Monetary Policy Rules in Practice: Evidence from Turkey and Israel

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Abstract

We estimate forward looking monetary policy rules for Israel and Turkey. When variable inflation targets are taken into consideration, as opposed to the fixed targets used in prior research that use data from developed countries, forward looking Taylor rules seem to provide reasonable description of Central Bank behavior in both countries. In general, it can be said that monetary policy appears to be quite strong in these countries, and especially so in Turkey, when compared with developed countries.

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1. Introduction

Monetary policy rules have been analyzed in considerable volume by researchers both from descriptive and prescriptive perspectives[1]. As stated by Svensson (2003), from a descriptive perspective, research has examined to what extent monetary policy rules proposed in the literature, explain actual Central Bank behavior. For developed countries, this question has been extensively analyzed (see Clarida, Gali and Gertler, 1998, 1999, 2000; Taylor, 1993, 1999; Nelson, 2000; Judd and Rudebusch 1998, among others). However, for developing countries the empirical literature has remained sparse. One reason for this could be the fact that flexible exchange rate system is a relatively new development for most developing countries after long periods of some form of a fixed exchange rate system in place. As stated by Taylor (2000), once an emerging market economy abandons the fixed exchange rate regime (a peg, a currency board, or a common currency – dollarization) the only sound monetary policy alternative is the one based on the trinity (Taylor, 1993) of a flexible exchange rate, an inflation target, and a monetary policy rule. As a result many developing countries have recently started implementing inflation targeting regimes with an accompanying monetary policy rule.

In this paper, we contribute to this empirical literature by considering two developing economies, namely, Turkey and Israel. Our aim is to see how reasonable it is to approximate the behavior of the central banks of these countries by our proposed Taylor rule. Turkey adopted a flexible exchange rate regime from February 2001 and explicitly declared its inflation targets starting from January 2002. Israel, on the other hand adopted inflation targeting in 1992 along with a flexible exchange rate regime.

The paper estimates forward looking monetary policy rules for both Turkey and Israel similar to those in Clarida, Gali and Gertler, (1998, 1999, 2000) (CGG, hereafter) and Taylor (1993). The specifications in CGG (1998, 1999, 2000) are forward looking rules estimated by the Generalized Method of Moments (GMM) with implicit fixed targets over long estimation periods. Although this specification is reasonable for developing countries for which the inflation targets stay almost fixed over years, for the two countries under consideration in this study, it cannot be acceptable because they had variable inflation targets over the period under

[1] This terminology belongs to Svensson (2003).

consideration. Therefore we will modify the CGG formulation to take into account the variable targets. To be able to do this we will use expected inflation data instead of actual inflation as CGG. This expected inflation data is based on surveys conducted by the Central Banks of the two countries[2].

Like CGG, we use GMM for the estimation of our monetary policy rules. However, recently GMM estimators have been severely criticized on the ground that inference based on these estimators is inconclusive. The related econometric literature indicates that there has been considerable evidence that asymptotic normality provides a poor approximation to the sampling distributions of GMM estimators. Particularly, the 2-stage least square (2SLS) estimator becomes heavily biased (in the same direction as the ordinary least squares estimator), and the distribution of the 2SLS estimator is quite far from the normal distribution (e.g. bimodal). Stock and Wright (2000) attribute this problem to "weak identification" or "weak instruments", that is, instruments that are only weakly correlated with the included endogenous variables. Stock, Wright and Yogo (2002) and Dufour (2003) provide a comprehensive survey on weak identification in GMM estimation. We also address this issue in our estimators by using recently developed statistics that are immune to weak identification.

The rest of the paper is organized as follows: Section 2 illustrates our modified specification developed to take into account the variables targets in the usual backward looking policy rules. Section 3 describes the data and provides the estimation results. Section 4 concludes.

