed period.[6]

As the table reveals, both output growth rate and inflation rate have fluctuated widely during the analyzed period. Average (quarterly) inflation rate during the period was around 9.8%, achieving its highest level of 34% in the second quarter of 1994, and lowest level of −0.4% in the third quarter of 2007. Similarly, output growth rate averaged 1.4%, with highest level of 10.7% in the first quarter of 1985 and lowest level of −11.4% in the second quarter of 1994. Output gap was at its lowest level of −11.7% in the same quarter of 1994.

Drastic fluctuations in output growth rate and inflation in Turkey may be attributed to economic crises in 1994 and 2001. As the above figures suggest, the highest (quarterly) inflation rate and output decline was observed during the second quarter of 1994. Annual inflation rate (measured as percentage change in GDP deflator) climbed to its historical high level of 107.3% in 1994.[7] Perhaps, the 2001 crisis was the severest one since the end of the World War 2. GDP fell by 9.5% in 2001, overnight interbank rates rose to above 4000% (the highest value since introduction of the interbank money market), Turkish Lira depreciated by 40% in a day against the USD. Both 1994 and 2001 crises caused output to realize far below its trend level. Measured output gap was negative for ten consecutive quarters following both crises. During the period from the second quarter of 1994 until the third quarter of 1996, inclusive, average output gap was −4.1% and output growth rate averaged 0.6% per quarter. On the other hand, during the period from the first quarter of 2001 until the second quarter of 2003, inclusive, average output gap and output growth rates were −5.2% and 0.3%, respectively, smaller when compared to the period after the 1994 crisis.

Another interesting feature of the data is that both output gap and inflation rate have fat-tailed distributions (i.e., suffer from excess skewness), and therefore, are not normally distributed. As the table shows, inflation rate is (statistically significantly) positively skewed, reflecting perhaps sharp increases during the crisis periods. Similarly, output gap is negatively skewed. Although inflation rate suffers from excess kurtosis, this is not the case for output gap. These features of the data imply that a model that captures nonlinear and time-varying dynamics might be more appropriate to model the relationship between the inflation rate and output gap.

Now, we turn to examination of stochastic properties of the data. The specification procedures described in the previous section rely on the assumption that both the output gap (y_t) and inflation rate (π_t) are $I(0)$ processes. Therefore, prior to estimation of the linear model, we tested stationarity of the variables concerned. It is well known that conventional unit-root tests have low power if the true data generating process is subject to regime changes (see, e.g., Kapetanios et al., 2003). If the process is globally stationary but exhibits unit root

[4] Gross Domestic Product (GDP) is the most appropriate measure of output for the purposes of this study. However, GDP is available only starting from 1987, whereas industrial production (IP) is available from 1980. Therefore, in order to obtain (economically and statistically) more meaningful results, we proxy output by the IP. The IP is, perhaps, the best alternative for the GDP-based measure of output. Indeed, the correlation coefficient between GDP and IP is equal to 0.996, and the time path of the IP mimics that of the GDP quite well.

[5] Nigmatullin et al. (2010) develop a new fractional filtering technique and compare properties of the newly developed procedure to those of the HP filter. They show that their new procedure and the HP filter capture the dynamics of the Turkish quarterly GDP series quite well.

[6] For a thorough discussion of developments in the Turkish economy since 1980 see, among others, Asikoglu and Uctum (1992), Müslümov et al. (2002), Telatar et al. (2003), Dibooglu and Kibritcioglu (2004), Hasanov and Omay (2008a), and works collected in Kibritcioglu et al. (2002).

[7] Annual inflation rate and GDP growth rate are taken from the State Planning Organization of Republic of Turkey.

Table 1
Summary statistics.

Series	Mean	S.E.	Min	Max	Sk	Ku	J–B
Inflation Rate	0.098	0.059	−0.004	0.340	0.683 [0.003]	1.256 [0.008]	16.358 [0.000]
Output growth rate	0.014	0.034	−0.114	0.107	−0.386 [0.097]	1.357 [0.004]	11.580 [0.003]
Output gap	0.000	0.038	−0.117	0.085	−0.473 [0.041]	0.734 [0.118]	6.889 [0.032]

S.E. denotes standard error, Sk denotes excess skewness, Ku denotes excess kurtosis, and J–B denotes Jarque–Berra's test for normality of series. p-values of diagnostic tests are provided in square brackets.

or explosive behavior in one of the regimes, then the test procedures that ignore regime-dependent dynamics and nonlinearities might be biased against stationarity (for a thorough discussion of this issue, see Taylor et al., 2001; Kapetanios et al., 2003). Therefore, in addition to conventional ADF and PP tests, we also applied the unit-root test procedure of Kapetanios et al. (2003) (the KSS test), which has a good power when the series under investigation follow a nonlinear STAR process. The results of these tests are provided below in Table 2.

As can be seen from the table, all three tests suggest that output gap is stationary. As regards inflation rate, only ADF test failed to reject the null hypothesis of unit root. This result may be due to the fact that ADF test has a low power against nonlinear stationary processes. Therefore we conclude that both the inflation rate and output gap are $I(0)$ processes, and proceed to model specification.

3.2. Model specification and linearity test results

As discussed in the previous section, both "specific-to-general" and "specific-to-general-to-specific" modeling strategies have drawbacks. However, as noted by Lundbergh et al. (2003), additional tests may be seen as additional helpful model selection devices. Therefore, taking into account of the fact that the first two steps in both model specification approaches are specification of an appropriate linear model and testing against linearity, we first estimate a bivariate model for the inflation rate and output gap. Then, in the next step, we test linearity separately against a STR-type nonlinearity and smoothly changing parameters as prescribed in step 2 of the "specific-to-general" modeling strategy as well as directly against the TV-STR alternative as in the step 2 of the "specific-to-general-to-specific" approach. The advantage of doing so is that if linearity tests in both approaches indicate either a STR nonlinearity or a smoothly changing parameter model, then burden on estimation of unnecessary nonlinear models may be alleviated.

As a first step, we estimate a linear bivariate model for the inflation rate and output gap. Since the nonlinearity tests are sensitive to autocorrelation, the autoregressive structure of the model should be specified so as to capture significant autocorrelation in the linear model. The lag lengths of each variable in each equation were selected by applying conventional Akaike Information Criterion (AIC), and then the resultant model was tested against autocorrelation of residuals. Considering the fact that the variables may exhibit serious seasonality, we added seasonal dummies to each equation. It is a well-known fact that if the residuals from the estimated equations are correlated, then Generalized Least Squares (GLS) estimates will bring efficiency gains over the Ordinary Least Squares (OLS) estimates (see, Zellner, 1962; Greene, 1997: 675–676). Therefore, we estimated the inflation and output equations simultaneously using the GLS estimator iteratively, which gives maximum likelihood (ML) estimates (Greene, 1997: 681–682).[8] The ML estimates of the inflation and output equations are given below in Table 3.[9]

Table 2
Unit-root test results.

