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DO TAX IMPLICATIONS CHANGE THE FISHER EFFECT FOR THE TURKISH ECONOMY?

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Abstract

This study investigates the validity of both the conventional and tax-adjusted Fisher effects using time series methods such as the *ARDL* bounds test, and Gregory-Hansen (G-H) cointegration test. To compare the conventional and tax-adjusted Fisher effects, we use two different time series data for interest rates: 1) interest rates adjusted for taxes, and 2) interest rates not adjusted for taxes. Using monthly changes in quarterly and annual interest and inflation rates, both the *G-H* and *ARDL* bound tests support the conventional and tax-adjusted Fisher effects. However, the magnitude of the inflation coefficient in the long-run relationship tends to decline for the tax adjusted Fisher effect.

Keywords: Fisher Effect, Taxation, Time Series Methods

JEL Code: C22, E43

Introduction

The existence of the Fisher effect in a country is important in deciding whether implemented economic policies are sustainable. When a country experiences deviations from its targeted inflation rate due to shocks to the system, the nominal interest rate could be a significant instrument for inflation targeting, given the cointegration between the nominal interest rate and inflation. Thus, for countries like Turkey, which aims at price stability, the nominal interest rate is a significant instrument for long-run inflation targeting.

The empirical validity of the Fisher effect has been restudied extensively in the literature in the last several decades, mostly due to the development of time series methods

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dealing with non-stationary data. Researchers now have various cointegration (Engle-Granger, Johansen) and VAR techniques in their hands to test the long-run relationship between individually non-stationary series (Atkins and Coe, 2002). However, potential shortcomings of these studies are 1) that these techniques require that individual series be non-stationary; and 2) most of these studies test the existence of a conventional Fisher effect rather than a tax-adjusted Fisher effect. The purpose of this study, then, is to test if the conventional and tax-adjusted Fisher effects are valid for Turkey, one of the most dynamic economies in the classification of developing countries. This study also aims to contribute to the current literature by including tax adaptation in the nominal interest rate for a developing country case. To that end, we use two different time series data for interest rates: 1) interest rates adjusted for taxes, and 2) interest rates not adjusted for taxes. Time series methods such as the ARDL bounds test of Pesaran, Shin and Smith (2001) and cointegration tests with a structural break are used to determine if consistent results for the existence of the Fisher effect can be reached.

In conclusion, for the post financial liberalization era in Turkey when tax adaptation has been especially implemented, declines in the inflation rate are associated with declines in the nominal interest rate. Therefore, in contrast to the findings for developed countries (see Atkins & Coe, 2002), this study finds support in favor of both types of Fisher effects in Turkey: the conventional and the tax-adjusted Fisher effect.

Theoretical Framework and Literature Review

Fisher (1930) hypothesized that the nominal interest rate was equal to the sum of the real interest rate and the expected inflation rate. He claimed that the nominal interest rate is comprised of the real interest rate and the expected inflation of the same period. Additionally, the Fisher effect asserts that there is a linear relationship between the nominal interest rate and the expected inflation, and assumes that the real interest rate does not change in the long run. If the real interest rate is not affected by monetary imbalances that affect inflation in the long term, this will cause a relationship between inflation and the nominal interest rate, leading to the likely existence of cointegration between the nominal interest rate and inflation.

To test the Fisher effect we use a standard model often encountered in the literature (e.g. Harrison (2010)):

$$int_t = real\,int_t + E_t(inf_{t+1})$$

According to the rational expectations;

$$E_t(inf_{t+1}) - inf_{t+1} = e_t$$

where $E\,(e_t) = 0$

Hence, the average expectational error is zero. Additionally, rational expectations suggest that the prediction error is uncorrelated with all the variables in the information set at the time of the prediction. Thus, substituting the realized inflation rate for the expected one the Fisher equation and our model is then;

$$int_t = real\,int_t + inf_{t+1} + e_t$$

More generally, this equation should be written as follows for developing countries;

$$(1+i_t) = (1+r_t)(1+\pi_{t+1})$$

$$i_t = r_t + \pi_{t+1} + r_t\pi_{t+1}$$

The $r_t\pi_{t+1}$ term is important for developing countries which have suffered from high inflation rates. Fisher (1930) examined the relationship between nominal interest rates and the rate of inflation for the U.S. and the U.K. using annual data over the 1890–1927 period for the US, and the 1820–1924 period for the U.K. Though Fisher (1930) asserted that the nominal interest rate and inflation should affect one another evenly, he failed at presenting this relationship empirically as his results indicated significant deviations from the Fisher effect.

Mundell (1963) attributed Fisher’s empirical results to the inflationary process by asserting that the inflationary process reduces the real rate through the wealth effect. Similarly, Tobin (1965) stated that nominal interest rates should adjust less than one to one due to inflationary pressures on the real interest rate. Darby (1975) and Feldstein (1976), on the other hand, argue that the nominal interest rate would adjust more than one for one ($1/(1-\tau)$) to the expected inflation rate due to the tax effect. Shome, Smith and Pinkerton (1988) argued that risk-averse investors need a premium to compensate them for any risks. Carmichael and Stebbing (1983) assumed that nominal interest rates on financial assets can be considered constant over time and that the real interest rate moves inversely with inflation.

Engsted (1996) and Crowder and Hoffman (1996) found support for the tax-adjusted Fisher effect for 13 OECD countries and the U.S., respectively. Carr, Pesando and Smith (1976), in contrast, could not find any evidence that supports the tax-adjusted Fisher effect for Canada while Atkins and Coe (2002) found evidence supportive of the conventional Fisher effect for the U.S. and Canada, but rejected the tax-adjusted Fisher effect for Canada and provided mixed results for the U.S. Evans and Lewis (1995) and Mishkin (1992) employed

Engle and Granger's (1987) bivariate cointegration test. Nevertheless, Mishkin's (1992) results did not support the Fisher hypothesis, while Evans and Lewis (1995) results supported the Fisher hypothesis only when regime shifts in the expected inflation process were taken into account. Since 1994 Turkey has been applying income tax on nominal interest incomes. In the event of a tax application on nominal interest incomes, the examination of the Darby-Feldstein effect is important in terms of policy proposals in countries with inflation targeting whose main instrument is short term interest rates.