2. The model

Consider the following monetary policy rule of the form (Taylor, 1993, CGG, 1998, 1999, 2000)

$$i_t^* = r + \pi_{t+k}^* + \beta\left[E\left(\pi_{t+k}/\Omega_t\right) - \pi_{t+k}^*\right] + \gamma E\left(x_t/\Omega_t\right) \tag{1}$$

[2] See Clarida, Gali, and Gertler (1999) Berument and Tasci (2004) estimates monetary policy rules for Turkey over the period from 1990-2001. They adopt the framework of CGG. This approach can be criticized on two grounds. As stated above, targets are treated as fixed and inflation targeting was adopted as a policy at the beginning of 2001.

where i_t^* denotes the target rate for nominal interest rate in period *t*. *E* is the expectation operator, and Ω_t is the information set at the time the interest rate is set. π_{t+k} denotes the percent change in the price level between periods *t* and *t+k* (expressed in annual rates). π_{t+k}^* is the (variable) target for inflation for period *t+k* formed at period *t*. x_t is a measure of the average output gap, with the output gap being defined as the percent deviation between actual output and the corresponding target (potential). *r* is the long-run equilibrium real rate[3]. As in Clarida, Gali, and Gertler (2000), we assume that the real rate is stationary and is determined by non-monetary factors in the long-run.

The policy rule given by (1) has been proven to be useful in both theoretical and empirical grounds. It has been used extensively in empirical research on developed countries as noted above. Approximate forms of this rule are optimal for a central bank that has a quadratic loss function in deviations of inflation and output from their respective targets, given a generic macroeconomic model with price inertia (see, Clarida, Gali, and Gertler 1999, Svensson , 2003). Moreover, as is shown by Clarida, Gali and Gertler (2001, 2002) this policy rule has the same form as in (1) for a small open economy, with possibly different coefficients than those of a closed economy[4]. However, all of these and previously mentioned studies consider a fixed inflation target over the period of estimation, which makes sense for developed countries, but has little relevance for the experiences of Turkey and Israel with inflation targeting. Therefore, we will modify Equation (1) to estimate it with variable inflation targets. Rearranging Equation (1) yields

$$r_t = r + \beta\left[E\left(\pi_{t+k} / \Omega_t\right) - \pi_{t+k}^*\right] + \gamma x_t + \varepsilon_t \tag{2}$$

where $\varepsilon_t = -\gamma\left[x_t - E\left(x_t / \Omega_t\right)\right] + \mu_t$, $r_t = i_t - \pi_{t+k}^*$, and i_t is the actual nominal interest rate. The term μ_t captures the difference between the desired and the actual nominal interest rate ,

[3] It should be noted that *r* is an "approximate" real rate since forecast horizon for inflation will generally differ from the maturity of the short-term nominal rate used as a monetary policy instrument. As noted by Clarida, Gali and Gertler (2000), in practice the presence of high correlation between short-term rates and at maturities associated with the target horizon (1 year) prevents this from being a problem.

[4] Clarida, Gali, and Gertler (2001, 2002) indicates that openness only affects the magnitude of the coefficients in the policy rule.

i.e. $\mu_t = i_t - i_t^*$ [5]. This difference, as put by Clarida, Gali and Gertler (2000), may result from three facts. Firstly, the specification in Equation (2) assumes an adjustment of the actual overnight rates to its target level, and thus ignores, if any, the Central Banks (CB)'s tendency to smooth changes in interest rates (we will address this issue below). Secondly, it treats all changes in interest rates over time as reflecting the CB's systematic response to economic conditions. Specifically, it does not allow for any randomness in policy actions, other than associated with misforecasts of the economy. Third, it assumes that the CB has perfect control over the interest rates, i.e. it succeeds in keeping them at the desired level (e.g., through open market operations).

Finally let $\mathbf{z}_t$ be a vector of variables within the central bank's information set at time it chooses the interest rate (i.e. $\mathbf{z}_t \in \Omega_t$) that are orthogonal to ε_t. Possible elements of $\mathbf{z}_t$ include any lagged variables that help to forecast output gap, as well as any contemporaneous variables that are uncorrelated with the current interest rate shock μ_t. Then since $E\left[\varepsilon_t / \mathbf{z}_t\right] = 0$, Equation (2) implies the following orthogonality condition

$$E\left[r_t - r + \beta\left[E\left(\pi_{t+k} / \Omega_t\right) - \pi_{t+k}^*\right] + \gamma x_t / \mathbf{z}_t\right] = 0 \tag{3}$$

By using this orthogonality condition we use GMM to estimate the parameter vector $\left[r, \beta, \gamma\right]$ as in Clarida, Gali and Gertler (1998, 2000).

