# Inflation Targeting matters! - Novel evidence from 'ex ante' Taylor rules in emerging markets

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Proponents of inflation targeting argue that such a strategy directly influences expectation formation processes in financial markets. This paper provides a novel test for the evidence that financial market expectations are formed differently under inflation targeting regimes. Using forecasts for the short-term interest rate, the inflation rate, and output growth for ten emerging markets in Latin-America, central and eastern Europe out of which six economies are inflation targeting economies we estimate expected Taylor-type rules. We find evidence for differences in the expectation formation process in the sense that the well-known Taylor principle fairly holds for countries which adopt an inflation targeting system, while for the other countries it does not.

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**JEL classification:** E52, D84, C33

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# 1 Introduction

In the 1980s Latin-American countries experienced the highest inflation of all countries of above 200 percent per year. In contrast, in 2006 they had an average inflation rate of about 6 percent. A similar process of declining inflation rates happened in many central and eastern European countries during the 1990s. As a group, these countries reduced their inflation rates substantially from, on average, 45 percent per year in the 1990s down to, on average, 5 percent per year in 2007. This process of stabilizing prices was achieved in individual countries under fairly different monetary and exchange rate regimes, ranging from the adoption of inflation targeting combined with floating exchange rates to the abandonment of independent monetary policy by introducing currency boards or even by dollarization of the economy.

Inflation targeting was adopted, for instance, by Brazil, Chile, the Czech Republic, Hungary, Mexico, Poland, and Slovakia. This raises the question: Does inflation targeting matter for the performance of the economy? Many empirical studies have focused on this issue. However, most studies have so far concentrated on the impact of inflation targeting on variables that are directly observable like output, unemployment and inflation. While these studies clearly show that the adoption of inflation targeting has significantly reduced inflation, the literature has not yet found a consensus whether the reduction of inflation is accompanied by a change in output growth.[1]

[1]While Mishkin and Schmidt-Hebbel (2006) find that inflation targeting helps countries to achieve lower inflation rates in the medium-term. Their evidence does not suggest that countries that adopt inflation targeting attain better monetary policy performance relative to non-inflation targeting countries. By analyzing 15 inflation targeting countries Landerrechte et al. (2001) find that output growth suffers in inflation targeting regimes. Compared to that Brito and Bysted (2006) show that inflation targeting is an efficient monetary policy to decrease the level and volatility of inflation in 13 Latin-American countries, but that the adoption of inflation targeting is not accompanied by a change in output growth or higher volatility in interest rates. Corbo and Schmidt-Hebbel (2001) compare inflation targeting in five Latin-American countries to other inflation targeting countries. They conclude that the adoption of inflation targeting is correlated with a large decline in output volatility. Finally, Corbo et al. (2002) find that inflation targeting countries have succeeded in reducing output costs of stabilization and in strengthening

The fact that most empirical studies dealing with the effects of adopting a regime of inflation targeting focus on actual economic performance of inflation targeting countries means that there is a gap in empirical research, because theoretically it is often argued that - in the first place - inflation targeting has important consequences for the expectation formation process which, in turn, leads to different economic performances. This paper contributes to close this gap in research by examining whether the introduction of inflation targeting has systematically changed the expectation formation process in financial markets. Since the academic literature has established a close theoretical link between inflation targeting and Taylor rules, we evaluate the performance of a group of ten (inflation targeting and non-inflation targeting) emerging markets by the means of the expected Taylor rules, i.e. 'ex-ante ' Taylor rules. More precisely, we investigate whether financial markets expect the central bank to behave in a manner in line with the well-known Taylor rule.

This paper, thus, changes the perspective on interest rate rules from the typical use in the academic literature as a ex-post reaction function to explaining central bank behavior. We use data from the Consensus Economic Forecast poll and examine whether 'ex-ante ' Taylor rules are present in the expectation formation process for emerging market variables. The data set includes Argentina, Brazil, Chile, Czech Republic, Hungary, Mexico, Poland, Slovakia, Turkey, and Venezuela. Since this country group includes inflation targeting as well as non-inflation targeting countries, we are able to test whether the adoption of inflation targeting matters for expectation formation process.

In order to do so, this paper is structured as follows: The subsequent section briefly reviews the theoretical link between inflation targeting and interest rate rules as well as the commonly applied empirical concept of

policy credibility.

Taylor-type rules. Section 3 explains the data set. Sections 4 and 5 present the empirical results for inflation targeting and non-inflation targeting countries. Subsequently section 6 studies the group of inflation targetors in more detail by looking at the importance of the time-varying inflation target and the credibility issue. Finally, section 7 concludes.

## 2 The theoretical and empirical morphology of Taylor-type rules

All major central banks in industrial countries currently conduct monetary policy by using market-oriented instruments in order to influence the short-term interest rate. Since the seminal paper of Taylor (1993), it has virtually become a convention to describe the interest rate setting behavior of central banks in terms of monetary policy reaction functions. In its plain form, the so-called Taylor rule states that the short-term interest rate, i.e., the instrument of a central bank, reacts to deviations of inflation and output from their respective target levels. Although the Taylor rule started out as an empirical exercise, there is a clear theoretical link between inflation targeting and Taylor rules. Svensson (1997) showed that a Taylor rule can be derived as the explicit solution of an optimal control problem within a stylized macro model which we briefly review subsequently.

Aggregate demand is described by the conventional dynamic IS-relation of the following form:

$$\hat{y}_{t+1} = \beta_1 \hat{y}_t - \beta_2 (i_t - \pi_t) + \eta_{t+1} \tag{1}$$

where $\hat{y}$ is the output gap defined as the deviation of real GDP from its natural level, $i$ is the short-term nominal interest rate that is also the instrument of the central bank and $\pi = p_t - p_{t-1}$ is the annual inflation rate with $p$ representing the aggregate price level. All variables except the interest rate are expressed in natural logarithms. The term $\eta$ is an i.i.d. shock variable with zero mean representing demand shocks. The structural parameters are such that $\beta_2 > 0$ and $0 < \beta_1 < 1$.

Aggregate supply is expressed in terms of the Phillips curve relation:

$$\pi_{t+1} = \pi_t + \alpha_1 \hat{y}_t + \epsilon_{t+1} \tag{2}$$

where $\epsilon$ represents an i.i.d. random cost-push shock, and the parameter $\alpha_1 > 0$ determines the slope of the short-term Phillips curve. Equation (2) states that inflation changes according to the size of the output gap and the supply shocks.

The central bank aims at minimizing the following intertemporal loss function with the size of inflation and the output gap as the two arguments:

$$\min E_t \sum_{\tau=t}^{\infty} = \delta^{\tau-t} L(\pi_\tau, \hat{y}_\tau), \tag{3}$$

where $0 < \delta < 1$ represents the discount factor. The period loss function is specified as:

$$L(\pi_\tau, \hat{y}_\tau) = \frac{1}{2}[(\pi_\tau - \pi^*)^2 + \lambda \hat{y}_\tau^2] \tag{4}$$

where $\pi^*$ is the inflation target defined by the monetary authority and $\lambda$ is the relative weight that is attached to stabilize output. For $\lambda > 0$ such preferences are usually described as *variable inflation targeting*, whereas a zero weight on output expresses a strategy of *strict inflation targeting*.

Optimizing the intertemporal loss function under the constraints of the structure of the economy displayed by the IS- and the Phillips curve leads to an optimal behavior that is commonly characterized as *inflation-forecast targeting*:

$$\pi_{t+2|t} - \pi^* = -\frac{\lambda}{\delta \alpha_1 k} \hat{y}_{t+1|t}, \tag{5}$$

with

$$k = \frac{1}{2}\left(1 - \frac{\lambda(1-\delta)}{\delta\alpha_1^2} + \sqrt{1 + \frac{\lambda(1-\delta)}{\delta\alpha_1^2} + \frac{4\lambda}{\alpha_1^2}}\right) \geq 1,$$

where $x_{t+j|t}$ denotes the expectations in time $t$ of variable $x$ for $j$ periods ahead $t$.[2] The two-period-ahead inflation forecast, $\pi_{t+j|t}$, should equal the inflation target only if the one-period-ahead output forecast equals the natural output rate, that is, if the expected next period's output gap, $\hat{y}_{t+1|t}$, is zero. Otherwise it should exceed (fall short of) the inflation target in proportion to how much the one-period output forecast falls short of (exceeds) the natural level of output. The proportionality is increasing in the relative weight $\lambda$. This essentially displays the gradual inflation stabilization strategy under variable inflation targeting. A higher weight on output stabilization leads to a slower adjustment of the inflation rate.