For Turkey, Turgutlu (2004), Kesriyeli (1994), Şimşek and Kadılar (2006) and Köse *et al* (2012) obtained findings supportive of the Fisher effect in Turkey. Additionally, Kasman *et al* (2006) found results in favor of the presence of the Fisher effect in a large majority of 33 countries, including Turkey. All the mentioned studies for Turkey, however, excluded the tax adjustment procedure, which led us to examine the tax adjusted Fisher effect for Turkey.

Data

This study examines the validity of the Fisher effect in Turkey over the period of 1990:01 through 2011:11. Interest rates are proxied by quarterly and annual deposit rates and are from international financial statistics (IFS). Two different time series are used for nominal interest rates:

i) Interest rates adjusted for taxes

ii) Interest rates not adjusted for taxes

The quarterly inflation rate is measured as the 3 month to 3 month percentage change in the consumer price index (CPI) multiplied by 400. The annual inflation is measured as the yearly change in the CPI multiplied by 100. Turkey has implemented income tax on interest earnings since 1994. Tax rates are obtained from the decisions of the Council of Ministers. Because no publicly available data exist for double taxation and tax immunities they cannot be taken into account. So, the tax adjusted interest rates are only imperfect proxies of the theoretical ones. Nevertheless, they are also the only ones that can be derived from the existing data sources. Variable definitions, the period covered and the data sources are given in Table 1.

Quarterly and annual inflation, interest rate and the tax adjusted interest rate series are shown in figures 1 and 2, respectively. According to the graphs, changes in interest rates follow the changes in the inflation rate. This is particularly true for the crises periods of 1994 and 2001. These are also the periods where the interest rates tend to be negative or *circa* zero. It is also noticeable that the declines in the inflation rate are associated with declines in nominal interest rates in the last decade.

Table 1: Data			
Variables	**Definition**	**Source**	**Period**
inf3	Quarterly inflation rate. (APR)	IFS	1990:01-2011:08
int3	Quarterly nominal interest rate. (APR)	IFS	1990:01-2011:08
tint3	Tax adjusted quarterly nominal interest rate. (APR)	IFS	1990:01-2011:08
inf12	Annual inflation rate. (APR)	IFS	1990:01-2010:11
int12	Annual nominal interest rate. (APR)	IFS	1990:01-2010:11
tint12	Tax adjusted annual nominal interest rate. (APR)	IFS	1990:01-2010:11
tax3 (τ)	Tax rates for quarterly deposit revenue.	Council of Minister Decisions	1990:01-2011:08
tax12 (τ)	Tax rates for annual deposit revenue.	Council of Minister Decisions	1990:01-2010:11

Figure1:

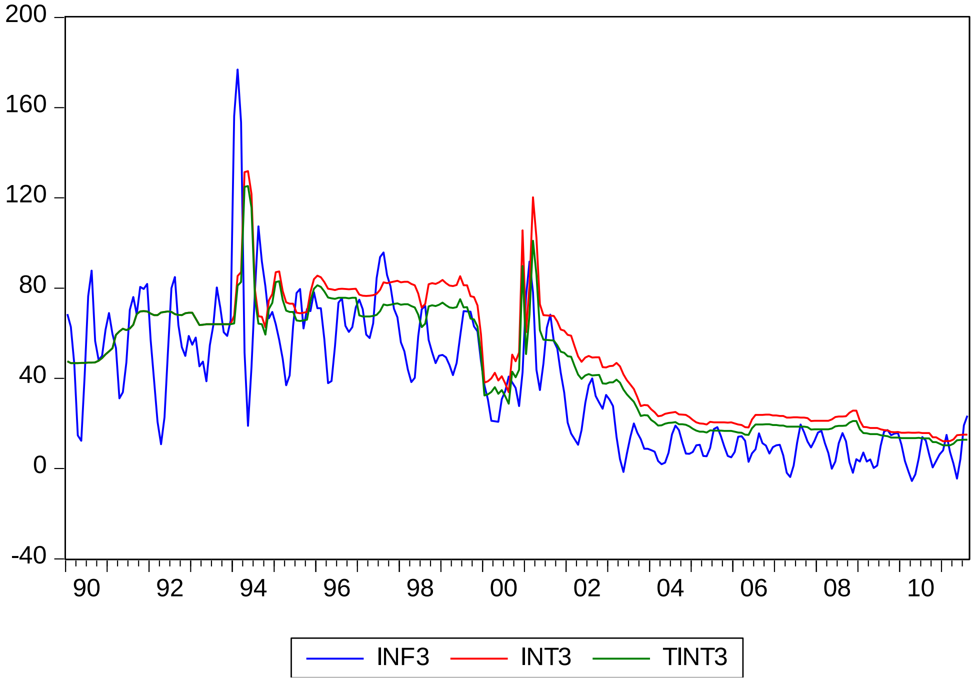
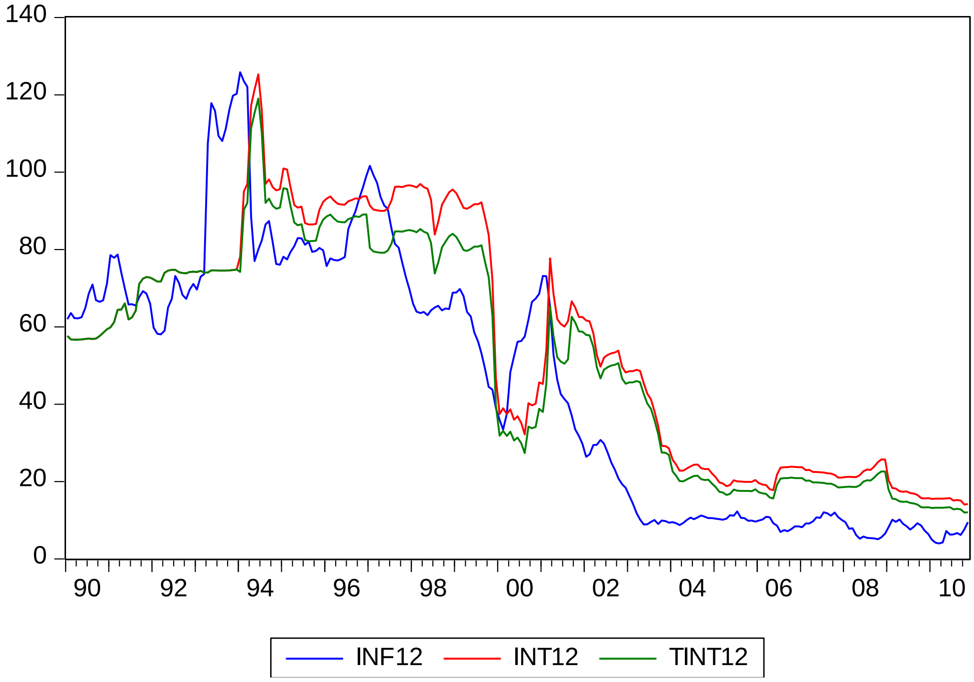