	ADF test	PP test	KSS test
Output gap	−3.947 (4)***	−5.149 (4)***	−2.080 (4)*
Inflation rate	−1.460 (3)	−6.140 (3)***	−2.490 (3)**

Notes: Test regressions include only an intercept term. Number of augmentation terms in the regressions is indicated in parenthesis. *, **, *** denote rejection of the null hypothesis of unit root at 10%, 5% and 1% significance levels, respectively.

As Table 3 reveals, residuals from both the inflation and output equations pass diagnostic tests. Particularly, residuals do not suffer from excess kurtosis and skewness, and are normally distributed. Furthermore, the tests reveal neither conditional heteroscedasticity nor autocorrelation. As both equations seem to be satisfactory, now we proceed to test linearity for the bivariate model.

The results of the linearity tests are provided in Table 4. In Panel A of Table 4 we report results of the linearity tests against STR-type nonlinearity as well as against TV-STR alternative as suggested by Lundbergh et al. (2003). In Panel B of the table we provide parameter constancy tests of Lin and Teräsvirta (1994). The tests of sub-hypotheses $H_0^{\text{TV-VAR}}$ and H_0^{STR} as prescribed in the step 3 of the "specific-to-general-to-specific" modeling strategy of Lundbergh et al. (2003) are reported in the Panel C of the table. Finally, Panel D of the table report specification tests of Lin and Teräsvirta (1994) for choosing appropriate transition function that governs structural change.

As can readily be seen from Table 4, the null hypothesis of linearity and structural break is rejected for most of the candidate transition variables. When the candidate transition variable is first lag of the output gap (y_{t-1}), the null hypothesis of linearity is rejected against a STR alternative with a p-value of 0.003. However, the null of linearity is rejected against the alternative of smoothly changing parameters with a p-value of 7.3×10^{-7} when we use LM3 test. This finding implies that a model that allows for both structural break and nonlinearity might be more appropriate for our data. We reach to the same conclusion when we follow the "specific-to-general-to-specific" specification procedure. Specifically, the null of linearity against a TV-STR alternative is rejected more strongly when the candidate transition variable is the second lag of inflation rate (π_{t-2}) with a p-value of 1.90×10^{-6}. When we use this transition variable (π_{t-2}), both H_0^{STR} and $H_0^{\text{TV-VAR}}$ sub-hypotheses are rejected at conventional significance levels, suggesting that the relationship between inflation rate and output gap can be best prescribed by a time-varying smooth transition regression model. We applied the H03 and H02 tests of Lin

[8] We also estimated the model with OLS and computed correlations between the residuals from the inflation and output equations. Although the correlation coefficient was relatively small (0.14) for the full sample, visual inspection of the residuals suggested that correlation between the residuals might have changed significantly during the analyzed period. Therefore, we also computed correlation coefficients for different sub-periods. The computed correlation coefficients varied quite considerably, from −0.79 to 0.32. In the face of such high correlation coefficients and considering that correlation coefficient measures only linear relationships, we opted to use GLS estimator iteratively, which gives ML estimates. In fact, the ML estimates, when compared to OLS estimates (not reported here, but available upon request), provide a better results in that, by using the GLS method iteratively we obtained lower coefficient standard errors, lower residual sum of squares, lower standard errors of estimate, and higher R^2 for both inflation and output equations.

[9] Each equation in the model includes dummy variables for outliers evident in the residuals. We first estimated the equations without dummy variables, and computed standard deviations of the residuals. Then, outliers are defined as those observations which are three times larger (in absolute value) than the standard deviation. We include these dummy variables into the equations in order to ensure that rejection of the null hypothesis of linearity is not due to the presence of big outliers. In addition, dummy variables ensure that the transition function does not simply capture small number of outliers. Therefore, we included the same dummy variables in the estimation of the nonlinear models as well. See also Öcal (2000) and Hasanov and Omay (2008b), among others, for use of outlier dummy variables within STR framework.

Table 3
ML estimates of the linear bivariate model.

	Output equation	Inflation equation
Constant	0.001 (0.005)	0.009 (0.006)
y_{t-1}	0.527 (0.077)***	0.133 (0.090)
y_{t-2}	0.198 (0.090)**	−0.066 (0.091)
y_{t-3}	0.167 (0.087)*	–
y_{t-4}	−0.292 (0.074)***	–
π_{t-1}	−0.131 (0.060)**	0.277 (0.070)***
π_{t-2}	0.020 (0.057)	0.334 (0.071)***
π_{t-3}	0.128 (0.060)**	0.132 (0.072)*
π_{t-4}	–	0.121 (0.070)*
$SD1_t$	0.013 (0.008)*	−0.022 (0.009)***
$SD2_t$	0.015 (0.008)**	−0.046 (0.009)***
$SD3_t$	0.001 (0.006)	−0.062 (0.009)***
D_y	−0.100 (0.014)***	–
D_π	–	0.154 (0.016)***
$\overline{R}^2$	0.691	0.820
Sum of squared residuals	0.055	0.074
Residual diagnostic tests		
Excess skewness	0.128 [0.595]	0.057 [0.811]
Excess kurtosis	−0.181 [0.711]	−0.033 [0.947]
J–B normality test	0.437 [0.804]	0.064 [0.969]
Ljung–Box Q(1)	0.184 [0.668]	0.405 [0.524]
Ljung–Box Q(4)	0.784 [0.941]	0.695 [0.952]
ARCH(1)	0.319 [0.572]	0.092 [0.762]
ARCH(4)	1.120 [0.891]	2.549 [0.636]

Notes: $SD1_t$, $SD2_t$ and $SD3_t$ are seasonal dummy variables. D_π and D_y are outlier dummy variables for inflation equation and output equation, respectively. Standard errors of parameter estimates are provided in parenthesis. ***, **, and * denote significance at 1%, 5%, and 10% significance levels, respectively. J–B is Jarque–Berra's test for normality of residuals. Ljung–Box Q(j) denotes Ljung–Box's (1979) Q test for residual autocorrelation of order j. ARCH(j) is Engel's (1982) LM test for jth order autoregressive conditional heteroscedasticity. p-values of residual diagnostic tests are shown in square brackets.