Solving the demand and supply equations for the respective expectations and substituting them into the optimality condition (5) yields a specific reaction function of the Taylor rule form:

$$i_t = \pi_t + \tilde{b}_1(\pi_t - \pi^*) + \tilde{b}_2 \hat{y}_t, \tag{6}$$

with

$$\tilde{b}_1 = \frac{1-c}{\beta_2 \alpha_1},\ \tilde{b}_2 \ = \frac{1-c+\beta_1}{\beta_2} \text{ and } c = \tfrac{\lambda}{\lambda + \delta \alpha_1^2 k}$$

For the purpose of empirical exercises Clarida et al. (1998) propose a forward-looking variant of the Taylor rule which takes into account the preemptive nature of monetary policy as well as interest smoothing behavior of central banks. This particular type of reaction function has become very popular in applied empirical research. Although it is still in the spirit of the Taylor rule, specifications of this type represent a modification of the original Taylor rule and, thus, the literature often refers to them as Taylor-type rules. Following Clarida et al. (1998, 2000) and Taylor (1999) the baseline forward-looking policy rule takes the form:

$$i_t^* = \bar{i} + \alpha_1 E_t(\pi_{t+k} - \pi^*) + \alpha_2 E_t(y_{t+k} - y_{t+k}^*), \tag{7}$$

[2] That is, $E_t[x_{t+j}] \equiv x_{t+j|t}$.

where $i^*$ is the desired level of the nominal short-term interest rate, and $\bar{i}$ is its equilibrium level.[3] The second term on the right-hand side is the expected deviation of the $k$-period ahead inflation rate ($\pi$) from the target rate ($\pi^*$) which is assumed to be constant over time.[4] The third term is the expected deviation of the $k$-period ahead level of output ($y$) from its natural level ($y^*$) (i.e., the expected output gap $E[\hat{y}_t]$). The coefficients $\alpha_1$ and $\alpha_2$ which will be the center of our estimates represent the reaction coefficients.[5] The additional assumption of interest rate smoothing behavior implies that:

$$i_t = (1-\rho)i_t^* + \rho i_{t-1} + \nu_t, \tag{8}$$

with the parameter $\rho$ (with $0 < \rho < 1$) representing the degree of interest rate smoothing and $\nu_t$ represents an i.i.d. exogenous random shock to the interest rate. Combining equations (7) and (8) lead to

$$i_t = (1-\rho)(\bar{i} + \alpha_1 E_t(\pi_{t+k} - \pi^*) + \alpha_2 E_t(y_{t+k} - y_{t+k}^*)) + \rho i_{t-1} + \nu_t \tag{9}$$

Equation (9) represents the econometric specification which is commonly used to describe central bank behavior.[6] It is reduced to the plain Taylor rule when $\rho$ is assumed to be zero and the horizon of the forward-looking behavior of the central bank, $k$, is also set equal to zero in econometric exercises.

The main messages generated by empirical studies focusing on central bank behavior can be summarized as follows. First, forward-looking specifi-

[3]The difference in the first term on the right hand side of equations (6) and (7) arises from the fact that the model of Svensson (1997) normalizes the equilibrium real interest rate to zero.

[4]In the subsequent analysis we allow the inflation target $\pi^*$ to be time variant. This actually fits reality very well against the background that inflation targeting countries frequently announce inflation targets reflecting nothing else than the desired long-term inflation rate. As these countries are trying to decrease the perceived long-term inflation level they are announcing decreasing inflation targets.

[5]We changed the notation to indicate that in our empirical exercises we do not estimate the optimal rule that we derived before.

[6]Since it contains expectations on the right-hand side that are not directly observable it is common to substitute them by the observed ex-post levels of the respective variables and rearrange the estimation equation into a form that contains the expectation errors of the central bank in the error term. This form is then estimated based on the General Methods of Moments.

cations seem to fit the central banks' behavior better than contemporaneous versions. Here the forward-looking feature is most relevant for the inflation gap with the horizon ($k$) being about one year. Second, the relevance of the Taylor principle for stability, which is a reaction coefficient for inflation being greater than unity, is well demonstrated and its presence is a strong feature for most central banks. Third, the reaction coefficient for the output gap is mostly significant but has a significant lower level compared to the inflation gap coefficient.[7] Fourth, persistence in the short-term interest rate is a strong feature in the data. However, what is not yet clear is whether this is due to intended interest rate smoothing or whether it is due to a strong autocorrelation in the shocks upon which monetary policy reacts.[8]

Our empirical analysis takes the afore-mentioned four empirical core results of Taylor-type rules as its starting point and interprets them as (historical) information that is available for financial market participants. We also assume that the agents in the financial market are aware of the theoretical link between inflation targeting and Taylor rules, that is that they in particular expect inflation targeting countries to be well described by Taylor-type rules. If, in turn, the agents believe in the Taylor-type rules and take this kind of analysis seriously, we would expect to observe this in their joint forecasts for the short-term interest rate, the inflation rate and the output development. In this case, the joint forecasts of the three variables can hardly be independent from each other. They should rather display the same links and dependencies that the estimated reaction functions of the central banks tell us. We refer to them as 'ex-ante' Taylor rules. In addition, because of the theoretical link between inflation targeting and Taylor rules, this form of

[7] In particular, for the output gap the literature demonstrated that it is relevant to discriminate between ex post and real-time data (Orphanides, 2001). Since we use observed expected variables in our analysis we do not need to take effects arising from ex post data into consideration.

[8] Again, since this issue is also not of a strong concern in the present paper, we refer to the recent literature. See, for instance, Rudebusch (2006).

expectation formation should be more relevant for inflation targeting countries compared to non-inflation targeting countries. We therefore estimate variants of equation (9) based on reported forecasts of financial market participants to support the claim that we raised in the title. Before we present the results, the next section briefly introduces our data set.

## 3 The data set

We use data from a survey conducted by Consensus Economics. The survey regularly asks professional forecasters about their projection of several financial and real economy variables such as short-term interest rates, unemployment rates and real GDP forecasts. The survey includes data for several countries. Since our analysis requires forecasts for short-term interest rates and due to data availability, our data set is limited to ten emerging markets, namely Argentina, Brazil, Chile, the Czech Republic, Hungary, Mexico, Poland, Slovakia, Turkey, and Venezuela. Out of this group, seven economies officially adopted inflation targets: Brazil, Chile, the Czech Republic, Hungary, Mexico, Poland and Slovakia.[9] We refer to them as the inflation targeting countries as opposed to non-inflation targeting countries.

This data set has several advantages over other surveys and is, thus, less subject to some of the weaknesses often associated with survey data. First, the individual forecasts are published together with the name of the employer of the forecaster. Given that this allows everybody to evaluate the performance of the individual participants, the accuracy of the forecasts can be expected to have an effect on the reputation of the forecasters.[10] Since an-

[9]Since Slovakia introduced the inflation targeting system as of 2007 we, however, treat Slovakia as a non-inflation targeting country in our study. This can be justified since the time period being a non-inflation targeting country prevails the sample period.

[10]Batchelor (2001) shows that the Consensus Economics forecasts are less biased and more accurate in terms of mean absolute error and root mean square error compared to OECD and IMF forecasts. He also shows that there is little information in the OECD and IMF forecasts that could be used to reduce significantly the error in the private sector forecasts. Mitchell and Pearce (2007) analyze individual forecasts of Wall Street Journal

alysts are bound in their survey answers by their recommendations to clients an analyst may find it hard to justify why he gave a recommendation different to the one in the survey. This all is expected to increase the incentives of the survey participants to submit their best rather than their strategic forecast (see Keane and Runkle, 1990).[11] Second, unlike some other surveys, forecasters participating in the Consensus Economic Forecast poll do not only submit the direction of the expected change of the macroeconomic variable, but forecast a specific level. Third, the survey data are readily available to the public so that our results can easily be verified.

For the five Latin-American countries in our sample the survey provides monthly data for the period from April 2001 to December 2007, hence our analysis covers 81 periods. During this time period 245 institutions participated at least in one survey. For the central and eastern European countries the survey is conducted on a bimonthly basis for the period from May 1998 to May 2007 and onwards on a monthly basis. It includes forecasts of 163 institutions over 63 periods. In order to investigate the time series characteristics of the expectation formation process, we only include professional forecasters participating at least in ten polls.[12] This applies to a total of 128 (116) participants and yields 5,433 (3,722) forecasts for the Latin-American (central and eastern European) countries.

The professional forecasters are requested to predict the economic variables for two different time horizons. The survey provides CPI and real GDP forecasts for the current and next year while the short-term interest is re-

economists'. They find that a majority of the professional forecasters produced unbiased interest rate forecasts, but the forecasts are indistinguishable from a random walk model and the economists are systematically heterogeneously distributed.

[11] In contrast to the view of Keane and Runkle (1990), Laster et. al (1999) develop a model in which forecasters are rewarded for forecast accuracy in statistical terms as well as by publicity in case of giving the best forecast at a single point in time. As a consequence those forecasters will differ the most from the consensus forecast whose wages depend the most on publicity.