Figure 2:



Methodology

We start the analysis first by investigating the order of integration of the individual series via *ADF* and *KPSS* unit root tests. The empirical validity of the Fisher effect has been restudied extensively in the literature, thanks to the development of various time series methods such as Engle-Granger and Johansen cointegration tests and VAR techniques. However, these techniques require that individual series be non-stationary. Therefore, using the ARDL bound test to determine long-run relationships for the Fisher effect provides an advantage since this technique does not require the nominal interest rate and the inflation rate to be integrated at the same order. On the other hand, due to the likely changes in the inflation - interest rate relationship in the long run, the G-H cointegration test which allows for the possibility of a structural break is used.

ADF and KPSS Unit Root Tests

To test the time series properties of individual series we start with the *ADF* and *KPSS* tests. We perform the *ADF* test, starting with the largest model which includes trend and drift, and test the significance of trend and then drift (φ_3, φ_1 tests) as well. If the null of unit root and insignificant trend or drift cannot be rejected, we move to the next model. Lag lengths are determined starting with the maximum length of 12 and reducing them until we find a significant *t-value* for the individual lag. Since the *ADF* tests have low power, we also use the *KPSS* tests to corroborate the results. For the *KPSS* tests both the trend and drift models are used to test the null of no unit root, using Newey-West Bandwidth criteria with Barlett kernel. However, neither of the above tests takes into account a potential structural break when testing for a unit root. For that, we turn to the Zivot-Andrew (Z-A) (1992) unit root test.

Zivot-Andrews Unit Root Test (Z-A)

Unit root tests can be biased if the structural breaks are not accounted for. Z-A (1992) proposed a test that determines the break in the series endogenously. The individual series has a unit root under the null and is trend stationary with a structural break at a time T_B ($1<T_B<T$) under the alternative. The test is based on a procedure that trims a small percentage of the data at two ends and calculates the *t-value* for each of the remaining observations as being the potential break point. The observation with the smallest *t-value*, the least favorable outcome for the null, then, is chosen as the break point. Z-A (1992) provided the critical values for the test. Three different models are proposed for the test.

Model A permits one time change in the level of the series:

$$y_t = \mu^A + \theta^A DU_t(\lambda) + \beta^A t + \alpha^A y_{t-1} + \sum_{j=1}^{k} c_j^{\,A} \Delta y_{t-j} + e_t$$

Model B allows one time change in the slope of the trend function occurring at time $T_{B:}$

$$y_t = \mu^B + \gamma^B DT^*{}_t(\lambda) + \beta^B t + \alpha^B y_{t-1} + \sum_{j=1}^{k} c_j^{\,B} \Delta y_{t-j} + e_t$$

Finally, Model C permits a onetime change both in the level and the trend of the series.

$$y_t = \mu^C + \theta^C DU_t(\lambda) + \beta^C t + \gamma^C DT^*{}_t(\lambda) + \alpha^C y_{t-1} + \sum_{j=1}^{k} c_j^{\,C} \Delta y_{t-j} + e_t$$

DU(λ) , *DT*(λ)* are dummy variables, representing the change in the level and the trend of the series:

$DU_t(\lambda) = \begin{cases} 1, t > TB, \\ 0, otherwise \end{cases}$ $\quad DT^*{}_t(\lambda) = \begin{cases} t - TB, t > TB, \\ 0, otherwise \end{cases}$ $\quad TB = \lambda T$, and λ is the break fraction.

ARDL Bound Test Approach

Pesaran, Shin and Smith (PSS) (2001) developed an ARDL based test to detect the long-run relationship between the variables regardless of the individual series being I(1) or I(0). The most important advantage of using this method for testing the validity of the Fisher effect is that it does not require the nominal interest rate and the inflation series to be integrated at the same order. The model to investigate the long-run level relationship is as follows:

$$\Delta Y_t = c_0 + c_1 Y_{t-1} + c_2 X_{t-1} + \sum_{i=1}^{m} c_{yi} Y_{t-i} + \sum_{i=0}^{m} c_{xi} X_{t-i} + e_t$$