Table 4
Linearity test results.

Candidate transition variable	LR test against STR-type nonlinearity	LR test against TV-STR alternative
Panel A: linearity tests		
y_{t-1}	40.261 (0.003)***	114.867 (8.84×10^{-6})***
y_{t-2}	38.702 (0.005)***	115.660 (7.15×10^{-6})***
y_{t-3}	33.129 (0.023)**	107.617 (5.85×10^{-5})***
y_{t-4}	26.207 (0.125)	93.253 (1.74×10^{-3})***
π_{t-1}	33.604 (0.020)**	105.376 (1.02×10^{-4})***
π_{t-2}	33.208 (0.023)**	120.497 (1.90×10^{-6})***
π_{t-3}	38.537 (0.005)***	115.588 (7.29×10^{-6})***
π_{t-4}	32.144 (0.030)**	113.903 (1.14×10^{-5})***
Panel B: linearity tests against smoothly changing parameters		
LM1	45.757 (0.001)***	
LM2	84.665 (1.06×10^{-4})***	
LM3	132.509 (7.3×10^{-7})***	
Panel C: model specification tests		
H_0^{STR}	77.289 (1.72×10^{-4})***	
$H_0^{TV\text{-}VAR}$	87.289 (9.51×10^{-6})***	
Panel D: transition function specification tests		
H03	47.844 (7.22×10^{-4})***	
H02	38.908 (0.010)**	
H01	45.757 (0.001)***	

Notes: p-values of the test statistics are provided in parenthesis. LM1, LM2 and LM3 are Lin and Teräsvirta's (1994) LM-type tests against parameter constancy whereas smooth change in parameters can be best prescribed by a first-order, second-order, and third-order logistic transition function, respectively. H03, H02 and H01 are specification tests for choosing the order of the logistic transition function. Note that H01 test is the same as LM1 test against parameter constancy where the transition function is a first-order logistic function in time trend. For further discussion see Lin and Teräsvirta (1994: pp. 215–217). *** and ** denote significance at 1% and 5% significance levels, respective.

and Teräsvirta (1994) to choose suitable transition function for structural break. Since the H03 test is rejected more strongly, we choose a third-order logistic transition function.

3.3. Estimation results

After choosing the appropriate model (TV-STR) for our data, we used maximum likelihood estimator to estimate the parameters of the model, which are provided in Table 5. Fig. 1 plots the graph of the

Table 5
Estimates of the TV-STR model.

	Output equation	Inflation equation
Constant	0.015 (0.013)	0.024 (0.014)*
y_{t-1}	0.806 (0.174) ***	−0.009 (0.163)
y_{t-2}	−0.059 (0.162)	−0.620 (0.159)***
y_{t-3}	0.351 (0.163)**	–
y_{t-4}	−0.531 (0.144)***	–
π_{t-1}	−0.277 (0.099) ***	0.063 (0.099)
π_{t-2}	0.187 (0.136)	0.561 (0.134)***
π_{t-3}	0.013 (0.119)	−0.056 (0.113)
π_{t-4}	–	0.259 (0.107)**
$F(\pi_{t-2})$	−0.033 (0.012)***	−0.024 (0.012)**
$F(\pi_{t-2})\cdot y_{t-1}$	−0.467 (0.239)*	0.366 (0.233)
$F(\pi_{t-2})\cdot y_{t-2}$	0.185 (0.220)	0.644 (0.218)***
$F(\pi_{t-2})\cdot y_{t-3}$	−0.368 (0.218)*	–
$F(\pi_{t-2})\cdot y_{t-4}$	0.397 (0.208)*	–
$F(\pi_{t-2})\cdot \pi_{t-1}$	0.002 (0.097)	0.216 (0.098)**
$F(\pi_{t-2})\cdot \pi_{t-2}$	−0.033 (0.064)	−0.057 (0.090)
$F(\pi_{t-2})\cdot \pi_{t-3}$	0.237 (0.068)***	0.002 (0.067)
$F(\pi_{t-2})\cdot \pi_{t-4}$	–	0.004 (0.072)
$G(t)$	−0.020 (0.020)	0.015 (0.018)
$G(t)\cdot y_{t-1}$	−0.230 (0.280)	−0.043 (0.272)
$G(t)\cdot y_{t-2}$	0.564 (0.297)*	0.057 (0.277)
$G(t)\cdot y_{t-3}$	−0.286 (0.275)	–
$G(t)\cdot y_{t-4}$	0.265 (0.231)	–
$G(t)\cdot \pi_{t-1}$	0.394 (0.209)*	0.394 (0.195)**
$G(t)\cdot \pi_{t-2}$	−0.424 (0.257)*	−0.937 (0.261)***
$G(t)\cdot {}^*\pi_{t-3}$	0.230 (0.227)	0.202 (0.208)
$G(t)\cdot \pi_{t-4}$	–	−0.695 (0.178)***
$F(\pi_{t-2})\cdot G(t)$	0.072 (0.025)***	0.069 (0.023)***
$F(\pi_{t-2})\cdot G(t)\cdot y_{t-1}$	0.413 (0.358)	−0.130 (0.350)
$F(\pi_{t-2})\cdot G(t)\cdot y_{t-2}$	−0.213 (0.354)	−0.198 (0.353)
$F(\pi_{t-2})\cdot G(t)\cdot y_{t-3}$	0.438 (0.321)	–
$F(\pi_{t-2})\cdot G(t)\cdot y_{t-4}$	−0.527 (0.306)*	–
$F(\pi_{t-2})\cdot G(t)\cdot \pi_{t-1}$	−0.251 (0.227)	−0.415 (0.212)*
$F(\pi_{t-2})\cdot G(t)\cdot \pi_{t-2}$	0.192 (0.173)	0.413 (0.216)*
$F(\pi_{t-2})\cdot G(t)\cdot \pi_{t-3}$	−0.494 (0.165)***	0.110 (0.154)
$F(\pi_{t-2})\cdot G(t)\cdot \pi_{t-4}$	–	0.230 (0.149)
$SD1_t$	−0.002 (0.007)	−0.071 (0.008)***
$SD2_t$	0.134 (0.008)***	−0.038 (0.008)***
$SD3_t$	−0.000 (0.008)	−0.029 (0.008)***
D_y	−0.102 (0.016)***	–
D_π	–	0.166 (0.013)***
Estimated transition functions:	$G(t) = \left[1 + \exp\left\{-\underset{(0.001)^{**}}{0.003}[t-\underset{(0.011)^{***}}{0.549}]^3\right\}\right]^{-1}$ $F(\pi_{t-2}) = [1 + \exp\{-\underset{(47.539)}{58.776}[\pi_{t-2}-\underset{(0.001)^{***}}{0.093}]/\delta_{\pi_{t-2}}\}]^{-1}$	
$\overline{R}^2$	0.756	0.893
Sum of squared residuals	0.039	0.039
Residual diagnostic tests		
Excess skewness	−0.152[0.530]	−0.044 [0.855]
Excess kurtosis	0.057 [0.907]	−0.072 [0.884]
J–B normality test	0.421 [0.810]	0.057 [0.972]
Ljung–Box Q(1)	2.231 [0.135]	1.207 [0.272]
Ljung–Box Q(4)	3.255 [0.516]	5.843 [0.211]
ARCH(1)	0.333 [0.564]	0.129 [0.720]
ARCH(4)	3.234 [0.519]	3.055 [0.549]