CBs may have a tendency to smooth changes in interest rate. Interest rate smoothing can be introduced into the model via the following partial adjustment mechanism (see Clarida, Gali, and Gertler, 1998, 2000)

$$r_t = \left(1-\rho\right) r_t^* + \rho r_{t-1} + v_t \tag{4}$$

where the parameter $\rho \in \left[0,1\right]$ captures the degree of interest rate smoothing. Equation (4) postulates that each period the CB adjusts the funds rate to eliminate a fraction $\left(1-\rho\right)$ of the

[5] We assume that μ_t is identically and independently distributed.

gap between its current target level and past value. ν_t is a independently and identically distributed error term. Substituting (2) into (4) and rearranging yields

$$r_t = (1-\rho)\left\{r + \beta\left[E\left(\pi_{t+k}/\Omega_t\right) - \pi^*_{t+k}\right] + \gamma x_t\right\} + \rho r_{t-1} + \xi_t \tag{5}$$

Where $\xi_t = -(1-\rho)\gamma\left[x_t - E\left(x_t/\Omega_t\right)\right] + \nu_t$. Apparently Equation (5) implies an orthogonality condition similar to (5).

3. Data and Time Series Properties

For Turkey, annual year-end CPI inflation targets are incorporated in the disinflation program implemented with the support of the IMF: the targets for 2002, 2003, and 2004 respectively 35 %, 20 %, and 12 % were indicated in the relevant review of the Standby Arrangement during the last quarter of 2001 and also explicitly declared in Central Bank of Turkey (CBT) website. As our inflation measure looks at the next 12 months, we need to find a solution to the discontinuity of these year-end targets, which we do by a linear transformation, whereby the inflation target in the formula falls each month to reach the next year's inflation target at the first month of the next year. For expected inflation we use CBT's expectations surveys that present expected figures for the inflation for next 12 months. We use the simple average of the two surveys published by CBT. The equivalent in Turkey of the money market short term interest rate of the Taylor rule is the overnight interest rate on borrowing published daily by the CBT: consensus again exists in the markets that this is the relevant indicator of monetary policy. This data is available over the period 2001M8-2004M4.

For Israel we use expected and target inflation rates provided by the Central Bank of Israel (CBI) over the period 1999M1-2002M12. Treasury bill rate is used as a proxy for the money market short term interest rate. This later variable is obtained from International Financial Statistics (IFS) CDROM published International Monetary Fund.

Measuring the output gap is a routine calculation in developed economies with low output volatility and smooth (relatively slow) changes in the structure of output and the composition of domestic demand. Especially for Turkey this variable raises several serious issues, further amplified by crisis conditions within the period of observation. The ideal measure should be based, in our opinion, on the comparison of the real domestic demand excluding the change in inventories with its long-term trend value, for which data is only published on a quarterly basis with a lag of three months. Therefore, we use the seasonally adjusted industrial production series (IPS) and use the definition of Khalaf and Kichian (2004) for the measure of output gap. That is, rather than detrending the log of IPS using the full sample, T, we proceed iteratively: to obtain the value of the gap at time t, we detrend IPS with the data ending in T. We then extend the sample by one more observation and re-estimate the trend. This is used to detrend IPS and yields a value for the gap at time t+1. This process is repeated until the end of the sample. In this fashion, our gap measures at time T do not use information beyond that period and can therefore be used as valid instruments. To detrend IPS. Hodrick-Prescot (HP) filter. IPS series for both countries are obtained from IFS CDROM.

Since the econometric estimation procedure that we use here (GMM) requires that all the variables (including instruments) used in the estimation should be stationary, all of the variables are tested by using the Augmented Dickey-Fuller (ADF) tests and we find that the null of unit root is rejected in all variables, at least at the 10 percent significance level, when tests are applied at different lags[6]

4. Estimation

In this section we will perform a linear GMM estimation to obtain the estimators of the parameters of Equation (2). The GMM estimator that we use is Limited Information Maximum Likelihood (LIML)[7]. The results are illustrated in Table 1 for Turkey and Israel. The instruments we use for GMM estimation consist of one lag of output gap and and one lag of output growth, both for Turkey and Israel [8].