In order to apply the bound test, first, the order of appropriate lag length is determined from a VAR model, using criterion such as AIC, SIC, or HQ. The model should be free of autocorrelation at the selected lag. So, the additional lags are added if needed. The null of no long-run relationship can be tested by an *F* test, restricting c_1=c_2=0. The distribution of the test statistic is non-standard and cannot be compared with the standard *F* distribution, but the critical values were provided in PSS (2001), If the test statistic is below the lower bound, the null cannot be rejected; if it is above the upper bound the null is rejected. In cases where the test statistic is between the lower and upper bound, no certain conclusion can be reached regarding the existence of a long-run level relationship. Once the existence of a long-run level relationship is concluded, the variables in the *ARDL* model are allowed to have different lag lengths and the model is reestimated to determine the long and short-run relationships. The long-run coefficients can be estimated using the *ARDL* model:

$$Y_t = \beta_0 + \sum_{i=1}^{m} \beta_{yi} Y_{t-i} + \sum_{i=0}^{n} \beta_{xi} X_{t-i} + e_t$$

To determine the short-run relationship between the variables, the error correction representation of the *ARDL* model is used:

$$\Delta Y_t = \delta_0 + \sum_{i=1}^{m-1} \delta_{yi} \Delta Y_{t-i} + \sum_{i=0}^{n-1} \delta_{xi} \Delta X_{t-i} + \rho ecm_{t-1} + e_t$$

Where ecm_{t-1} represents the long-run equilibrium, and the differences of *X* and *Y* stand for the short-run dynamics.

Gregory- Hansen Cointegration Test

Gregory and Hansen (G-H) (1996a, b) proposed a residual based procedure for testing the null of no cointegration against the alternative of cointegration with a structural break. The time of the break point is decided endogenously. The dummy variable, specifying the timing of the structural change, is defined as follows (Gregory Hansen, 1996a, p. 102-103):

$$\theta_{t\tau} = \begin{cases} 0\ , if \ \ t \leq [n\tau], \\ 1\ , if \ \ t > [n\tau], \end{cases}$$

where n is the number of observations, $\tau \in (0,1)$ indicates the timing of the structural change point and [ ] denotes the integer part. The four models proposed are as follow:

Model 1: Level Shift (C) Model

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha^T y_{2t} + e_t\,,\ t=1,\dots,\ n$$

This model allows a break in the intercept, where μ_1 represents the intercept before the break, and μ_2 represents the change in the intercept at the time of the break.

Model 2: Level Shift with trend (C/T) Model

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t\,,\ t=1,\dots,\ n$$

This model takes the break into account in the intercept in the presence of a deterministic trend.

Model 3: Regime Shift (C/S) Model

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T \varphi_{t\tau} y_{2t} + e_t\ ,\ \ t=1,\dots,\ n.$$

The model does not include a deterministic trend, but permits a break in the intercept as well as in the slope coefficients. μ_1 and μ_2 are as in Model 1, α_1^T and α_2^T, on the other hand, represent the slope coefficients before the regime shift and the changes in the slope coefficients after the regime change, respectively.

Model 4: Trend and Regime Shift (C/T/S) Model

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta_1 t + \beta_2 \varphi_{t\tau} t + \alpha_1^T y_{2t} + \alpha_2^T \varphi_{t\tau} y_{2t} + e_t\,,\ \ t=1,\dots,\ n$$

The model was proposed by G-H (1996b) and is the largest model which allows a break in the level, trend and slope coefficients. μ_1, β_1, and α_1^T represent the intercept, trend and slope coefficients, respectively, before the regime change, μ_2, β_2 and α_2^T, in contrast, denote the changes in the parameters after the regime change.

The test is based on a procedure that trims a small percentage of the data at two ends and calculates the *ADF*, and the Phillips (1987) Z_α and Z_t statistics for each of the remaining observations as being the potential break point. The observation with the smallest test-*value*, the least favorable outcome for the null, then, is chosen as the break point. Hence the three test statistics proposed are as follows (Gregory and Hansen 1996a, b).

$$ADF^* = \min_{\tau \in T} ADF(\tau)\ , \qquad Z_\alpha^* = \min_{\tau \in T} Z_\alpha(\tau)\ , \qquad Z_t^* = \min_{\tau \in T} Z_t(\tau)$$

Empirical Results

Unit Root Tests:

The results for the *ADF* and *KPSS* unit root tests are given in Table 2. All the *ADF* and *KPSS* tests indicate that all the variables are non-stationary at conventional significance levels, except *int3* and *tint3* which are significant at the 5% level, using the largest model for the *ADF* test. However, the null of no unit root is rejected for both of these variables by the *KPSS* tests. So, we conclude that all the variables are integrated order of one according to the results.

Table 2: ADF and KPSS Unit Root Tests

Series	**ADF**								**KPSS**			
	Trend&Drift			**Drift**			**None**		**Trend&Drift**		**Drift**	
	k	τ_3	ϕ_3	*k*	τ_2	ϕ_1	*k*	τ_1	*K*	η_μ	*k*	η_τ
inf3	10	-2.8432	4.0595	10	-1.1428	0.9036	10	-1.2229	8	**0.2809**	11	**1.8850**
int3	3	**-3.8686**	**7.6801**	3	-1.8934	1.8334	5	-0.9276	11	**0.3430**	12	**1.5652**
tint3	5	**-3.4418**	6.2384	5	-1.1195	0.7405	5	-0.9505	11	**0.3007**	12	**1.7075**
inf12	12	-1.8449	1.7682	12	-0.4587	1.0586	12	1.3958	11	**0.2522**	12	**1.7637**
int12	4	-2.6175	3.9127	4	-0.7918	0.4849	4	-0.8896	12	**0.2795**	12	**1.6112**
tint12	4	-2.7388	4.1892	4	-0.7046	0.4825	4	-0.9398	11	**0.2790**	12	**1.6100**
Critical Values	**%1** **%5** **%10**	**-3.98** **-3.42** **-3.13**	**8.34** **6.30** **5.36**		**-3.44** **-2.87** **-2.57**	**6.47** **4.61** **3.79**		**-2.58** **-1.95** **-1.62**		**0.216** **0.146** **0.119**		**0.739** **0.463** **0.347**

k is the number of the lags and the bandwidth parameter for the *ADF* and *KPSS* tests, respectively.

