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The Month and Holy Days Effects on the Volatility of Trade Deficit: Evidence from Turkey

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Abstract. This paper tries to examine the month and Holy Days (the feast of Ramadan and the feast of Sacrifice) effects on the volatility of trade deficit of the Turkish economy by using conditional variance framework. Three separate models including dummy variables are estimated in both conditional mean and variance equations. The estimation results reveal that the Holy Days' effect is significant in the mean equation and that trade deficit occurs mostly in December and least in January. The volatility content of trade deficit is found at a maximum in December, whereas it is at minimum in September.

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Keywords: Month Effect, Holy Days Effect, Volatility, Conditional Heteroskedasticity Models, and Trade Deficit.

1. Introduction

Most economic time series exhibit wide swings with respect to a certain mean level and such an effect can be persistent even for long time periods. The volatility clustering of economic time series, in this sense, gives knowledge of high variability seen especially in financial time series such as in exchange rates, inflation, and stock exchanges. An interesting area of

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research related to stock market returns is the presence of calendar effects. The most common calender effects are the day of the week effect, the weekend effect, the January effect, the trading month effect, and the holiday effect. Many papers in finance literature have examined these effects in stock market returns.

The first paper considering the day of the week effect with respect to average returns is by Cross (1973). The findings of this paper demonstrate that average returns on Friday are higher than average returns on Monday. French (1980) analyzed S&P 500 index for the day of the week and found that stock exchange returns are high on Fridays, but low on Mondays. Jaffe and Westerfield (1985) also observed weekend anomaly in the Australia, Canada, Japan and UK markets. French and Roll (1986) point out that return variances for the days following an exchange holiday should be larger than other days. Kim (1988) indicates that stock exchange returns are high before holidays but low for the following days in the US, Australia, UK, Japan, Canada, and Korea. Berument and Kıymaz (2001) examined the volatility of S&P 500 index and reveal the day of the week effect in both the mean and variance equations. Their findings show that the maximum and minimum returns are observed on Wednesday and Monday, whereas the maximum and minimum volatility are observed on Friday and Wednesday, respectively. Kenourgios et al. (2005) investigated the day of the week effect in the Athens Stock Exchange (ASE) using daily observations, and found evidence in favor of the day of the week effect. Ndu (2006) considers the trends in 15 developed financial markets using annual and daily returns and reported negative returns on Mondays for seven countries and on Wednesdays for the others.

A set of empirical studies examines not only the day of the week effect but also the month and Holy Days' effect on the volatility of returns. Rozeff and Kinney (1976) indicate that returns in January are higher than in other months for the New York Stock Exchange (NYSE). Cadsby (1989) found that a negative return exists in the NYSE in October as a regular pattern. Using the months January and October as dummy variables in the conditional mean and variance equations, Glosten et al. (1993) found a negative relationship between the conditional variance of monthly returns and conditional expected returns. Beller and Nofsinger (1998) examine the seasonal volatility in stock exchange returns using generalized autoregressive conditional heteroskedasticity (GARCH) modeling and find differences in volatilities between the months. Coutts et al. (2000) consider both the day of the week effect and the month and weekend effect for the three leading industry indices of the Athens Stock Exchange, and reveal the January effect in these indices. They also found significant holiday anomaly and positive return on Fridays. Husain (1998) analyzed the Pakistani stock market and demonstrated that volatility is significantly lower during the weeks of Ramadan. Oğuzsoy and Güven (2004) investigated the effect of Holy Days (the feast of Ramadan and the feast of Sacrifice) on stock returns at the Istanbul Stock Exchange (ISE) and confirmed high returns two days before religious holidays. Seyyed et al. (2005) examined sector indexes in the Saudi Arabian stock market and did not find any significant effects in average returns during Ramadan. Abadir and Spierdijk (2005) found that anomalies exist around festive times and that index returns tend to be negative before the festivities and positive after the festivities.

This paper seeks to investigate the month and Holy Days' effect (the feast of Ramadan and the feast of Sacrifice) on the volatility of trade balance as a predominant characteristic of the Turkish economy. There are many empirical papers in the Turkish economics literature constructed upon the trade balance and trade deficit, but these papers are concerned in general with the possible reasons of the trade deficit, historical development, and the relationship between the trade balance and the exchange rate[1]. In this paper, an attempt has been made to analyze the month and Holy Days' effect (the feast of Ramadan and the feast of Sacrifice) on Turkey's volatility of trade deficit. For this purpose, the next section is devoted to the course of trade balance in the Turkish economy. Section three discusses the methodological issues in modeling volatility content of the trade balance. Section four presents the estimation results, and the final section provides conclusions.

2. Some Stylized Facts of the Turkish Trade Balance

The Turkish economy has witnessed vital characteristic changes during the post-1980 period. Especially, liberalization of capital account and an ever-increasing openness of the domestic economy to world markets associated with an enlarging trade volume gave rise to the developments in this period. As in many emerging markets and developed countries, the main

[1] Berument and Dinçer (2005) reveal that appreciation of Euro against US Dollar would increase the output in the long run, appreciate the local currency, and improve the trade balance. Brada et al. (1997) find that exchange rate and domestic and world real income have significant effects on the trade balance in the long run.

determinant taking the Turkish economy into this process is the globalization phenomenon.