Notes: $SD1_t$, $SD2_t$ and $SD3_t$ are seasonal dummy variables. D_π and D_y are outlier dummy variables for inflation equation and output equation, respectively. Standard errors of parameter estimates are provided in parenthesis. ***, **, and * denote significance at 1%, 5%, and 10% significance levels, respectively. J–B is Jarque–Berra's test for normality of residuals. Ljung–Box Q(j) denotes Ljung–Box's (1979) Q test for residual autocorrelation of order j. ARCH(j) is Engel's (1982) LM test for jth order autoregressive conditional heteroscedasticity. p-values of residual diagnostic tests are shown in square brackets.

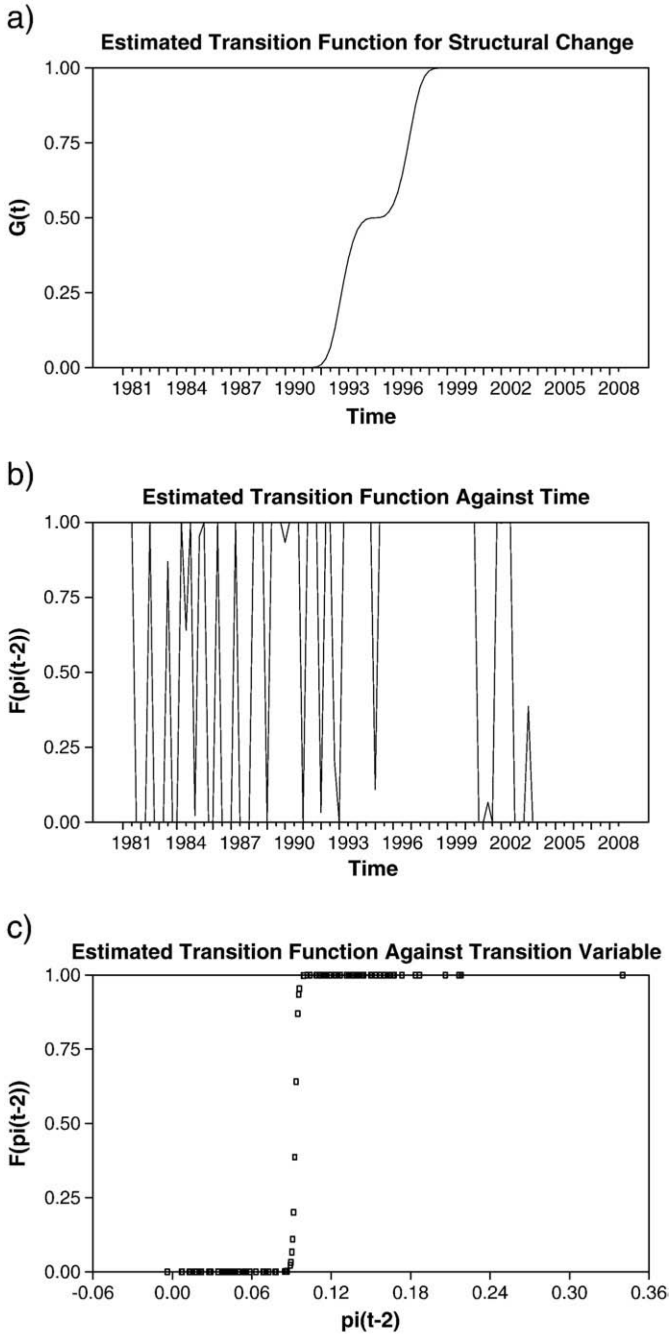

Fig. 1. Estimated transition functions.

estimated transition function $\hat{G}(t)$ against time as well as the graph of $\hat{F}(\pi_{t-2})$ against time and against transition variable π_{t-2}.

The estimated transition function $\hat{G}(t)$ in Fig. 1(a) implies that the change in parameters of the estimated model happened gradually over time. The structural change is monotonic, starts (roughly) in the first quarter of 1992, slows temporarily in mid 1990s, and is over by 1998. This finding is very impressive, especially when the developments in the Turkish economy during the last three decades are taken into account. Specifically, as discussed briefly earlier, starting from 1980, Turkey has abandoned import-substitution policies and adopted export-led growth strategies. Price controls have been lifted, import regime was gradually liberalized, and foreign direct investments and export encouraging policies were followed. In addition, interest rates and foreign exchange regime were gradually liberalized, and new market institutions were established step-by-step. By early 1990s Turkey has managed to liberalize its economy and integrate to the world economy to a large extent. The estimated transition function indicates that these reforms in the Turkish economy have caused to structural changes in the inflation–output relationship whereas the change was gradual. The slowdown in parameter change in mid 1990s may be explained by the economic crisis in April 1994.

The estimated slope parameter $\hat{\gamma}=58.776$ in the second transition function $\hat{F}(\pi_{t-2})$ as well as the graphs of the transition function against time and the transition variable $s_t=\pi_{t-2}$ in Fig. 1(b) and (c) indicates that the transition between two regimes, $\hat{F}(\pi_{t-2};\gamma_1, c_1)=0$ and $\hat{F}(\pi_{t-2};\ \gamma_1,\ c_1)=1$, is rather rapid. The two regimes $\hat{F}(\pi_{t-2};\ \gamma_1,\ c_1)=0$ and $\hat{F}(\pi_{t-2};\ \gamma_1,\ c_1)=1$ are identified by the value of the transition variable (π_{t-2}) relative to the location parameter $\hat{c}_1=0.093$. Particularly, for relatively lower values of inflation rate (approximately for $\pi_{t-2}<0.085$, i.e., 8.5% per quarter) $\hat{F}(\pi_{t-2})\approx 0.0$, and for higher values of inflation rate (approximately for $\pi_{t-2}>0.10$, i.e., 10% per quarter) $\hat{F}(\pi_{t-2})\approx 1.0$. These results imply that the dynamic interrelationship between output gap and inflation rate depends not only time, but on past inflation rates as well.[10]

3.4. Impulse response functions and discussion

The parameters of the estimated VAR model are difficult to interpret. Therefore, one may use impulse response functions to figure out the dynamic relationships between variables in the estimated VAR model. Since nonlinear time series models do not have a Wold representation, however, traditional impulse response functions cannot be computed for the nonlinear models. Koop et al. (1996) introduced the generalized impulse response function (GIRF) that may be used for defining impulse responses in nonlinear models. The method of computation and features of the generalized impulse response functions are discussed by Koop et al. (1996) in a great detail. For a brief discussion, see the Appendix in Weise (1999).