[12] We also used other minimum participation rates. The results, however, do not qualitatively change and are available upon request.

quested to be predicted for the next three and twelve months. Hence, this information covers forecast periods of three and twelve months. In order to equalize the beginning and end of the forecast period, we generate a synthetic short-term and medium-term CPI and GDP forecast by weighting the forecast with the remaining months at the time of the forecast. This procedure is quite common in the literature (Heppke-Falk and Hüffner, 2004, and Beck, 2001). For instance, the synthetic medium-term forecast in July is the weighted average of the GDP of the current and next year while the synthetic short-term forecast is, of course, the forecast of the current year only. The Appendix provides an overview of the calculation of the synthetic short-term and medium-term forecasts. Using these alternative time horizons we distinguish between a short-term and a medium-term Taylor rule specification.

Table 1 and 2 summarize the main features of the data set. The predicted interest rate is either the Funds rate or a three-month interest rate. Since the Taylor rule is suggested to work for the Funds rate our data set seems to have a deficiency in case of Hungary, Poland, Slovakia and the Czech Republic as the forecasted series is a three-month interest rate. However, the correlation coefficient between the actual three-month rate and the Funds rate for these countries is about 0.98. Moreover, we potentially would find even stronger evidence on favor of the Taylor rule in financial market expectations if we could observe expectations on the prime rate instead of the three-months interest rate. Hence, this should not diminish the quality of our empirical analysis. Tables 1 and 2 also show that the expectations on the macroeconomic variables are on average a good predictor of their actual value. For instance, the forecasts for the Mexican interest rate (7.8 percent) and inflation rate (4.3 percent) are close to the actual values of 7.7 and 4.4 percent, respectively. Only for Venezuela the forecasts differ from the actual values. While the financial market overestimated both the inflation and the interest rate by about 5 percentage points, their real interest rate forecasts were correct. However, we leave the discussion of the accuracy of the forecasts to further research and turn to the empirical analysis of the expectation formation process in emerging markets.

– Insert Tables 1 and 2 here –

# 4 Estimation results

For our empirical analysis we start from the econometric specification of the Taylor-type rule as derived in section 2:

$$i_t = (1-\rho)(\bar{i} + \alpha_1 E_t(\pi_{t+k} - \pi^*) + \alpha_2 E_t(y_{t+k} - y^*_{t+k})) + \rho i_{t-1} + \nu_t \quad (9)$$

The most difficult variable to quantify in this framework is the expected output gap which is defined as $E_t(\tilde{y}_{t+k}) = E_t(y_{t+k} - y^*_{t+k})$. In line with Clarida et al. (1998), we take the industrial production index and the expected growth rate to measure the expected contribution to the industrial production $E_t(\Delta y_{t+k})$ for the period $t+k$. To calculate the output trend $y^*_{t+k}$ we apply a standard Hodrick–Prescott filter (with the smoothing parameter set at $\lambda = 14{,}400$) and define the expected output gap as $E_t(\tilde{y}_{t+k}) = y_t + E_t(\Delta y_{t+k}) - y^*_{t+k}$.

Next we modify equation (9) in the following way: We use the expected interest rate $E_t i_{t+l}$ as the dependent variable rather than the actual interest rate. Furthermore, in order to arrive at a testable relationship, the unobservable terms in equation (9) have to be eliminated. Since we are able to directly observe the expectations on the short-term interest rate, the inflation rate and the output development, we only lack information on the equilibrium interest rate and the inflation target of the respective central bank. Given the short sample period we treat them as time-invariant for the time being and summarize both in the constant.[13] Thus, we rewrite equation (9) as:

$$E_t i_{t+l} = (1-\rho)\alpha_0 + \alpha_1(1-\rho)E_t\pi_{t+k} + \alpha_2(1-\rho)E_t(\tilde{y}_{t+k}) + \rho i_t + \epsilon_t \quad (10)$$

[13] In the subsequent section 6 we allow for an observable time varying inflation target,

where

$$\alpha_0 = \bar{i} - \alpha_1 E_t \pi^* \tag{11}$$

In equation (10) we already use the expected short-term interest rate forecast as the left-hand side variable. In the subsequent regressions we look at two different forecast horizons. We apply three month forecasts of the short-term interest rate as the left-hand side variable when referring to the short-term forecast. For the medium-term forecast we use the twelve months forecasts of the short-term interest rate as the dependent variable. Note that we do not need to apply the General Methods of Moments when estimating equation (10), since all expectational variables on the right-hand side are also observed data. Thus, we rely on OLS in our panel setting. However, our econometric analysis is impaired by the problem of overlapping forecast horizons since the monthly data set provides three months forecasts. This obviously leads to serial correlation in the error terms by construction. In order to overcome this problem we apply a serial correlation model:

$$\epsilon_{t,i} = \varphi_{t,1}\epsilon_{t-1,i} \tag{12}$$

where the autoregressive term $\varphi$ measures the degree of persistence in the error term. Additionally, we use Prais-Winsten panel corrected standard errors to account for cross section correlation among the survey participants.

Tables 3 and 4 display the estimation results of equation (10). The short-term and medium-term regressions are contemporaneous versions, i.e. all variables enter with the same time index. The short-term equation (called 'Short') regresses the three months interest rate forecast on the forecasts of inflation and output gap for three months (i.e., $l = k = 3$). The medium-term regression (called 'Medium') uses forecast horizons of twelve months

when we limit our analysis to the inflation targeting countries only. At this point of the analysis we do not account for the inflation target because it is not observable for the non-inflation targeting countries and we want to treat both groups identically in our regression specification.

forecasts for all variables instead (i.e., $l = k = 12$). The lagged interest rate is the actual (observable) three months interest rate.[14] In the forward-looking specification (called 'Forward') the dependent variable is the three months interest rate forecasts (i.e., $l = 3$) while the independent variables reflect twelve months forecasts (i.e., $k = 12$). This implies that the monetary policy is expected to affect the inflation rate and GDP growth with a time lag of nine months. Against the background that the time-lag of the monetary policy is about nine to twelve months, the forward-looking specification fits the central bank reaction function very well.

For the Latin-American countries (Table 3) the results show that the expected inflation rate and the expected output gap are indeed significant predictors for the forecasted interest rates. Furthermore, the coefficient of the inflation rate is significantly higher than unity for Brazil, Chile and Mexico as indicated by the $Chi^2$ test in Table 3 which represents the significance level rejecting the null hypothesis that $\alpha_1 \leq 1$. The $Chi^2$ test suggests that the *Taylor principle* only holds (i.e. $\alpha_1 > 1$) in economies which are classified as inflation targeting countries (IT). In the case of Argentina and Venezuela, two non-inflation targeting countries, the inflation coefficients are instead significantly lower from what the *Taylor principle* suggests. Moreover, for Chile the expected output gap coefficients are also in line with our theoretical considerations and of reasonable size. In contrast, for the non-inflation targeting countries and Mexico the output coefficient has the wrong sign which contradicts the Taylor rule.[15] However, a significant negative sign turned out to be a strong feature in the estimation results. Fendel et al. (2008) find

[14]More precisely, in order to avoid daily volatility effects we use the monthly average of the short-term interest rate. However, our results do not qualitatively change using the interest rate at the beginning or the end of the month. Results can be obtained on request.

[15]Interestingly, Schmidt-Hebbel and Werner (2002) report a significant output gap coefficient for Chile estimating the central bank reaction function by the means of equation (10). By contrast they find an insignificant value for the output gap for Brazil which is supported by our results.

the same feature for expectations about the Fed behavior and refer to this as a 'reversed causality'. Forecasters only observe the ex post causality, i.e. that output growth rates react to changes in the official interest rate: If the short-term interest rate is increased output growth slows down. However, forecasters do not base their forecasts on the fact that when the central bank expects or observes higher output growth it tends to increase official interest rates due to the inflationary pressure. The latter can be referred to as the ex ante causality.

The results also hold when we allow for a forward-looking version of the Taylor rule. In the third regression (called 'Forward') we regress the three-month forecast of the short-term interest rate on the twelve-month forecasts of the output gap and the inflation rate (i.e., $l = 3$ and $k = 12$). While the coefficient on the inflation rate is still significantly higher than unity for Brazil, Chile and Mexico, the inflation coefficient is significantly lower in the cases of Argentina and Venezuela. In order to account for the severe financial crisis in Argentina from 1998 - 2002 we also estimate regression (10) for Argentina for the time period January 2003 until December 2007.[16] Yet, the *Taylor principle* remains violated. Besides that, the autoregressive parameter in the error terms is significant and ranges between 0.41 and 0.91. This basically supports our model specification.

Not surprisingly, the results indicate that the predicted interest rate is highly dominated by the current interest rate for the inflation targeting countries as suggested by a significant smoothing parameter ranging around 0.5 to 0.9.[17] Argentina (full sample) and Venezuela experienced low interest

[16]We choose to start the curtailed sample for Argentina from 2003 onwards although we are aware that the aftermaths of the financial crisis still exist. The reason is that the interest rate came down from 50 percent in August 2002 to about 5.7 percent in January 2003 which is close to the post crisis medium-term level of about 4.6 percent.