Table 1 is about here

The first three columns of Table 1 and 2 report the estimates of r, β, and γ. The estimate of the coefficient on difference between expected and targeted inflation is around 0.95 for

[6] These results are available upon request.
[7] The usual Two Stage Least Square estimators yield exactly the same results. For LIML estimators we used the GAUSS code originally used by Stock and Wright (2000).
[8] By choosing these instruments, we implicitly assume that these two variables are strong instruments for predicting output gap.

Turkey and 0.73 for Israel. That is, if expected inflation were 1 percentage point above the target, the Central Banks (CB) would set the real interest rate 95 and 73 basis point above its equilibrium value. This coefficient also appears to be highly significant for both countries when we use asymptotic normality as an approximation to the sampling istribution of GMM estimators. The response of the CBs to the deviations of expected output gap from its target (assumed to be zero) is around 0.54 for Turkey and is 0.22 for Israel. In other words holding the difference between expected and targeted inflation constant one percent increase in output gap induces the CBs to increase the real rates by 54 and 22 percent in Turkey and Israel respectively. These coefficients are also statistically significant at 5 percent significance (though not at 1 percent) level. Another difference in the two countries lies in estimated coefficient of the equilibrium real interest rate. This is estimated as 13 and 6 percent for Turkey and Israel respectively. They are also highly significant using normal asymptotics. These estimates indicate the fact that the CB of Turkey implemented much stronger monetary policy in the period of observation than that of the CB of Israel.

The Hansen's *J*-statistic reported in Table 1, does not reject the null hypothesis that the overidentifying restrictions are satisfied at conventional significance levels for Turkey. However, for Israel *J*-statistic indicates that overidentfying restrictions are not satisfied.

Despite their significance (or insignificance), as we mention in the introduction, one should wary about GMM-based results that are obtained under the asymptotic normality of the sampling distributions that obtained under conventional asymptotics. However, under weak-identification asymptotics, the sampling distributions are quite far from being normally distributed. In this paper we address the problem of weak identification by using Anderson and Roubin (1949), test (AR test) in its general form developed by Dufour and Jasiak (2001) (see also Dufour 2003). This test is robust in the case of nonlinear models (see Dufour, 2003; Stock, Wright and Yogo, 2002), and perhaps, more importantly, are even robust to excluded instruments (see Dufour, 2003). Since it is rarely possible to use all possible instruments, this latter property is quite important from applied point of view.

AR test statistic is used to test the null hypothesis that (for Turkey's estimated parameters), $H_0 : r = 12.6725; \beta = 0.9409; \gamma = 0.5383$, i.e. given the instruments that we used, whether the estimated parameters of Equation (3) are compatible with the data or not. Since the test is fully robust to weak instruments (see Stock, Wright, and Yogo, 2002, pp.522), a non-rejection of this null hypothesis means that our estimates are also "data-admissible" even under the case of weak instruments.

The AR-statistics, under the above null hypothesis has an exact Fisher distribution with *k* and *T-k-x* degrees of freedom (where *k* is the number of instruments, *x* is the number of exogenous variables, and *T* is the number of observations), given that the error terms are i.i.d. normal and the instruments are strictly exogenous. $k \times AR$ statistics are asymptotically distributed chi-square with *k* degrees of freedom, even without i.i.d. normal errors under standard regularity conditions (see Dufour and Jasiak, 2001, pp. 829, and Dufour 2003, pp.20). As can be followed from Table 1 given the high *p*-value of the *AR*-test, our parameter estimates cannot be rejected. In other words our GMM estimates of monetary policy rules given Equation (3), cannot be refuted by neither the Turkish nor Israelian data.

Table 2 summarizes the estimation results of Equation (5), i.e. when interest rate smoothing is introduced into the equation into Equation (3). The instrument set for this estimation contains first and second lags of output gap and growth[9].