Since GIRF is history dependent (i.e., impulse response depends on the particular history when the shock hits the system), we computed GIRF for three different time periods. The first sub-period spans 1981Q3–1991Q4,[11] when the estimated transition function $\hat{G}(t)$ takes on value less than 0.01, the second sub-period spans 1994Q1–1996Q1 when $0.45<\hat{G}(t)<0.55$, and the third sub-period spans 1998Q1–2008Q3 when $\hat{G}(t)>0.99$. This choice of sub-periods is dictated by the nature of parameter changes in the estimated TV-STR model and is interesting in that each sub-period can be characterized by different macroeconomic environment.

In order to see whether the dynamic effects of one variable on the other is regime dependent, we calculate GIRF for each regime separately, i.e., for lower inflation regime that is associated with $\hat{F}(\pi_{t-2})=0.0$ and for higher inflation regime that is associated with $\hat{F}(\pi_{t-2})=1.0$ for each mentioned sub-period. In addition, a large body of theoretical and empirical research suggests that negative and positive demand shocks may have asymmetric effects on output (e.g., Tsiddon, 1993; Ball and Mankiw, 1994; Cover, 1992; Telatar and Hasanov, 2006). Therefore, for each sub-period and regime we calculate the effects of both negative and positive shocks. Figs. 2–4 graph the computed cumulative responses of one variable to a one-time shock to the other variable. The magnitude of the shock is set to the standard deviation of the innovations computed from the TV-STR model. The responses to negative shocks are plotted with reversed sign so as to compare them with the responses to positive shocks.

10 This particular threshold value of inflation rate that separates the two regimes is consistent with the earlier literature on the inflation–output relationship. Particularly, Fischer (1993), Barro (1996), Khan and Senhadji (2001), among others, examine how inflation rate affects output growth when inflation rate exceeds some threshold values.

11 Due to data transformation and lags used in the estimated model, we lose 6 observations. Therefore, although our data span the period 1980Q1–2008Q3, actual estimations were carried for the 1981Q3–2008Q3 period.

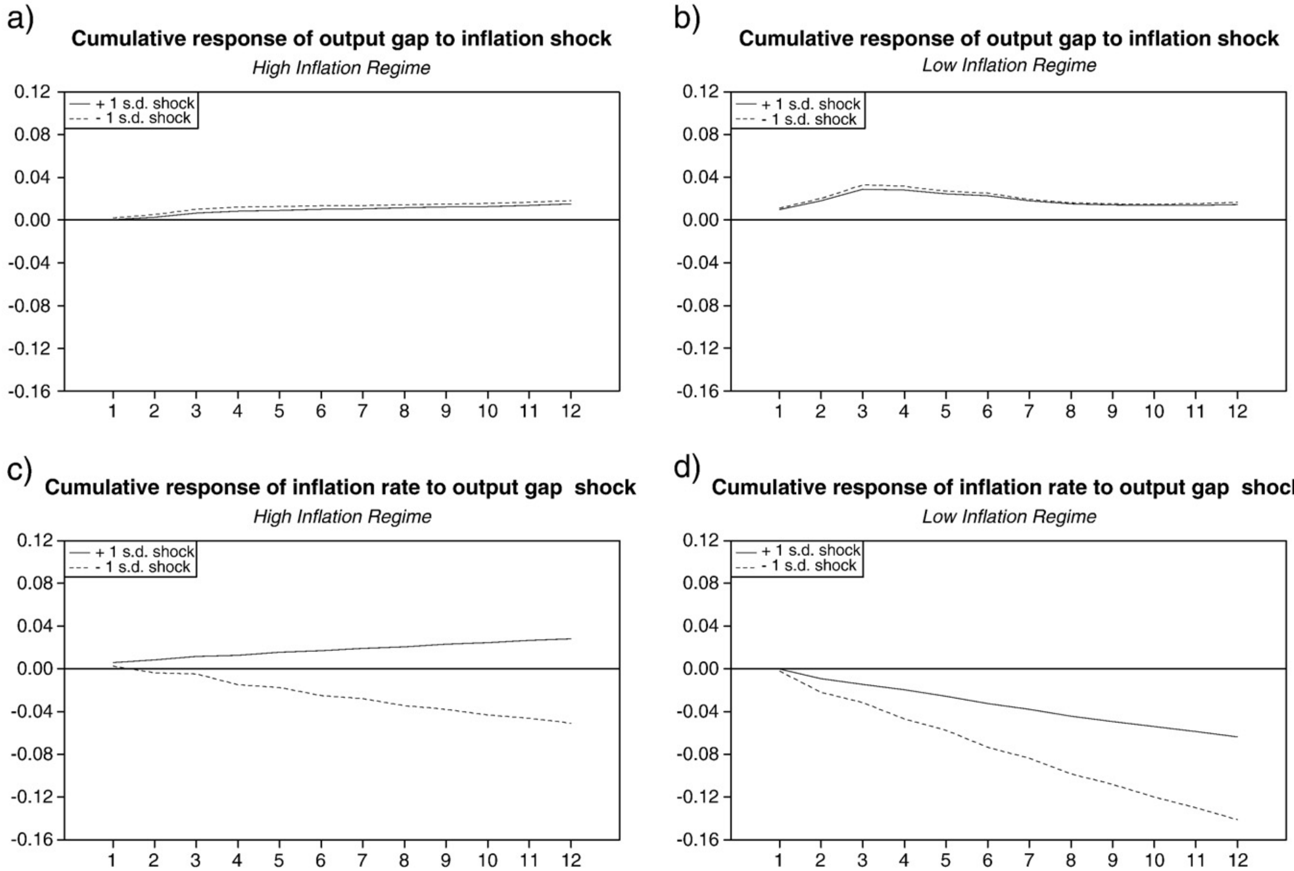

Fig. 2. Cumulative response functions in the sub-period 1981:03–1991:04.