[17]This result is in line with the stylized fact that inflation expectations are more persistent in inflation targeting countries compared to non-inflation targeting countries (Levin et. al, 2004). This finding also matches the well-demonstrated phenomenon that expectations in financial markets are rather static than dynamic (Mitchell and Pearce, 2007). Furthermore, Krueger and Kuttner (1996) found that the Federal Funds future market

rate persistence ($\rho$ is smaller than 0.3) and hence, reflecting high expected changes in the interest rates during that period.

Table 4 reports the corresponding results for the central and eastern European countries. Again, the results suggest that the Taylor principle only holds (i.e. $\alpha_1 > 1$) in inflation targeting countries (i.e. Poland in the short-term and forward-looking version). For the Czech Republic the *Taylor principle* cannot be rejected in the forward-looking version. Apparently, this supports our previous conclusion that inflation targeting matters for expectation formation: the *Taylor principle* if at all only holds in inflation targeting countries. The Taylor principle does not hold for Hungary, Slovakia and Turkey, where the latter two are the non-inflation targeting countries.

With respect to the output coefficient we find again the wrong or insignificant sing for the two non-inflation targeting countries, namely Slovakia and Turkey while the output gap coefficient is significantly positive in the cases of the Czech Republic and Poland.

In sum, the inspection of Tables 3 and 4 suggests that the Taylor-type rules seem to explain the forecasts very well for the majority of inflation targeting countries while this conclusion does not hold for non-inflation targeting countries. This result is most pronounced for the subsample of Latin-American countries, while it is (so far) less striking for the central and eastern European countries.[18]

– Insert Tables 3 and 4 here –

provide efficient predictions on the future path of the Federal Funds rate. As the future and actual path of the Federal Funds rate are close to each other, static expectations seem reasonable as a means to forecast interest rates. Furthermore, applying actual data Schmidt-Hebbel and Werner (2002) estimate the Taylor rule by means of equation (10) for Brazil, Chile and Mexico. They find smoothing coefficients similar to ours for Brazil, Chile and Mexico of about 0.83, 0.96 and 0.62, respectively.

[18]Note that, so far, we did not control for a possibly time-varying inflation target would could potentially significantly distort our results for the IT countries.

# 5 Expectations on the long-term inflation rate

The estimation procedure enables us to investigate another feature inherent in the Taylor rule. Equation (10) allows us to calculate the long-term inflation rate ($\pi^*$) expected by the financial market. In order to recover the expected inflation target ($\pi^*$) we can use the parameter estimates $\alpha_0$ and $\alpha_1$ from Table 3 and 4. Recall that

$$\alpha_0 = \bar{i} - \alpha_1 E_t \pi^* \tag{11}$$

and given the Fisher relation

$$\bar{i} = i^{real} + E(\pi^*) \tag{13}$$

which together yields

$$\alpha_0 = i^{real} + (1 - \alpha_1) E_t \pi^*. \tag{14}$$

This implies that

$$E_t \pi^* = \frac{\alpha_0 - i^{real}}{1 - \alpha_1}. \tag{15}$$

According to Clarida et al. (1998) we use the expected sample average real interest rate among all individuals to provide an estimate of $i^{real}$ as our sample is sufficiently long. With these estimates it is possible to construct the expected inflation target rate $E_t(\pi^*)$ by the means of the short-term results shown in Table 3 and 4 reflecting the estimation of equation (10).[19]

The expected real interest rate ($i^{real}$), the expected long-term inflation rate $E(\pi^*)$ and the actual inflation rate ($\pi^{act}$) are shown in Table 5 for the

[19] In order to obtain a real-interest rate forecast $i^{real}$ with the same maturity as in Tables 3 and 4 we cannot use the forward-looking version. Therefore, we use the short-term version since Tables 3 and 4 show that the Taylor principle holds more frequent in this version compare to the medium-term version. However, results in the medium-term version are of comparable magnitude and available upon request.

ten emerging markets. The expected long-term inflation rate is the highest for Turkey amounting to 26 percent on an annual basis. Indeed, with a rate of 34 percent annually Turkey experienced the highest inflation rate in our sample. The expected long-term inflation rate is also very close to the actual inflation rate for the majority of countries. The expected long-term inflation rate is not different from the actual (realized) inflation rate for Brazil (6.9 compared to 7.8 percent), Mexico (3.6 compared to 4.4 percent), the Czech Republic (5.1 compared to 3.6), Slovakia (5.4 compared to 6.1 percent) and Turkey (26.4 compared to 33.8 percent). Interestingly for Argentina, the expected long-term inflation rate is not different from the actual inflation rate for the full sample period and for the curtailed period. This implied that the financial market adopts a lower long-term inflation rate for Argentina after the peak of the Argentinean crisis. While for Chile (Venezuela) the financial market slightly underestimated (overestimated) the actual inflation rate, for Hungary the actual inflation rate was twice as high as the expected long-term inflation rate. Before drawing the conclusion that the financial market provides inaccurate expectations one has to recall that the Taylor rule estimated in our analysis is based on expectations which might not coincide with the realized inflation rate. Moreover it seems possible, if not likely, that the financial market learns during the inflation process and the assumption of a constant long-term inflation rate might not be appropriate. Therefore the next section accounts for the possibility of a time-variant long-term inflation rate.

– Insert Table 5 here –

# 6 The role of time-varying inflation targets and credibility

In this section we analyze the link between expected interest rates, projected output development and inflation rate forecasts in more detail by allowing for a time-varying inflation target. As our sample consists of six inflation targeting countries (Brazil, Chile, Mexico, Poland, Hungary and the Czech Republic) which publicly announce inflation targets we now incorporate the time variance of the official inflation target into our analysis. The inflation targets are obtained from the monthly inflation reports of the respective central bank. Figures 1 through 6 show the inflation target (thick solid line), the mean of the three-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line). Not surprisingly, the expected and the actual inflation rate move in line. This feature can be attributed to the fact that the financial market uses the actual inflation rate as a benchmark to build inflation expectations. Figures 1 through 6 also show that the inflation target in the cases of the Latin-American countries is relatively stable compared to the cental and eastern European countries. Notwithstanding, we now take the feature of the time-varying inflation target into account and analyze whether the financial market treats the long-term inflation rate $\pi^*$ to be time-invariant.[20]

– Insert Figures 1 through 6 here –

Starting from equation (9) we now treat the inflation target as observable but time-varying and, thus, do not need to include it in the constant. Therefore we depart from the former specification (10) and estimate[21]

$$E_t i_{t+l} = \alpha_0(1-\rho)+\alpha_1(1-\rho)E_t(\pi_{t+k}-\pi_t^*)+\alpha_2(1-\rho)E_t(\tilde{y}_{t+k})+\rho i_t+\epsilon_t \quad (16)$$

with $\alpha_0 = \bar{i}$.

The estimation results are presented in Table 6. Compared to Tables 3 and 4 (constant implicit inflation target) the results do not noticeably change for

[20] Since the Czech Republic and Hungary introduced the inflation targeting system in June 1999 and December 2001, respectively, we drop all observations prior to this date.

[21] Since we learned from the previous analysis that interest rate smoothing is crucial for the analysis we jointly estimate the smoothing parameter and refrain from the assumption that $\rho = 0$.

Brazil and Chile. This is due to the fact that these countries did not change their targets substantially during the observation period. For the remaining inflation targeting countries, considerable differences between the specifications of the time-constant (Tables 3 and 4) and time-varying inflation targets (Tables 6) emerge. For the Czech Republic and Poland the inflation coefficient increases while only for Poland the output coefficient gains magnitude and significance when the inflation target is included. Although Hungary has been changing the target quite frequently, the results do not change with respect to the market forecasts in the Taylor rule framework.

For Mexico the *Taylor principle* no longer holds as the inflation coefficient is insignificant (short-term and medium-term version) or has even the wrong sign (forward looking version). This means that the financial market does not expect the Central Bank of Mexico to respond to inflation changes in the direction suggested by the Taylor rule including the announced inflation target as the long-term inflation rate. An interpretation of this finding is that financial market participants do not trust the official announced inflation targets in case of Mexico compared to Brazil and Chile. This confirms the results of Schmidt-Hebbel and Werner (2002) analyzing the credibility of the inflation targeting of these three countries. Their results suggest that, in the case of Mexico, a substantial deviation exists between actual and expected inflation, whereas this is not found in the cases of Brazil and Chile. The authors also find substantial deviations of the inflation target from the expected inflation rate which supports our previous finding according to which the incorporation of the inflation target is not being reflected in the expectations of the financial market.

An additional feature for Mexico is revealed when comparing Tables 3 and 6. We find an increase of the smoothing parameter from about 0.5 to 0.7 for the short-term (forward looking) version and from 0.3 to 0.4 for the medium-term version. This means that the financial market again expects the Mexican central bank to change its interest rates to a lesser extent in the specification with the inflation target announcement. Put differently, the persistence of the expected short-term interest rate reflects that the financial market does not expect the central bank of Mexico to respond to inflation changes relative to the inflation target in the specification of the time-varying inflation target, but rather sticks to the current interest rate when predicting the interest rate.