In the case of the *Z-A* test, the effect of a potential structural break on the individual series' degree of integration is investigated for all the three models. The results are presented in Table 3. First of all, all the break dates for all the series occur in or around the two crises years of 1994 and 2001 for Turkey. Secondly, none of the tests can decisively reject the null of unit root (at 1% significance level), except the results for *int3* and *tint3* from Model B. Nonetheless, Sen (2003) suggested that Model C is preferable to the other models. Although the results from Model C suggest that *int3*, t*int3* and *inf12* are stationary with a structural break at the 5% level, the null of unit root cannot be rejected at the 1% significance level, even after allowing a break both in the level and trend. So, the conclusion based on the *ADF* and the *KPSS* tests regarding the degree of integration of all the variables may still very well be valid.

Table 3: Zivot- Andrews Unit Root Test														
Series	**Model A**				**Model B**				**Model C**					
	K	**Test statistic**	**T_B**	**DU Prob.**	**K**	**Test statistic**	**T_B**	**DT Prob.**	**K**	**Test statistic**	**T_B**	**DU Prob.**	**DT Prob.**
Inf3	10	**-4.60**	142	0.000	10	-3.66	49	0.023	10	-4.35	142	0.000	0.827
int3	3	**-5.25**	48	0.000	3	**-5.25**	52	0.000	3	**-5.57**	49	0.045	0.046
tint3	5	-4.52	136	0.001	3	**-5.36**	52	0.000	3	**-5.55**	49	0.110	0.041
Inf12	12	-3.09	133	0.007	12	-4.03	40	0.014	9	**-5.10**	133	0.000	0.937
int12	12	-3.09	133	0.007	12	-2.87	40	0.025	12	-3.04	37	0.196	0.473
tint12	1	-4.15	116	0.002	1	-3.78	50	0.014	1	-4.58	116	0.000	0.023
Critical Values	**%1** **%5** **%10**	**-5.34** **-4.8** **-4.58**				**-4.93** **-4.42** **-4.11**				**-5.57** **-5.08** **-4.82**			
k is the number of the lags determined by a significant t - ratio reduced from 12 and T_B is the break point.													

Table 4-a : ARDL (4,10) for int3			
Regressor	Coefficient	Standard Error	T-Ratio[Prob]
inf3	**0.96853**	0.060533	16.00[.000]
C	13.8137	2.7836	4.96[.000]
95% Lower Bound	95% Upper Bound	90% Lower Bound	90% Upper Bound
4.9791	5.7037	5.7370	4.0755
F-statistic ARDL(4,4)	**14.1673**		
ARDL(4,4): Serial Correlation:CHSQ(12) = 12.1411[.434] F(12,226) = .96946[.479]			

Table 4-b: ARDL(4,10) for tint3			
Regressor	Coefficient	Standard Error	T-Ratio[Prob]
tinf3	**0.93717**	0.049058	19.1033[.000]
C	10.1258	2.2563	4.4878[.000]
95% Lower Bound	95% Upper Bound	90% Lower Bound	90% Upper Bound
4.9791	5.7037	5.7370	4.0755
F-statistic ARDL(1,1)	**22.0055**		
ARDL(1,1): Serial Correlation:CHSQ(12) = 19.8984[.069] F(12,232) = 1.6865[.071]			

Table 4-c: ARDL(5,3) for int12			
Regressor	Coefficient	Standard Error	T-Ratio[Prob]
inf12	**0.93664**	0.071472	13.1050[.000]
C	10.7852	4.1144	2.62[.009]
95% Lower Bound	95% Upper Bound	90% Lower Bound	90% Upper Bound
5.1343	5.6226	4.1847	4.7689
F-statistic ARDL(2,2)	**17.1184**		
ARDL(2,2): Serial Correlation:CHSQ(12) = 11.2128[.511] F(12,221) = .90655[.541]			

Table 4-d: ARDL(2,3) for tint12			
Regressor	Coefficient	Standard Error	T-Ratio[Prob]
inf12	**0.86864**	0.066707	13.0217[.000]
C	9.8691	3.8497	2.56[.011]
95% Lower Bound	95% Upper Bound	90% Lower Bound	90% Upper Bound
5.1343	5.6226	4.1847	4.7689
F-statistic ARDL(2,2)	**14.9639**		
ARDL(2,2): Serial Correlation:CHSQ(12) = 10.5441[.568] F(12,221) = .84999[.599]			

Cointegration Test Results

The *ARDL* bound test results and the estimated long-run level relationships are displayed in tables 4-a to 4-d. The calculated *F-values* are above the upper bounds in all cases, indicating the existence of a long-run relationship between the nominal interest rate and the expected inflation rate. The *LM* tests as well indicate that the errors are free of serial correlation at the 5% significance level. Once the cointegration relationship is verified, each model is reestimated, allowing the right-hand variables to have different lag lengths. In all the cases the point estimates for the expected inflation coefficients are slightly less than unity, but the 95% confidence intervals contain unity. Thus, the *ARDL* results support both the conventional and tax adjusted Fisher effect for Turkey.

In order to take into account a potential structural break in the long-run relationship, we use the *G-H* cointegration test. The results are provided in Table 5. The *G-H* tests provide a substantial support for cointegration with a structural break. In particular, using a 3-month inflation and interest rate series, all three tests statistics reject the null of no cointegration for all the structural change models and for both the tax adjusted and non-adjusted interest rates. For the 12-month inflation and interest rate series, the *ADF* tests further verify the existence of cointegration for all the models except the level shift (C) and the trend and the regime shift (C/T/S) models.