3.4.1. The output–inflation relationship during the 1981–1991 sub-period

The first sub-period from 1981 to 1991 can be characterized as the years of liberalization and integration of the Turkish economy to the world economy.[12] Starting with the economic stabilization program that was launched on January 24, 1980 in response to a severe balance-of-payment crisis and economic recession, Turkey has abandoned inward oriented import-substitution strategies and adopted export-led growth strategies. The reforms implemented throughout 1980s aimed at liberalization of import regime and encouraging exports and foreign direct investments, adopting a flexible exchange rate regime, liberalization of current account transactions, removing price controls and interest rate restrictions. In addition, in order to foster economic growth further, the government undertook huge infrastructure investments and established new market institutions. The budget deficits were initially financed from central bank resources. After the introduction of domestic borrowing instruments in 1984, however, domestic borrowing has gained importance in financing budget deficits. In 1989, capital flows were liberalized and Turkey adopted a convertibility policy for the Turkish Lira. In 1990 foreign exchange regime was further liberalized, and controls on interest rates were fully removed. Thus, by early 1990s Turkey completed the first stage of liberalization reforms.

Turkey enjoyed relatively high growth rates during the period from 1981 to 1991. Average GDP growth rate was 4.81% and average annual inflation rate was 48.28% during this period. The engine of economic growth during this period was huge infrastructure investments and growth in exports. Turkey allowed real exchange rates to depreciate in order to promote exports until 1988. However, huge public expenditures accelerated inflation rate, and government's concern has shifted from external competitiveness to domestic stability. In order to prevent inflationary effects, real exchange rate depreciation policy aimed at export promotion was abandoned in 1989, and more real appreciations have been allowed. Capital account liberalization and high domestic interest rates intensified capital inflows, which contributed to real appreciation of the domestic currency further. Loss of external competitiveness coupled with high interest rates has slowed output growth starting from late 1980s. The failure to control huge budget deficits accelerated inflation rate. The exchange rate policy aimed at stabilization of exchange rates was unsustainable in the face of increasing foreign trade deficits. In fact, the central bank of Turkey was obliged to abandon exchange rate policy in April, 1994 and Turkey underwent a severe crisis in that year. Now, we turn to examination of output–inflation relationships during the period from 1981 to 1991.

Fig. 2 plots computed cumulative responses of output to a one-time shock to inflation rate as well as of inflation rate to a one-time shock to output in the first sub-period 1981Q3–1991Q4. As can readily be seen from the figure, a shock to inflation rate increases the output gap both in low and high inflation regimes. Such a positive effect of inflation on output is consistent with sticky price and sticky wage models (Fischer, 1977; Taylor, 1980; Caplin and Spulber, 1987; Blanchard and Kiyotaki, 1987). The effects of a shock to inflation rate on output gap are slightly larger in low inflation regimes when compared to high inflation regimes. This result complies with findings of Telatar and Hasanov (2006), who also found that the effects of demand shocks on output are dependent on initial state of the economy in the case of Turkey. However, for this sub-period there is no apparent evidence of asymmetric response of output to negative and positive shocks to inflation rate, which suggests that in this period

[12] See, for example, Asikoglu and Uctum (1992), Müslümov et al. (2002), Dibooglu and Kibritcioglu (2004), Hasanov and Omay (2008a).

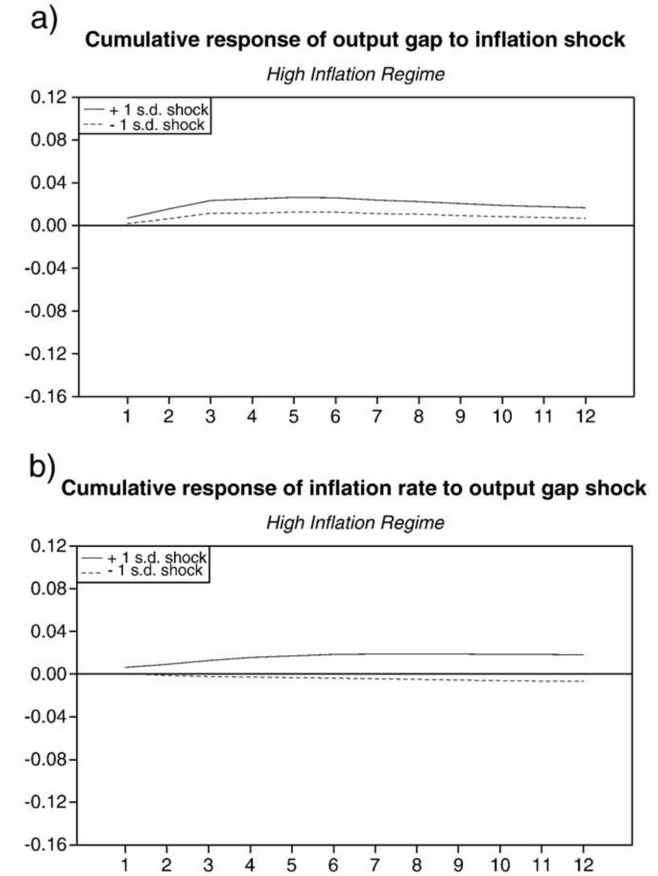

Fig. 3. Cumulative impulse response functions in the sub-period 1994:01–1996:01.

the Phillips curve might be almost linear. On the other hand, we find that the effects of a shock to output gap on inflation rate are both regime dependent and asymmetric. In particular, the graphs of the GIRF in Fig. 2(c) and (d) suggest that the effects of negative shock to output gap is greater than the effects of positive shocks as well as that output shocks affect inflation rate stronger in low inflation regimes. The negative effect of output shock on inflation rate may be due to the fact that massive liberalization reforms implemented in Turkey during the 1980'ies as well as huge public infrastructure investments might have increased potential output in Turkey and shifted long-run aggregate supply curve rightwards.

3.4.2. The output–inflation relationship during the 1994–1996 sub-period

As briefly discussed above, the main objective of monetary policies starting from late 1980s was to stabilize exchange rates in order to fight accelerating inflation rate. However, increasing trade deficits and the failure to control budget deficits rendered exchange rate policy unsustainable. In the face of intensified speculative attacks the central bank abandoned exchange rate stabilization policy and allowed domestic currency to depreciate. As a result, Turkey underwent a severe exchange rate crisis in the first half of 1994. Annual inflation rate achieved its historical high level 107.3% and GDP fell by 6.1% in that year.