– Insert Table 6 here –

In sum, the financial market does not incorporate the announced inflation target of the central bank of Mexico in forecasts of the short-term interest rate by the means of the Taylor rule. However, the Taylor rule works when we assume that the financial market expects constant a long-term inflation rate that is different to the inflation target. Apparently the central bank of Mexico does not seem to be credible in the eyes of the financial market to reduce the inflation rate. This feature is also reflected in the statement of the central bank of Mexico (2004, p.45) that 'both the level of total core CPI inflation and its expectations suggest that the 3 percent inflation target has still not been included in a widespread fashion into the price determination process.'

By contrast, the announcements of the inflation target of the central banks of Brazil, Chile, Hungary, Poland and the Czech Republic seem to be regarded as being credible as the results do not considerably suffer when including the inflation target. For Chile, Poland and the Czech Republic of the expected Taylor rule even gains fit if time-varying inflation target is included.

# 7 Conclusion

This paper investigates whether the adoption of inflation targeting matters for the expectation formation process in the financial market. Using survey data of the Consensus Economic Forecast poll we analyze whether the financial market predicts short-term interest rates on the basis of the Taylor rule for five Latin-American countries and five central and eastern European countries. We refer to them as 'ex-ante' Taylor rules. We find that for most inflation targeting countries financial markets adopt the Taylor rule framework for their forecasts at least at some time horizon and, furthermore, provide interest rate forecasts close to the realized interest rate. These economies are Brazil, Chile, Mexico, and Poland. Only Hungary and the Czech Republic seem to be exceptions among the group of inflation targeting economies. Compared to that, for non-inflation targeting countries the financial market does not adopt the Taylor rule framework which is in every case reflected in the violation of the *Taylor principle.* We interpret this finding as a strong evidence that inflation targeting matters in the sense that expectations are formed differently in IT regimes like it has been claimed i the IT related literature.

We also included a time-varying inflation target which is the announced inflation target for the inflation targeting countries. The results indicate that for Brazil, Chile, Poland and the Czech Republic the short-term interest rate forecasts are provided by means of the Taylor rule whereas for Mexico the Taylor rule is now violated. Apparently, for Mexico the financial market incorporates a long-term inflation rate that is different to the announced inflation target rate.

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Table 1: Overview of the average monthly forecasts for the Latin-American countries

| Country | Argentina | Brazil | Chile | Mexico | Venezuela |
|---|---|---|---|---|---|
| Introduction of Inflation Targeting | No | Since June 1999 | Since January 1991 | Since January 1999 | No |
| Forecasts | | | | | |
| Interest Rate | Buenos Aires Interbank Offering Rate (BAIBOR) | Funds Rate | Monetary Policy Rate (MPR) | Funds Rate | Funds Rate |
| Short-term | 12.8 | 17.7 | 3.8 | 7.8 | 16.1 |
| Medium-term | 12.2 | 15.6 | 4.4 | 7.7 | 16 |
| GDP Growth | | | | | |
| Short-term | 3.2 | 2.8 | 4.4 | 2.9 | 2.0 |
| Medium-term | 3.5 | 3.0 | 4.5 | 3.2 | 2.7 |
| CPI | | | | | |
| Short-term | 16.4 | 6.5 | 2.9 | 4.3 | 23.7 |
| Medium-term | 12.1 | 6.1 | 2.9 | 4.2 | 23.2 |
| Real Interest Rate | | | | | |
| Short-term | -3.6 | 10.8 | 0.9 | 3.6 | -7.9 |
| Medium-term | -2.4 | 9.2 | 1.6 | 3.6 | -7.8 |
| Actual Series | | | | | |
| CPI | CPI | IPCA | CPI | CPI | CPI |
| Mean | 17.8 | 7.8 | 2.7 | 4.4 | 18.5 |
| Interest Rate | BAIBOR | Funds Rate | MPR | Funds Rate | Interbank Rate |
| Mean | 13 | 18.1 | 3.7 | 7.7 | 11.1 |
| Source | Banco Central de la Republica Argentina | OECD | Banco Central de Chile | Banco de Mexico | Banco Central de Venezuela |
| Period | 2001 - 2007 | 2001 - 2007 | 2001 - 2007 | 2001 - 2007 | 2001 - 2007 |

Notes: Table 1 shows the expected and actual variables under consideration; the real interest rate forecast is the mean of the differences between the forecasts of the nominal interest rate and the inflation rate and is defined as $\bar{E}[i^{real}] = E_t[i] - E_t[\pi]$.

Table 2: Overview of the average monthly forecasts for the central and eastern European countries

| Country | Czech Republic | Hungary | Poland | Slovakia | Turkey |
|---|---|---|---|---|---|
| Introduction of Inflation Targeting | Since January 1999 | Since December 2001 | Since January 1990 | Since January 2007 | No |
| Forecasts | | | | | |
| Interest Rate | 3-months PRIBOR | 90-Day Treasury bill Rate | 3-months Interbank Rate | 3-months BRIBOR | Overnight Interbank Rate |
| Short-term | 4.3 | 9.6 | 9.4 | 7.5 | 35.4 |
| Medium-term | 4.6 | 8.6 | 8.7 | 7 | 28.5 |
| GDP Growth | | | | | |
| Short-term | 3.2 | 3.7 | 4.3 | 4.6 | 3.6 |
| Medium-term | 3.4 | 3.8 | 4.4 | 4.6 | 3.8 |
| CPI | | | | | |
| Short-term | 3.4 | 7.1 | 4.4 | 6.4 | 33.2 |
| Medium-term | 3.7 | 5.6 | 4.2 | 5.3 | 23.9 |
| Real Interest Rate | | | | | |
| Short-term | 0.9 | 2.6 | 5.0 | 1.2 | 3.0 |
| Medium-term | 1.1 | 2.2 | 4.3 | 1.1 | -0.6 |
| Actual | | | | | |
| CPI | CPI | CPI | CPI | CPI | CPI |
| Mean | 3.2 | 7.0 | 3.9 | 6.1 | 31.2 |
| Interest Rate | 3-months PRIBOR | 90-Day Treasury bill Rate | 3-months Interbank Rate | 3-months BRIBOR | Overnight Interbank Rate |
| Mean | 4.4 | 10.3 | 9.9 | 7.7 | 39.3 |
| Source | OECD | OECD | OECD | OECD | OECD |
| Period | 1998 - 2007 | 1998 - 2007 | 1998 - 2007 | 1998 - 2007 | 1998 - 2007 |

Notes: Table 2 shows the expected and actual variables under consideration; the real interest rate forecast is the mean of the differences between the forecasts of the nominal interest rate and the inflation rate and is defined as $\bar{E}[i^{real}] = E_t[i] - E_t[\pi]$.

Table 3: Expected Taylor-type rules for Latin-American countries (April 2001 - December 2007)

| Country | | $\alpha_0$ | $\alpha_1$ | $\alpha_2$ | $\rho$ | $\varphi$ | $\alpha_1 > 1$ | $\alpha_2 > 0$ | $R^2$ | Obs. | Groups |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Short | 3.01* | .61* | .17+ | .17* | .79* | .99 | .05 | .78 | 796 | 27 |
| | | (.92) | (.03) | (.10) | (.02) | | | | | | |
| Argentina | Medium | 2.89* | .58* | -.28+ | .07* | .77* | .99 | .99 | .68 | 498 | 27 |
| | | (.26) | (.04) | (.12) | (.02) | | | | | | |
| | Forward | 1.95 | .80* | .16 | .18* | .82* | .99 | .11 | .75 | 588 | 27 |
| | | (1.29) | (.06) | (.13) | (.02) | | | | | | |
| | Short | 1.41* | .55* | .07 | .46* | .90* | .99 | .06 | .47 | 598 | 22 |
| | | (.11) | (.07) | (.05) | (.05) | | | | | | |
| Argentina | Medium | 2.87* | .48* | -.00 | .23* | .83* | .99 | .50 | .30 | 371 | 22 |
| (w/o crisis) | | (.23) | (.07) | (.06) | (.07) | | | | | | |
| | Forward | .66* | .43* | .01 | .41* | .91* | .99 | .43 | .37 | 409 | 22 |
| | | (.11) | (.08) | (.07) | (.07) | | | | | | |
| | Short | 7.01* | 1.55* | -.03 | .71* | .77* | .00 | .72 | .85 | 1,259 | 30 |
| | | (.43) | (.11) | (.06) | (.02) | | | | | | |
| Brazil (IT) | Medium | 8.44* | .97* | -.03 | .42* | .81* | .30 | .85 | .55 | 1,085 | 30 |
| | | (.31) | (.06) | (.03) | (.03) | | | | | | |
| | Forward | 8.25* | 1.55* | -.05 | .70* | .77* | .00 | .84 | .81 | 1,108 | 30 |
| | | (.52) | (.11) | (.05) | (.03) | | | | | | |
| | Short | -1.18+ | 1.91* | .26+ | .87* | .41* | .00 | .04 | .92 | 1,204 | 25 |
| | | (.59) | (.20) | (.14) | (.01) | | | | | | |
| Chile (IT) | Medium | .00 | 2.09* | .32* | .69* | .64* | .00 | .00 | .82 | 1,009 | 24 |
| | | (.18) | (.17) | (.08) | (.03) | | | | | | |
| | Forward | -1.48+ | 2.02* | .39* | .86* | .41* | .00 | .00 | .91 | 1,050 | 24 |
| | | (.84) | (.29) | (.11) | (.02) | | | | | | |
| | Short | 2.09* | 1.42* | -.11+ | .56* | .64* | .00 | .99 | .73 | 1,253 | 26 |
| | | (.15) | (.10) | (.05) | (.02) | | | | | | |
| Mexico (IT) | Medium | 2.44* | 1.33* | -.07+ | .29* | .75* | .00 | .99 | .54 | 1,091 | 26 |
| | | (.11) | (.09) | (.03) | (.03) | | | | | | |
| | Forward | 1.67* | 1.56* | -.09+ | .54* | .64* | .00 | .99 | .72 | 1,115 | 26 |
| | | (.18) | (.11) | (.04) | (.02) | | | | | | |
| | Short | 11.73* | .18* | .02 | .07* | .86* | .99 | .06 | .49 | 772 | 25 |
| | | (1.08) | (.03) | (.01) | (.02) | | | | | | |
| Venezuela | Medium | 7.32* | .44* | .02 | .08* | .74* | .99 | .50 | .40 | 595 | 25 |
| | | (.38) | (.04) | (.01) | (.03) | | | | | | |
| | Forward | 11.02* | .24* | .02+ | .08* | .85* | .99 | .03 | .42 | 630 | 25 |
| | | (1.32) | (.04) | (.01) | (.03) | | | | | | |