For the models where the null of no cointegration is rejected in favor of cointegration with a structural break, we estimate the following regressions to see how the cointegration relationships may have changed over time.

$$C : \text{int}_t = \beta_0 + \beta_1 dummy + \beta_2 E_t(\inf_{t+k}) + e_t$$

$$C/T : \text{int}_t = \beta_0 + \beta_1 trend + \beta_2 dummy + \beta_3 E_t(\inf_{t+k}) + e_t$$

$$C/S : \text{int}_t = \beta_0 + \beta_1 dummy + \beta_2 E_t(\inf_{t+1}) + \beta_3 dummyE_t(\inf_{t+k}) + e_t$$

$$C/T/S: \text{int}_t = \beta_0 + \beta_1 trend + \beta_2 dummy + \beta_3 E_t(\text{inf}_{t+1}) + \beta_4 dummyE_t(\text{inf}_{t+k})$$
$$+ \beta_4 dummytrend + e_t$$

where ***tint**$_t$* would replace ***int**$_t$* when the model is estimated for the tax adjusted Fisher effect.

Table 5 :Gregory Hansen Test Results

Model	k	ADF	Break Point	Z_t	Break Point	Z_a	Break Point
C (int3,inf3)	1	-7.27*	0.6538 (170)	-6.57*	0.6538 (170)	-76.03*	0.6538 (170)
C/T (int3,inf3)	1	-7.43*	0.2076 (54)	-6.55*	0.2076 (54)	-74.25*	0.2076 (54)
C/S (int3,inf3)	1	-7.36*	0.2576 (67)	-8.24*	0.1923 (50)	-112.47*	0.1923 (50)
C/T/S (int3,inf3)	1	-7.47*	0.4307 (112)	-7.19*	0.4307 (112)	-88.61*	0.4307 (112)
C (tint3,inf3)	1	-7.55*	0.6500 (169)	-6.61*	0.6269 (163)	-77.18*	0.6269 (163)
C/T (tint3,inf3)	1	-7.47*	0.2076(54)	-6.52*	0.2000 (52)	-73.37*	0.2000 (52)
C/S (tint3,inf3)	1	-7.70*	0.35384 (92)	-8.21*	0.1923 (50)	-112.35*	0.1923 (50)
C/T/S (tint3,inf3)	1	-7.27*	0.43076 (112)	-6.96*	0.4307 (112)	-83.57*	0.4307 (112)
C (int12,inf12)	9	-4.21	0.59362 (149)	-4.02	0.2191 (55)	-32.08	0.2191 (55)
C/T (int12, inf12)	9	-5.13*	0.23505 (59)	-4.21	0.2191 (55)	-34.09	0.2191 (55)
C/S (int12, inf12)	9	-5.07*	0.15139 (38)	-4.20	0.1832 (46)	-34.70	0.1832 (46)
C/T/S (int12, inf12)	9	-5.27*	0.49800 (125)	-4.63	0.4581 (115)	-41.51	0.4581 (115)
C (tint12, inf12)	9	-4.85*	0.71314 (179)	-3.86	0.2191 (55)	-29.75	0.2191 (55)
C/T (tint 12, inf12)	9	-5.09*	0.25498 (64)	-4.05	0.2191 (55)	-31.52	0.2191 (55)
C/S (tint 12, inf12)	9	-4.89*	0.72509 (182)	-4.04	0.1832 (46)	-32.17	0.1832 (46)
C/T/S (tint 12, inf12)	9	-5.04	0.49800 (125)	-4.45	0.4541 (114)	-38.46	0.4541 (114)
Critical Values							
C	%1/5/10	-5.13 / -4.61 / -4.34		-5.13 / -4.61 / -4.34		-50.07 / -40.48 / -36.19	
C/T	%1/5/10	-5.45 / -4.99 / -4.72		-5.45 / -4.99 / -4.72		-57.28 / -47.96 / -43.22	
C/S	%1/5/10	-5.47 / -4.95 / -4,68		-5.47 / -4.95 / -4,68		-57.17 /- 47.04 / -41.85	
C/T/S	%1/5/10	-6.02 / -5.50 / -5.24		-6.02 / -5.50 / -5.24		-69.37 / -58.58 / -53.31	

Corresponding Break Dates

Observation	Date	Observation	Date	Observation	Date	Observation	Date
38	1993:02	55	1994:07	112	1999:04	163	2003:07
46	1993:10	59	1994:11	114	1999:06	169	2004:01
50	1994:02	64	1995:04	115	1999:07	170	2004:02
52	1994:04	67	1995:07	125	2000:05	179	2004:11
54	1994:06	92	1997:08	149	2002:05	182	2005:02

The results in Table 6 show that the effect of inflation on the nominal interest rate is quite close to one both for the conventional and tax-adjusted Fisher effect at least in one of the regimes. When the standard errors are taken into consideration, the presence of the Fisher effect cannot be rejected. Endogenously determined break dates follow the 1994 and 2001 economic crises in general. Typically for breaks following the 1994 crisis the impact of inflation on the nominal interest rate increases significantly and the Fisher effect holds. This can be seen, for example, for the tax-adjusted and non-adjusted quarterly interest rates (tint3 and int3) C/S models with break points 92 and 67. Before the break point 67, one percentage point increase in the inflation rate causes a 0.036 percentage point increase in the nominal interest rate. However, after the break, a one percentage point increase in inflation causes a rise of 0.92 percentage point, almost one to one, in the interest rate. This can be explained by the fact that the 1990s in Turkey is the high inflation period with rising inflationary expectations, especially after 1994.