Table 2 is about here

The estimation results indicate that smoothing parameters for both countries are highly significant, similar to the other parameters, and their magnitudes equal to 0.51 (Turkey) and 0.78 (Israel). These estimates imply that although Israel's CB allows less variability in its interest rate, both countries' CBs put significant efforts for smoothing interest rate. The inclusion of the interest rate smoothing parameter also leads to some significant changes in the estimated parameters of the previous equation. When this parameter is included, the only negligible change occurs in *r* (the estimated value of this coefficient moves from 12.7 to 11.1 for Turkey, from 5.9 to 6.1 for Israel. However, there are considerable changes in the estimated of the remaining parameters especially in β, and to a lesser extent in γ (for Turkey only). When interest rate smoothing is present in the model, the estimate of the coefficient on difference between expected and targeted inflation is considerably higher than in the absence

[9] In this case the estimation method is nonlinear LIML with continuous updating algorithm. We again used the GAUSS code of Stock and Wright (2000). The asymptotic standard errors of the estimators are calculated by using delta method.

of smoothing. This coefficient is now estimated as 1.19 for Turkey and 1.35 for Israel. That is, if expected inflation were 1 percentage point above the target, the CBs would set the real interest rate 1.19 and 1.35 basis point above its equilibrium value. The response of the CBs to the deviations of expected output gap from its target is now around 0.73 for Turkey, i.e., with smoothing Turkish CB appears to be more reactive to the deviations in the output gap. There is not much significant change in the estimated value of this variable compared to the other case.

Similar to the above case J-Statistics indicates that the validity of our instruments is not rejected by the data and AR statistics confirms the validity of the model even in the case of weak instruments.

5. Conclusion

When variable targets are taken into account, forward looking Taylor rules seem to provide reasonable description of CBs behavior, in both Turkey and Israel, even with only two response variables such as deviation from targets and output gap. It should be pointed out that we also include some other variables in Taylor rules, such as money growth, real exchange rate, deviation of the real exchange from an “equilibrium” level, nominal exchange rate growth. None of these variables turn out to be significant in these countries. In general, it can be said that monetary policy appears to be quite strong, especially in Turkey, in these countries when compared with the developed countries’ policy function as outlined in the above mentioned papers.

Tables

Table 1
Results of Equation (3) for Turkey and Israel

Country	r	β	γ	*AR-stat*		*J-stat*	*Sample size (n)*
				$F(2;29)$	$\chi^2(2)$	$\chi^2(1)$	
Turkey	12.6725 (1.1839) [0.000]	0.9409 (0.1518) [0.000]	0.5383 (0.2833) [0.0287]	0.8512 [0.4373]	1.7024 [0.4269]	0.8931 [0.3446]	33
				$F(2;44)$	$\chi^2(2)$	$\chi^2(1)$	
Israel	5.9550 (0.1952) [0.000]	0.7365 (0.1303) [0.000]	0.2172 (0.0494) [0.000]	0.8507 [0.434]	1.7015 [0.4271]	10.8609 [0.001]	48

Notes: *Standard errors are in parantheses and p-values are in brackets.*

Table 2

Results of Equation (5) for Turkey and Israel

Country	r	β	γ	ρ	*AR-stat*		*J-stat*	*Sample size (n)*
					$F(4;25)$	$\chi^2(4)$	$\chi^2(3)$	
Turkey	11.0984 (1.9625) [0.000]	1.1937 (0.2759) [0.000]	0.7337 (0.3439) [0.0164]	0.5073 (0.1458) [0.0003]	0.9387 [0.4578]	3.7547 [0.4402]	4.4106 [0.2204]	32
					$F(4;40)$	$\chi^2(4)$	$\chi^2(3)$	
Israel	6.1303 (0.2717) [0.000]	1.3591 (0.2346) [0.000]	0.2237 (0.0650) [0.0003]	0.7782 (0.0390) [0.000]	1.0490 [0.3942]	4.1962 [0.3801]	8.7991 [0.0321]	47

Notes: *Standard errors are in parantheses and p-values are in brackets.*

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