On April 5, 1994 Turkey launched IMF-supported stabilization program. The main objective of the stabilization program was to ensure financial and economic stability. Under this program, the Turkish Lira was devaluated, government expenditures were cut, prices of the public goods and tax rates were increased, and new taxes were introduced. The stabilization program has been successful in stabilizing financial markets and reducing budget deficits. As a result, the economy grew by 8% in 1995, while inflation rate remained high at 87.2%. After the exchange rate crisis was defeated successfully, however, the government once more switched to expansionary fiscal policies, mainly due to political considerations. Tight fiscal policies were in place only until the second-half of the year 1995, when general elections were held. Since comprehensive reforms were not undertaken, the April 1994 stabilization program had only a limited effect, and fiscal condition began to deteriorate starting from 1995. Turkey joined the Customs Union with the European Union in 1996, which resulted in further liberalization of the foreign trade regime. Short-term capital inflows continued to finance increased imports.

Now, we discuss output–inflation relationship in the second sub-period 1994Q1–1996Q1. Computed cumulative responses of output to a one-time shock to inflation rate as well as of inflation rate to a one-time shock to output in this sub-period are plotted in Fig. 3. During this sub-period the quarterly inflation rate never fall below 8% and therefore, no low inflation regime was realized during this period. Hence, we computed GIRF only for high inflation regime. It is interesting to note that the type of asymmetry in the effects of shocks has reversed in this period. As the graph suggests, the effects of a positive shock to one variable on the other variable is greater than the effects of a negative shock. Furthermore, negative shocks to output gap had almost no effect on inflation rate. The fact that the effect of a positive shock to inflation rate on output gap is greater than the effect of a negative shock is consistent with a concave Phillips curve, implying that the Phillips curve might has changed its shape during the crisis years.

3.4.3. The output–inflation relationship during the 1998–2008 sub-period

The Turkish economy has experienced wide fluctuations during the period from 1998–2008. The 1998 Russian crisis, two earthquakes in 1999, and the 2000–2001 crises contributed to output volatility during this period (see, e.g., Müslümov et al., 2002; Dibooglu and Kibritcioglu, 2004).

Turkey adopted an exchange rate-based stabilization program in December 1999 under support of the IMF. In addition to using exchange rates as a nominal anchor, the program aimed at implementation of comprehensive structural reforms, tight fiscal policies and large-scale privatizations. However, the failure of the government to implement massive structural reforms and privatization has diluted the credibility of the stabilization program, and inflation rate did not fall as envisaged. Fixed devaluation rates coupled with initial high interest rates had increased short-term capital inflows. Domestic currency appreciated sharply, which, in turn, increased trade deficit by more than twice. The banking sector increased its foreign currency denominated liabilities. Deterioration of the current account balance and increasing liability dollarization cast doubts on credibility of the exchange rate-based stabilization program and sudden capital outflow in November 2000 triggered banking crisis. The CBRT used a significant part of its reserves in order to maintain the program. The financial support provided by the IMF under extra reserve facility in November/December 2000 just postponed a new crisis and failed to prevent it.

In the face of speculative attacks, the central bank abandoned crawling peg regime and floated the Turkish Lira in February 2001. Although the exchange rate anchoring was abandoned, the stabilization program was in force thereafter with major modifications. One of the major amendments introduced in the program was amendment of the law on the Central Bank of Republic of Turkey. The central bank was given independence in conduct of monetary policy, and adopted inflation targeting regime. In order to achieve debt sustainability and support the inflation targeting policies of the central bank, fiscal policy was tightened further, increasing the ratio of primary surplus to GDP to 6.5%. In addition, Turkey has implemented massive reforms in order to liberalize the goods and financial markets, and established new regulatory and supervision agencies. As a result of these reforms, Turkey enjoyed relatively high growth and low inflation rates. Annual