Table 6: Estimated Cointegration Regressions with a Structural Break for Selected Models

Model	Coefficient	Std. Error	t-Stat	Prob.
GH- 3 **(67) 1995:07** $C/S: \text{int}\,3_t = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + \beta_3 dummy \inf_{t+3} + e_t$	65.18344 -47.6515 0.036104 **0.883771**	3.98714 4.260705 0.058165 0.069701	16.34842 -11.1839 0.620721 12.67954	1.26E-41 6.61E-24 5.35E-01 6.45E-29
GH- 3 **(50) 1994:02** $C/S: \text{int}\,3_t = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + \beta_3 dummy \inf_{t+3} + e_t$	55.22035 -34.3302 0.104499 **0.723845**	5.120188 5.373494 0.076923 0.085755	10.78483 -6.38881 1.35849 8.440893	1.33E-22 7.84E-10 1.76E-01 2.36E-15
GH- 3 **(92) 1997:08** $C/S: t\,\text{int}\,3_t = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + \beta_3 dummy \inf_{t+3} + e_t$	66.49513 -53.4715 0.023885 **0.847529**	3.297014 3.558686 0.047959 0.063401	20.16829 -15.0256 0.498033 13.3677	8.38E-55 5.14E-37 6.19E-01 2.85E-31
GH- 3 **(50) 1994:02** $C/S: t\,\text{int}\,3_t = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + \beta_3 dummy \inf_{t+3} + e_t$	56.26918 -39.4145 0.084485 **0.691695**	4.742464 4.977083 0.071248 0.079428	11.86497 -7.91919 1.185789 8.708408	3.63E-26 7.30E-14 2.37E-01 3.88E-16
GH-2 **(59) 1994:11** $C/T: \text{int}\,12_t = \beta_0 + \beta_1 dummy + \beta_2 trend + \beta_3 \inf_{t+12} + e_t$	28.72996 23.85144 -0.17684 **0.880572**	5.750068 3.078034 0.032274 0.08004	4.996456 7.74892 -5.47935 11.00169	1.11E-06 2.40E-13 1.05E-07 3.58E-23
GH-4 **(125) 2000:05** $C/T/S: \text{int}\,12_t = \beta_0 + \beta_1 dummy + \beta_2 trend + \beta_3 dummytrend + \beta_4 \inf_{t+12} + \beta_5 dummy \inf_{t+12} + e_t$	20.33818 58.48628 0.227903 -0.50574 **0.834225** **-0.56428**	6.645581 11.95197 0.029973 0.053558 0.1052 0.164487	3.060406 4.893443 7.603667 -9.44287 7.929883 -3.43056	0.002456 1.80E-06 6.13E-13 2.99E-18 7.78E-14 0.000707
GH-1 **(179) 2004:11** $C: t\,\text{int}\,12_t = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+12} + e_t$	24.38205 -14.2655 **0.891125**	2.443596 2.542056 0.048956	9.977937 -5.61181 18.20273	6.25E-20 5.34E-08 1.36E-47
GH-2 **(64) 1995:04** $C/T: t\,\text{int}\,12_t = \beta_0 + \beta_1 dummy + \beta_2 trend + \beta_3 \inf_{t+12} + e$	33.29126 12.72898 -0.15449 **0.805539**	5.476005 2.923747 0.030786 0.075516	6.079479 4.353652 -5.0182 10.6671	4.55E-09 1.96E-05 9.98E-07 4.26E-22
GH- 3 **(182) 2005:02** $C/S: t\,\text{int}\,12_t = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+12} + \beta_3 dummy \inf_{t+12} + e_t$	22.94279 -6.96744 **0.917369** **-0.74958**	2.376743 6.72416 0.048013 0.758202	9.653039 -1.03618 19.10667 -0.98862	6.51E-19 0.301131 1.40E-50 0.323815

An opposite picture arises for the break points around or after 2001. In this case the impact of inflation on the interest rate becomes weaker after the break. This can be attributed to decreasing inflationary expectations as the last decade is a relatively lower inflation period. See for example the C/S model for the tax-adjusted and non-adjusted annual interest rates (tint12, int12) with break points 182 and 125.

Conclusion

This study investigated the conventional and tax adjusted Fisher effect for Turkey to humbly suggest a policy for an incumbent government. Because the real interest rate for developed small open economy countries is closer to the world interest rate in comparison to developing countries, it is reasonable to expect that the Darby and Feldstein effect should apply for the developed countries.

Mundell (1963) and Tobin (1965) indicated that inflation would affect the real monetary stability when the economic agents prefer holding other assets rather than money due to the increase in the anticipated inflation. This would lower the nominal rates instantly. With regard to the Mundell-Tobin effect for tax implications, the welfare loss caused by the inflationary process can be the reason for the tendency of inflation coefficients to be lower than unity in the Fisher equation.

In this study, using the CPI-based inflation rate and the tax-adjusted and non-adjusted deposit rates, we investigated the validity of the Fisher effect with the *ARDL* bounds test and the Gregory-Hansen cointegration test. The results from the *ARDL* method are, in general, supportive of both the conventional and tax-adjusted Fisher effects, but the magnitude of the inflation coefficients tends to decline for the tax adjusted Fisher effect. The *G-H* tests provide substantial evidence for the existence of the cointegration relationship between the expected inflation and the nominal interest rate. However, about half the coefficients for the inflation rate in the cointegration relationships are less than unity and do not significantly change for the tax adjusted Fisher effect, a fact in contradiction to the Darby-Feldstein effect, but compatible with Mundell-Tobin effect.

Real interest rates in Turkey have been above the world interest rate in the post financial liberalization era. Hence, investors may continue to prefer investing in TL assets as long as real interest rates are higher than the rest of the world. According to the results from the G-H cointegration regressions, tax implications on the nominal interest rate do not affect the saving behavior. Therefore, public revenues (taxes) should be obtained with direct rather than indirect taxation for price stability and lower inflation.

It should be noted that deposit interest rates do not account for exceptional cases such as tax immunities and double taxation. Savings deposits are preferred by small investors rather than professional ones. Therefore, the results here should be compared with results using other measures of nominal interest rates such as Treasury bill rates. Furthermore, the G-H test was designed to account for only a single structural break. The results might be biased, if there are multiple breaks, given the two important economic crises (1994, 2001) in the last two decades, it very likely that the Fisher equation in Turkey had suffered more than one structural break (see Figure 2, p. 6). So, the results here should be interpreted with caution. Future research is needed to scrutinize the Fisher effect further with multiple structural breaks and the alternative measures of interest rates.