Notes: Estimated equation (10) $E_t i_{t+l} = (1-\rho)\alpha_0 + \alpha_1(1-\rho)E_t\pi_{t+k} + \alpha_2(1-\rho)E_t(\tilde{y}_{t+k}) + \rho i_t + \epsilon_t$ by the means of a serial correlation model where (12) $\epsilon_{t,i} = \varphi_i \epsilon_{t-1,i}$; to estimate Argentina (w/o crisis) we skip the time period before 2003; values in parentheses present panel corrected standard errors applying the Prais-Winsten model; the Hausman test indicates to use the fixed-effects estimator on a one percent significance level; $\alpha_1 > 1$ ($\alpha_2 > 0$) represents the significance level of a $Chi^2$ test to test whether the *Taylor-principle* holds with the null hypothesis that $\alpha_1 \leq 1$ ($\alpha_2 \leq 0$); the $R^2$ refers to the overall coefficient of determination; for readability within and between $R^2$ are left from the Table but available upon request; * (+) indicates significance at the one (ten) percent level, respectively.

Table 4: Expected Taylor-type rules for central and eastern European countries (May 1998 - December 2007)

| Country | | $\alpha_0$ | $\alpha_1$ | $\alpha_2$ | $\rho$ | $\varphi$ | $\alpha_1 > 1$ | $\alpha_2 > 0$ | $R^2$ | Obs. | Groups |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Short | 3.04* | .58* | .06 | .90* | .39* | .99 | .08 | .99 | 823 | 24 |
| | | (.34) | (.09) | (.04) | (.01) | | | | | | |
| Czech Republic (IT) | Medium | 5.00* | .32* | .06+ | .78* | .56* | .99 | .02 | .96 | 763 | 24 |
| | | (.61) | (.12) | (.03) | (.02) | | | | | | |
| | Forward | 1.57* | .94* | .06+ | .89* | .37* | .73 | .03 | .99 | 768 | 24 |
| | | (.42) | (.08) | (.03) | (.01) | | | | | | |
| | Short | 4.74* | .39* | .04 | .84* | .44* | .99 | .32 | .94 | 655 | 24 |
| | | (.55) | (.12) | (.08) | (.02) | | | | | | |
| Hungary (IT) | Medium | 2.48* | .55* | -.11* | .62* | .53* | .99 | .99 | .87 | 574 | 24 |
| | | (.29) | (.09) | (.03) | (.03) | | | | | | |
| | Forward | 3.21* | .84* | -.06+ | .74* | .40* | .99 | .96 | .94 | 576 | 24 |
| | | (.36) | (.08) | (.04) | (.03) | | | | | | |
| | Short | 2.90* | 1.22* | .28* | .87* | .52* | .02 | .00 | .99 | 893 | 29 |
| | | (.32) | (.11) | (.07) | (.01) | | | | | | |
| Poland (IT) | Medium | 4.28* | .40* | .12* | .76* | .54* | .99 | .00 | .96 | 785 | 29 |
| | | (.43) | (.18) | (.04) | (.03) | | | | | | |
| | Forward | 3.06* | 1.32* | .22* | .87* | .51* | .01 | .00 | .98 | 791 | 29 |
| | | (.46) | (.15) | (.06) | (.02) | | | | | | |
| | Short | 3.02* | .66* | -.03 | .79* | .20* | .99 | .62 | .91 | 497 | 18 |
| | | (.77) | (.13) | (.09) | (.02) | | | | | | |
| Slovakia | Medium | 5.17* | -.10 | .04 | .11* | .93* | .99 | .05 | .60 | 427 | 17 |
| | | (.14) | (.08) | (.03) | (.02) | | | | | | |
| | Forward | 5.35* | -.02 | .03 | .29* | .92* | .99 | .20 | .90 | 427 | 17 |
| | | (.17) | (.09) | (.03) | (.02) | | | | | | |
| | Short | 10.19* | .73* | -.35* | .32* | .59* | .99 | .99 | .88 | 699 | 24 |
| | | (1.17) | (.03) | (.10) | (.03) | | | | | | |
| Turkey | Medium | 5.82* | .69* | -.00 | .18* | .69* | .99 | .51 | .77 | 520 | 24 |
| | | (1.58) | (.04) | (.09) | (.04) | | | | | | |
| | Forward | 8.20* | .91* | -.28* | .24* | .56* | .91 | .99 | .89 | 541 | 24 |
| | | (.60) | (.04) | (.09) | (.03) | | | | | | |

Notes: Estimated equation (10) $E_t i_{t+l} = (1-\rho)\alpha_0 + \alpha_1(1-\rho)E_t\pi_{t+k} + \alpha_2(1-\rho)E_t(\tilde{y}_{t+k}) + \rho i_t + \epsilon_t$ by the means of a serial correlation model where (12) $\epsilon_{t,i} = \varphi_i \epsilon_{t-1,i}$; values in parentheses present panel corrected standard errors applying the Prais-Winsten model; the Hausman test indicates to use the fixed-effects estimator on a one percent significance level; $\alpha_1 > 1$ ($\alpha_2 > 0$) represents the significance level of a $Chi^2$ test to test whether the *Taylor-principle* holds with the null hypothesis that $\alpha_1 \leq 1$ ($\alpha_2 \leq 0$); the $R^2$ refers to the overall coefficient of determination; for readability within and between $R^2$ are left from the Table but available upon request; * (+) indicates significance at the one (ten) percent level, respectively.

Table 5: Expected long-term inflation target and actual inflation rate

| | Argentina | w/o crisis | Brazil | Chile | Mexico | Venezuela |
|---|---|---|---|---|---|---|
| Real Interest Rate Forecast: $i^{real}$ | -3.64 | -3.19 | 10.84 | 0.89 | 3.57 | -7.86 |
| Expected Inflation Rate: $E(\pi^*)$ | 17.15 (2.09) | 10.22 (1.66) | 6.90 (1.50) | 2.28 (.20) | 3.57 (.54) | 23.84 (.94) |
| Actual Average Inflation Rate: $\pi^{act}$ | 17.75 | 9.45 | 7.80 | 3.40 | 4.35 | 21.30 |
| Test: $E(\pi^*) = \pi^{act}$ | .78 | .64 | .55 | .00 | .16 | .01 |

| | Czech Rep. | Hungary | Poland | Slovakia | Turkey |
|---|---|---|---|---|---|
| Real Interest Rate Forecast: $i^{real}$ | .92 | 2.60 | 4.97 | 1.18 | 2.97 |
| Expected Inflation Rate: $E(\pi^*)$ | 5.05 (.64) | 3.49 (.33) | 9.19 (3.35) | 5.41 (.91) | 26.39 (2.82) |
| Actual Average Inflation Rate: $\pi^{act}$ | 3.60 | 7.38 | 4.60 | 6.20 | 33.82 |
| Test: $E(\pi^*) = \pi^{act}$ | .03 | .00 | .17 | .38 | .01 |

Notes: The expected real interest rate is the average of the real interest rate forecast over the sample period; the expected inflation rate is calculated by the means of (15) $E_t\pi^* = \frac{\alpha_0 - i^{real}}{1-\alpha_1}$ and based on the estimation results of Tables 3 and 4; in order to estimate Argentina (w/o crisis) we skip the time period before 2003; standard errors in parenthesis; the actual inflation rate reflects the average inflation rate as displayed in Tables 1 and 2; the sources of the actual inflation rate are presented in Tables 1 and 2; the last row reflects the significance level of a two-sided t-test under the null hypothesis that the expected long-term inflation rate equals the actual average inflation rate.