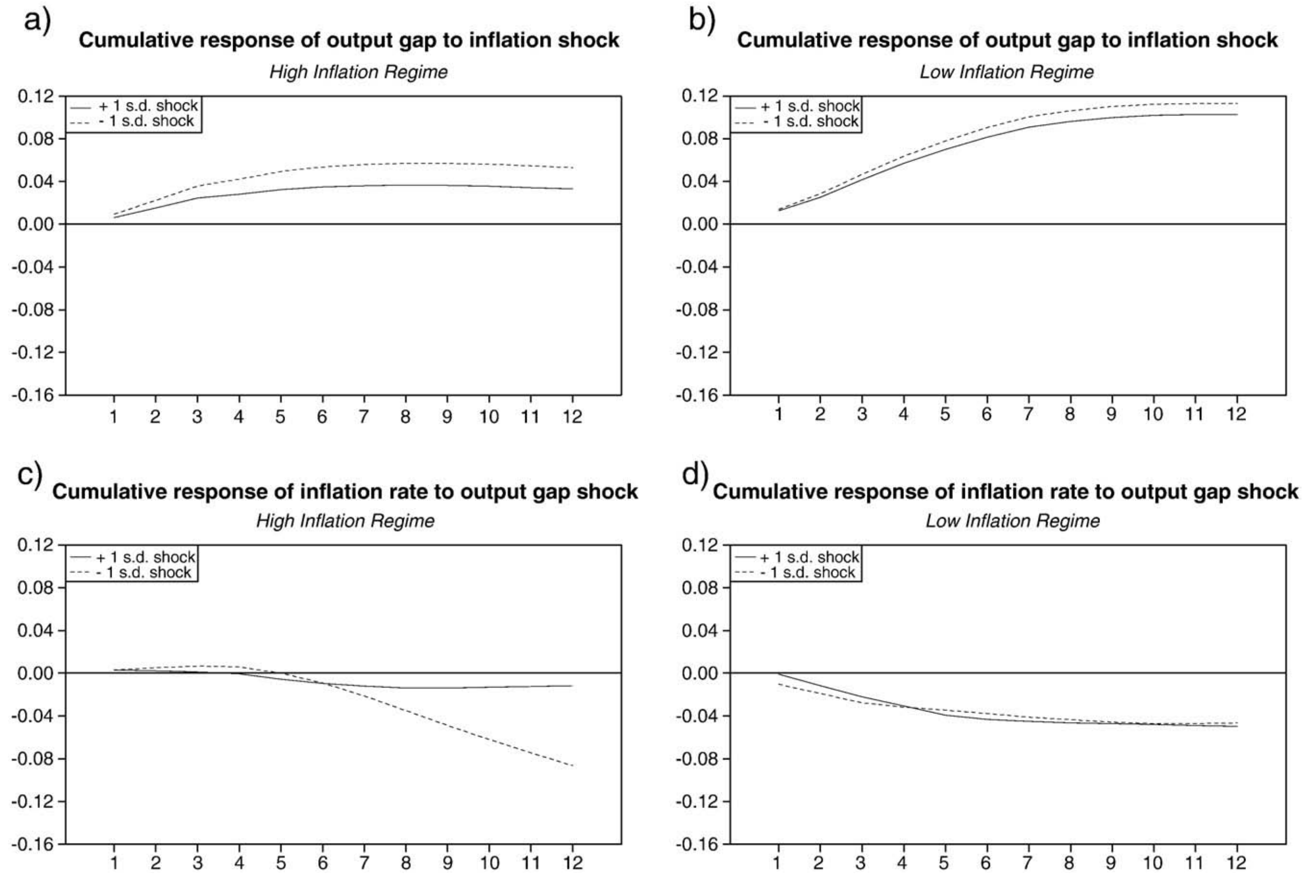

Fig. 4. Cumulative impulse response functions in the sub-period 1998:01–2008:03.

inflation rate reduced to 5.3% in 2005 from 75.3% in 1998. Average annual growth rate of GDP in the post-crisis period was 6.2%.

In Fig. 4 we provide the graphs of GIRF for the third sub-period 1998Q1–2008Q3. The graphs suggest that the effects of a shock to inflation rate had greater effects on output gap in low inflation regimes. Furthermore, the graphs reveal that negative shocks to inflation rate had a greater effect in both regimes, implying that the Phillips curve was convex in this period. The effects of output shocks on inflation rate are also asymmetric though such asymmetry is more pronounced for high inflation regime. In high inflation regime, a shock to output gap had a positive effect on inflation rate during the first four quarters whereas the effect turned to negative starting from the fifth quarter. On the other hand, during low inflation regimes output shocks had a negative effect on inflation rate always. The plotted GIRFs imply that during this sub-period output shocks reduced inflation rate.

The dynamic interrelationship between inflation rate and output gap implied by the GIRFs are consistent with actual Turkish experience, who succeeded to reduce inflation rate considerably and enjoyed relatively high levels of economic growth during the last years. Higher rates of inflation distort efficient allocation of resources by causing reallocation of scarce resources to unproductive activities, and thus reduce potential output. Furthermore, Friedman (1977) argued that a higher inflation rate increases inflation uncertainty and reduces output. Indeed, Telatar and Telatar (2003) find that inflation rate increases inflation uncertainty in the case of Turkey. Therefore, a reduction in inflation rate and inflation uncertainty starting from 2001 supplemented with structural reforms might have increased potential output. In addition the European Union membership perspective of Turkey has intensified foreign capital inflows into Turkey, which add to productive capacity.

Another interesting aspect of the output–inflation relationship in Turkey during the analyzed period is that, as the Figs. 2–4 reveal, the effects of inflation shocks on output in the period 1998 to 2008 is much greater when compared to previous periods. This result may be due to rapid decline of inflation rate and inflationary expectations during the post-crisis period. If past values of inflation rate play a significant role in formation of inflation expectations, then expected inflation rate shall be higher during high inflation periods.[13] The Phillips curve models that incorporate inflation expectations imply that the effects of inflation shocks on output shall be smaller (greater) when economic agents expect that inflation rate shall be high (low) in the future. Therefore, it is not surprising to observe that the effects of inflation shocks were relatively greater during the third sub-period, when Turkey succeeded to reduce inflation rate considerably.

All in all, our results indicate that the relationship between output and inflation is not invariant of economic environment, thus providing evidence in support of the Lucas critique. Particularly, our results suggest that inflation shocks have asymmetric effects on output, consistent with the findings of Telatar and Hasanov (2006). Furthermore, we find that output cost of disinflation is higher in low inflation regimes, implying that policy authorities must be cautious when fighting inflation in high inflation countries like Turkey.

4. Conclusion

In this paper, we investigate possible nonlinearities and instabilities in the inflation–output relationship in Turkey. For this purpose, we first estimate a linear bivariate model for the inflation rate and

[13] Dibooglu (2002) examines the role of inflation expectations on inflation dynamics in the case of Turkey. He finds that inflation dynamics in Turkey are consistent with costless disinflation path. Our results here that output cost of disinflation is considerably low in high inflation periods (especially during the periods from 1981 to 1991 and from 1994 to 1996) are consistent with the findings of Dibooglu (2002).

output gap for the 1980–2008 period using quarterly data, and test for nonlinear alternatives. The test results suggest that the relationship between inflation rate and output gap can be best prescribed by a TV-STR model. The nonlinearity in the TV-STR model is governed by two logistic transition functions. We estimate the TV-STR model employing nonlinear least squares estimators and discuss dynamic relationships between the inflation rate and output gap by using generalized impulse response functions. Our findings can be summarized as follows. First, estimated transition function in inflation rate suggest that the relationship between inflation rate and output gap depends on the initial rate of inflation. Particularly, we find that, the effects of inflation (output) shocks on output (inflation rate) are higher in low inflationary regimes when compared to high inflationary regimes.

Second, we find that, on average, there are asymmetries in the effects of negative inflation shocks on output when compared to the effects of positive shocks, providing evidence in favor of a convex Phillips curve. Although asymmetry in the effects of negative versus positive shocks varies considerably across time, such an asymmetry is more pronounced in low inflation regimes, consistent with the convex Phillips curve.

Third, we find that the dynamic relationship between inflation rate and output gap varies considerably across time, in addition to being regime dependent. As the transition function in time trend evidences, the change in parameters of the estimated nonlinear model happened gradually over time. The structural change started in early 1992, slowed temporarily in mid 1990s, and was over by 1998. This implies that structural change was rather gradually, started after liberalization of the Turkish economy in 1980s, and slowed down temporarily during the economic crisis of 1994. This result indicates that the relationship between output and inflation is not invariant of economic policies, thus providing evidence in support of the Lucas critique.

Fourth, and perhaps most importantly, we find that the shape of the Phillips curve is not independent of economic environment. The results for the first sub-period 1981–1991 suggest that Phillips curve in this period was almost linear. For the second sub-period 1994–1996, on the other hand, our results suggest that Phillips curve was concave. However, the evidence in favor of a convex Phillips curve is more pronounced for the third sub-period 1998–2008. The finding of a convex Phillips curve implies that output cost of disinflation is higher in low inflation regimes when compared to high inflation regimes, suggesting that policy authorities must be cautious when fighting inflation in high inflation countries like Turkey.

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