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Appendices

Table 7 : : Estimated Cointegration Regressions with a Structural Break

Model	Coefficient	Standard Error	Test Statistic	Prob.
GH-1 **(170) 2004:02** $C:\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + e_t$	50.00289 -32.6933 0.308653	2.304709 2.480339 0.03868	21.69597 -13.181 7.979694	5.07E-60 1.18E-30 4.88E-14
GH-2 **(54) 1994:06** $C/T:\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 trend +$ $\beta_3 \inf_{t+3} + e_t$	56.2898 25.51869 -0.28249 0.283219	3.68262 3.150306 0.023143 0.043906	15.28526 8.100385 -12.2064 6.450615	6.41E-38 2.25E-14 2.59E-27 5.53E-10
GH- 3 **(67) 1995:07** $C/S:\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3}$ $+ \beta_3 dummy \inf_{t+3} + e_t$	65.18344 -47.6515 0.036104 0.883771	3.98714 4.260705 0.058165 0.069701	16.34842 -11.1839 0.620721 12.67954	1.26E-41 6.61E-24 5.35E-01 6.45E-29
GH- 3 **(50) 1994:02** $C/S:\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3}$ $+ \beta_3 dummy \inf_{t+3} + e$	55.22035 -34.3302 0.104499 0.723845	5.120188 5.373494 0.076923 0.085755	10.78483 -6.38881 1.35849 8.440893	1.33E-22 7.84E-10 1.76E-01 2.36E-15
GH-4 **(112) 1999:04** $C/T/S:\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 trend$ $+ \beta_3 dummytrend_t + \beta_4 \inf_{t+3}$ $+ \beta_5 dummy \inf_{t+3} + e_t$	56.63458 3.810938 0.282944 -0.48941 -0.00497 0.623352	3.12398 6.937777 0.030858 0.041782 0.04086 0.074601	18.12898 0.549302 9.16916 -11.7134 -0.1216 8.355836	1.07E-47 5.83E-01 1.68E-17 1.27E-25 0.903308 4.28E-15

Table 8 : : Estimated Cointegration Regressions with a Structural Break

Model	Coefficient	Standard Error	Test Statistic	Prob.
GH-1 **(169) 2004:01** $C:t\,\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + e_t$	44.75839 -30.8462 0.310637	2.146305 2.297241 0.035917	20.8537 -13.4275 8.648784	3.38E-57 1.67E-31 5.73E-16
GH-1 **(163) 2003:07** $C:t\,\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+3} + e_t$	48.51882 -33.719 0.258099	2.1786 2.260112 0.035829	22.27064 -14.9192 7.203581	6.25E-62 1.11E-36 6.49E-12
GH-2 **(54) 1994:06** $C/T:t\,\text{int}\,3_t \quad = \beta_0 + \beta_1 dummy + \beta_2 trend +$ $\beta_3 \inf_{t+3} + e_t$	57.25922 17.8836 -0.27022 0.252094	3.286902 2.811788 0.020656 0.039188	17.42042 6.360223 -13.0817 6.432972	2.33E-45 9.21E-10 2.73E-30 6.11E-10

Table 9 : : Estimated Cointegration Regressions with a Structural Break

Model	Coefficient	Standard Error	Test Statistic	Prob.
GH-2 **(59) 1994:11** $C/T: \text{int}12_t \quad = \beta_0 + \beta_1 dummy + \beta_2 trend + \beta_3 \inf_{t+12} + e_t$	28.72996 23.85144 -0.17684 0.880572	5.750068 3.078034 0.032274 0.08004	4.996456 7.74892 -5.47935 11.00169	1.11E-06 2.40E-13 1.05E-07 3.58E-23
GH- 3 **(38) 1993:02** $C/S: \text{int}12_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+12} + \beta_3 dummy \inf_{t+12} + e_t$	43.90485 32.14544 -0.27204 0.534311	6.429235 4.063556 0.040792 0.103177	6.828938 7.910668 -6.66894 5.178612	6.60E-11 8.58E-14 1.68E-10 4.64E-07
GH-4 **(125) 2000:05** $C/T/S: \text{int}12_t \quad = \beta_0 + \beta_1 dummy + \beta_2 trend + \beta_3 dummytrend_t + \beta_4 \inf_{t+12} + \beta_5 dummy \inf_{t+12} + e_t$	20.33818 58.48628 0.227903 -0.50574 0.834225 -0.56428	6.645581 11.95197 0.029973 0.053558 0.1052 0.164487	3.060406 4.893443 7.603667 -9.44287 7.929883 -3.43056	0.002456 1.80E-06 6.13E-13 2.99E-18 7.78E-14 0.000707

Table 10 : : Estimated Cointegration Regressions with a Structural Break

Model	Coefficient	Standard Error	Test Statistic	Prob.
GH-1 **(179) 2004:11** **2005:02** $C: t\,\text{int}12_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+12} + e_t$	24.38205 -14.2655 0.891125	2.443596 2.542056 0.048956	9.977937 -5.61181 18.20273	6.25E-20 5.34E-08 1.36E-47
GH-2 **(64) 1995:04** $C/T: t\,\text{int}12_t \quad = \beta_0 + \beta_1 dummy + \beta_2 trend + \beta_3 \inf_{t+12} + e_t$	33.29126 12.72898 -0.15449 0.805539	5.476005 2.923747 0.030786 0.075516	6.079479 4.353652 -5.0182 10.6671	4.55E-09 1.96E-05 9.98E-07 4.26E-22
GH- 3 **(182) 2005:02** $C/S: t\,\text{int}12_t \quad = \beta_0 + \beta_1 dummy + \beta_2 \inf_{t+12} + \beta_3 dummy \inf_{t+12} + e_t$	22.94279 -6.96744 0.917369 -0.74958	2.376743 6.72416 0.048013 0.758202	9.653039 -1.03618 19.10667 -0.98862	6.51E-19 0.301131 1.40E-50 0.323815