Table 6: Interest rate smoothing and time-varying inflation target

| Country | | $\alpha_0$ | $\alpha_1$ | $\alpha_2$ | $\rho$ | $\varphi$ | $\alpha_1 > 1$ | $\alpha_2 > 0$ | $R^2$ | Obs. | Groups |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Short | 13.50* | 1.64* | -.03 | .79* | .77* | .00 | .67 | .86 | 1,259 | 30 |
| | | (1.01) | (.16) | (.08) | (.02) | | | | | | |
| Brazil | Medium | 12.11* | .87* | -.03 | .52* | .80* | .95 | .83 | .55 | 1,085 | 30 |
| | | (.48) | (.08) | (.04) | (.03) | | | | | | |
| | Forward | 14.73* | 1.64* | -.07 | .78* | .76* | .00 | .83 | .81 | 1,108 | 30 |
| | | (1.11) | (.17) | (.07) | (.02) | | | | | | |
| | Short | 5.15* | 2.32* | .31+ | .90* | .41* | .00 | .05 | .93 | 1,179 | 25 |
| | | (.36) | (.29) | (.19) | (.01) | | | | | | |
| Chile | Medium | 8.38* | 2.21* | .43* | .81* | .64* | .00 | .00 | .82 | 1,009 | 24 |
| | | (.89) | (.37) | (.14) | (.02) | | | | | | |
| | Forward | 5.25* | 2.60* | .51* | .89* | .41* | .00 | .00 | .91 | 1,050 | 24 |
| | | (.41) | (.41) | (.15) | (.02) | | | | | | |
| | Short | 8.44* | .03 | -.17+ | .70* | .61* | .99 | .99 | .72 | 1,253 | 26 |
| | | (.38) | (.14) | (.07) | (.02) | | | | | | |
| Mexico | Medium | 8.24* | -.03 | -.09+ | .44* | .74* | .99 | .99 | .52 | 1,091 | 26 |
| | | (.30) | (.10) | (.04) | (.03) | | | | | | |
| | Forward | 8.76* | -.03 | -.13+ | .70* | .59* | .99 | .97 | .72 | 1,115 | 26 |
| | | (.42) | (.16) | (.07) | (.02) | | | | | | |
| | Short | 6.89* | 1.31* | .07 | .95* | .42* | .24 | .12 | .96 | 523 | 23 |
| | | (1.71) | (.42) | (.06) | (.02) | | | | | | |
| Czech Republic | Medium | 11.57* | .83+ | .10 | .92* | .48* | .36 | .10 | .83 | 443 | 23 |
| | | (4.46) | (.49) | (.07) | (.03) | | | | | | |
| | Forward | 6.18* | 1.42* | .06 | .94* | .41* | .16 | .11 | .96 | 444 | 23 |
| | | (1.23) | (.42) | (.05) | (.02) | | | | | | |
| | Short | 6.71* | .05* | -.02 | .82* | .44* | .99 | .60 | .89 | 436 | 23 |
| | | (.30) | (.12) | (.10) | (.02) | | | | | | |
| Hungary | Medium | 5.59* | .12 | -.05 | .61* | .60* | .99 | .81 | .68 | 355 | 22 |
| | | (.22) | (.12) | (.06) | (.03) | | | | | | |
| | Forward | 6.52* | .69* | -.08 | .76* | .37* | .99 | .87 | .91 | 357 | 23 |
| | | (.31) | (.12) | (.07) | (.02) | | | | | | |
| | Short | 7.49* | 1.52* | .43* | .93* | .50* | .02 | .00 | .99 | 893 | 29 |
| | | (.39) | (.27) | (.12) | (.00) | | | | | | |
| Poland | Medium | 5.33* | .15 | .14* | .79* | .54* | .99 | .00 | .95 | 785 | 29 |
| | | (.18) | (.16) | (.05) | (.01) | | | | | | |
| | Forward | 8.21* | 1.53* | .35* | .93* | .49* | .07 | .00 | .98 | 791 | 29 |
| | | (.47) | (.36) | (.10) | (.01) | | | | | | |

Notes: Estimated equation (10) $E_t i_{t+l} = (1-\rho)\alpha_0 + \alpha_1(1-\rho)E_t\pi_{t+k} + \alpha_2(1-\rho)E_t(\tilde{y}_{t+k}) + \rho i_t + \epsilon_t$ by the means of a serial correlation model where (12) $\epsilon_{t,i} = \varphi_i \epsilon_{t-1,i}$; values in parentheses present panel corrected standard errors applying the Prais-Winsten model; the Hausman test indicates to use the fixed-effects estimator on a one percent significance level; $\alpha_1 > 1$ ($\alpha_2 > 0$) represents the significance level of a $Chi^2$ test to test whether the *Taylor-principle* holds with the null hypothesis that $\alpha_1 \leq 1$ ($\alpha_2 \leq 0$); the $R^2$ refers to the overall coefficient of determination; for readability within and between $R^2$ are left from the Table but available upon request; * (+) indicates significance at the one (ten) percent level, respectively.

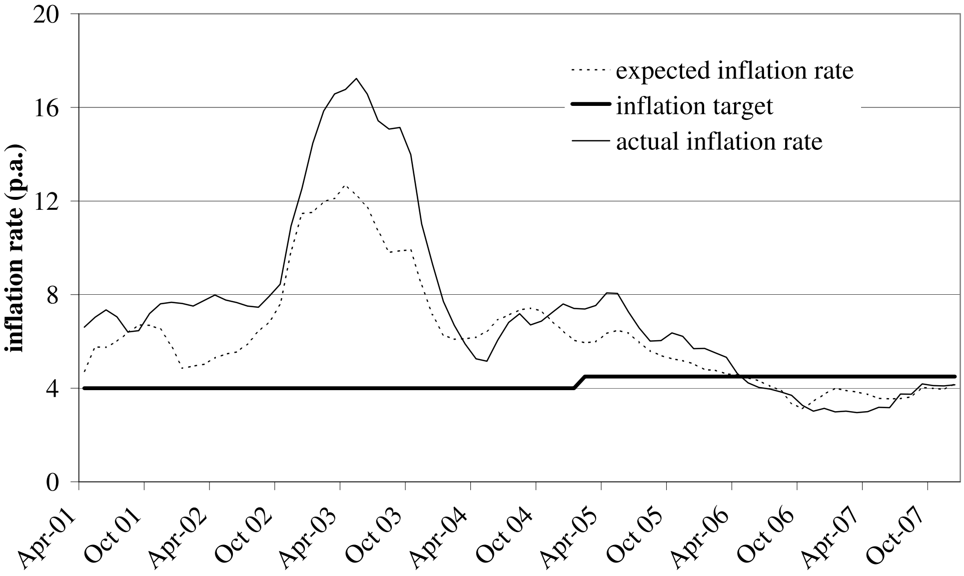


Figure 1: Inflation target, expected and actual inflation rate in Brazil

Note: Figure 1 shows the inflation target (thick solid line), the mean of the six-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line) at time $t$.

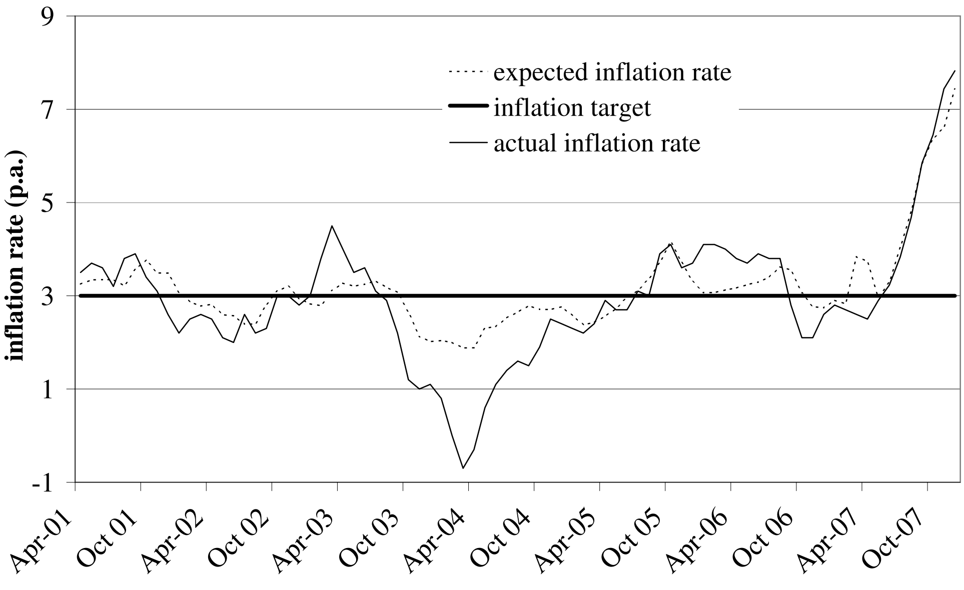


Figure 2: Inflation target, expected and actual inflation rate in Chile

Note: Figure 2 shows the inflation target (thick solid line), the mean of the six-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line) at time $t$.

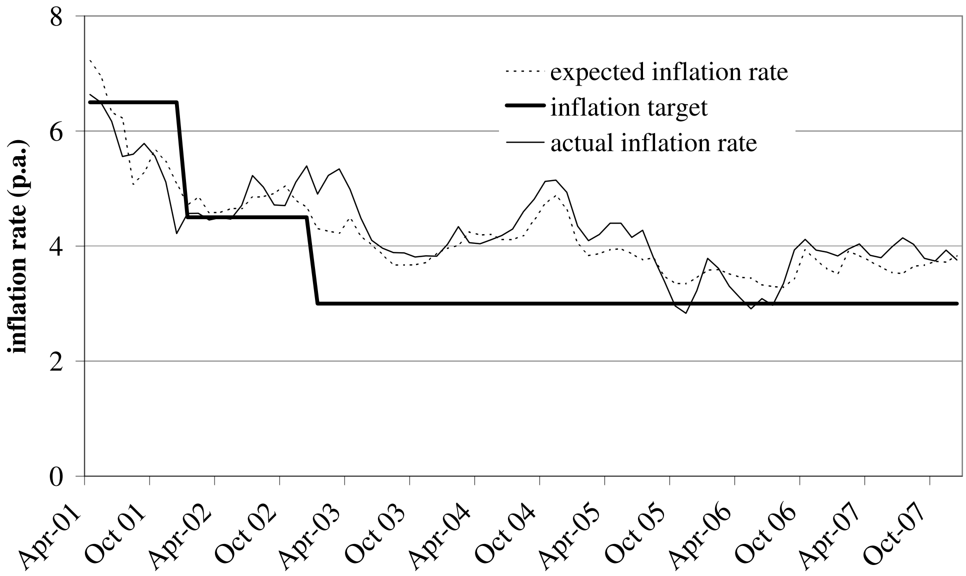


Figure 3: Inflation target, expected and actual inflation rate in Mexico

Note: Figure 3 shows the inflation target (thick solid line), the mean of the six-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line) at time $t$.

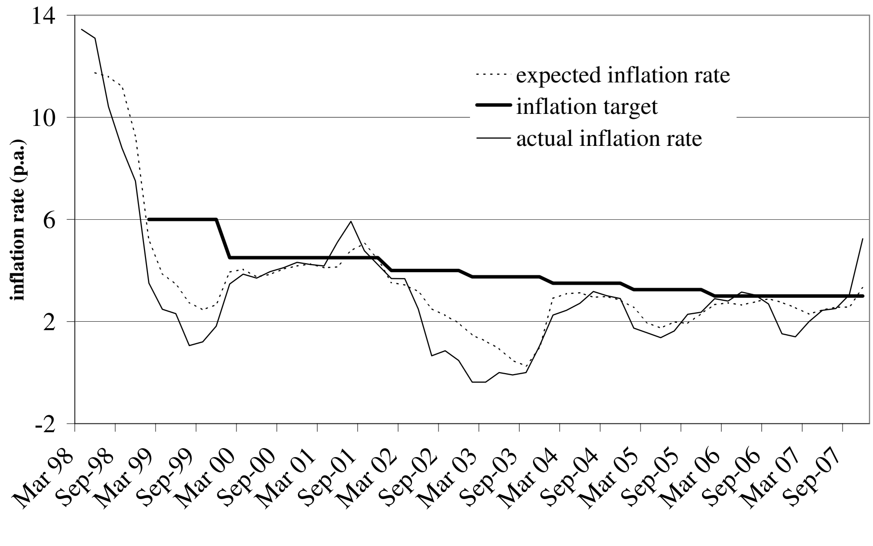


Figure 4: Inflation target, expected and actual inflation rate in the Czech Republic

Note: Figure 4 shows the inflation target (thick solid line), the mean of the six-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line) at time $t$.

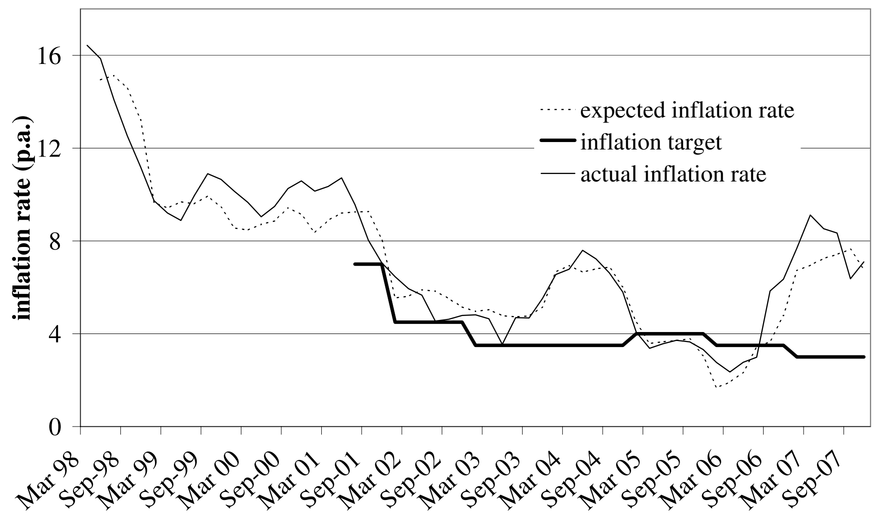


Figure 5: Inflation target, expected and actual inflation rate in Hungary

Note: Figure 5 shows the inflation target (thick solid line), the mean of the six-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line) at time $t$.

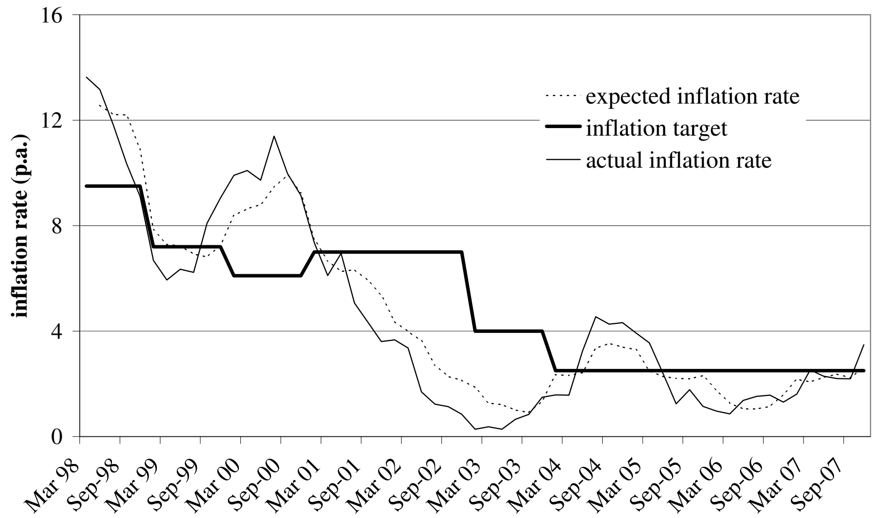


Figure 6: Inflation target, expected and actual inflation rate in Poland

Note: Figure 6 shows the inflation target (thick solid line), the mean of the six-months expected inflation rate $\bar{E}_t[\pi_{t+3}]$ (dotted line) and the actual inflation rate (solid line) at time $t$.

# Appendix: Calculation of the Weighted Average of Expected GDP and CPI

In order to generate a three months forecast we set the forecasted variable $f_t$ at time $t$ (= 1,2,.., 63 and 81, respectively) equal to the forecast of the current year $f_t^{cur}$ for forecasts collected before November of any ear (i.e. the remaining three months are all in the current year). For forecasts collected in November or December of any year, the three month forecast $f_t$ is calculate as a weighted arithmetic average of the forecast for the current year $f_t^{cur}$ and the next year $f_t^{next}$. We weight the forecast $f_t$ with the remaining number of months m (with $m = 2$ (for November forecasts) and $m = 1$ (for December forecasts)) at the time of the forecast $t$:

$$f_t = \frac{f_t^{cur} \cdot m + (3 - m) \cdot f_t^{next}}{3} \qquad (A.1)$$

In order to generate a twelve months forecast horizon which is consistent with the forecast horizon of the twelve months interest rate forecast we apply the outlined procedure with 1 (= December ) ≤ m ≤ 12 (= January). The twelve months $GDP$ and $CPI$ forecasts $f_t$ are as follows:

$$f_t = \frac{f_t^{cur} \cdot m + (12 - m) \cdot f_t^{next}}{12} \qquad (A.2)$$

This procedure is also applied by Heppke-Falk and Hüffner (2004) and Beck (2001). Both studies deal with data of the Consensus Economic Forecast poll and construct the arithmetic average as outlined above.