A turning point in the Turkish economic policy came in January 1980. At the time, import substitute industrialization strategy was replaced by an export-led growth strategy that relies more on the market-based economy. The policies based on adjustments upon tariffs rather than quantity restrictions were adopted, and also protection rates in imports regime were steadily lowered. Besides, export licenses were abolished, and export liberalization was put in effect as a major policy issue in the Turkish economy politics (Varol, 2003). The share of exports and imports within the gross national product (GNP) increased in time. However, the increase was more rapid in the ratio of import /GNP. Thus, Turkey faced a negative balance of trade.

In 1989, capital account liberalization was completed, which enabled domestic residents and private firms to borrow freely in the international financial markets. This allowed residents to make financial transactions in foreign currencies and nonresidents were allowed to invest freely in the domestic markets. Integration into the European Customs Union in 1996 affected the Turkish trade balance predominantly until as recently as 2006.

The 1997 Asian financial crisis and the subsequent Russian crisis in 1998 affected the Turkish trade balance in a negative way because while the exports volume could not be increased, the imports volume maintained a growing trend. And as the anti-inflationary stabilization program based on nominal exchange anchor appreciated the domestic currency, it also deteriorated the trade balance. The stabilization program was unsuccessful in attaining the ex-ante crawling-peg regime leading to the February 2001 Turkish economic crisis. This crisis resulted in a massive depreciation of domestic currency against hard currencies such as the US Dollar and Euro. These developments brought about a narrowing effect upon imports volume, and supported the exports volume in a positive way. However, having a stabilized economy and having attained a sustainable growth path, trade balance again began depreciating until 2006, although policymakers could provide massive increases in exports volume.

3. Methodology

Many papers in the literature examine the day of the week effect by employing the Ordinary Least Squares (OLS) methodology (Cross, 1973; French, 1980; Jaffe and Westerfield, 1985). However, the use of this methodology has two drawbacks. The first is that the errors in the model may be autocorrelated. To prevent this problem, lagged values of the variables must be included in the analysis. The second problem is the residuals change conditional over time. (Berument et al. 2006). Autoregressive conditional heteroskedasticity (ARCH) and generalized autoregressive conditional heteroskedasticity **(**GARCH) models can be employed to remove this problem. In this sense, the ARCH methodology of Engel (1982) assumes the conditional heteroskedasticity (h_t) as a function of past values of the residual squared terms. ARCH (p) model can be expressed as follows:

$$\begin{aligned} h_t &= \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 e_{t-2}^2 + \ldots + \alpha_p e_{t-p}^2 \\ &= \alpha_0 + \sum_{i=1}^{p} \alpha_i e_{t-i}^2 \end{aligned} \tag{1}$$

The GARCH model proposed by Bollerslev (1986) is a popular extension of standard ARCH model. The conditional variance of residual term at time t in the GARCH models depends not only on the squared error term in the previous time period but also on its conditional variance in the previous time period.

$$h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \ldots + \alpha_p e_{t-p}^2 + \beta_1 h_{t-1} + \ldots + \beta_q h_{t-q} \tag{2}$$

This specification is known as GARCH (p, q) modeling. Because both ARCH and GARCH models make the symmetry assumption that deals with the volatility of financial time series, asymmetric conditional heteroskedasticity models are required for considering asymmetry in financial time series. In line with such requirements, exponential generalized autoregressive conditional heteroskedasticity (EGARCH) model of Nelson

(1991) and threshold generalized autoregressive conditional heteroskedasticity (TGARCH) model of Zakoian (1994) are models that attempt to consider the asymmetric conditional heteroskedasticity. EGARCH model can be shown as follows:

$$\ln h_t = \alpha_0 + \beta_1 \ln h_{t-1} + \theta \frac{e_{t-1}}{\sqrt{h_{t-1}}} + \gamma \left| \frac{e_{t-1}}{\sqrt{h_{t-1}}} \right| \quad (3)$$

Here, when the parameter of θ is considerably different from 0, the effect of shocks on volatility is asymmetric. The γ parameter measures the effects of shocks on volatility. The β parameter represents the weight of conditional volatility at specified time.

The conditional variance of TGARCH model is given by:

$$h_t = w + \alpha e_{t-1}^2 + \gamma e_{t-1}^2 D_{t-1} + \beta h_{t-1} \quad (4)$$

where $D_{t-1} = 1$ if $e_{t-1} < 0$ and $D_{t-1} = 0$ if $e_{t-1} > 0$. In this model specification, the γ parameter is used to capture the asymmetrical effect. If $\gamma > 0$ then the leverage effect exists, while if $\gamma \neq 0$ the effects of shocks on volatility is asymmetric.

Dummy type variables can also be included into the abovementioned conditional heteroskedasticity models[2]. Berument and Kıymaz (2003) attempt to test the presence of the day of the week effect by including dummy variables in both mean and variance equations. The equations can be given as in (5) and (6):

$$R_t = \alpha_0 + \alpha_M M_t + \alpha_T T_t + \alpha_H H_t + \alpha_F F_t + \sum_{t=1}^{n} \alpha_i R_{t-i} + e_t \quad (5)$$

[2] Karolyi (1995) includes the volatility of foreign stock returns to explain the conditional variance of home country stock return. Hsieh (1988) includes the day of the week effect in volatility for various exchange rates.

$$h_t = V_c + V_m M_t + V_t T_t + V_h H_t + V_F F_t + V_{j1} e_{t-1}^2 + V_{1b} h_{t-1} \qquad (6)$$

where R_t represents returns on a selected index; M, T, H, and F are the dummy variables for Monday, Tuesday, Thursday, and Friday, respectively.

Similarly, the month and Holy Days (the feast of Ramadan and the feast of Sacrifice) dummies in both conditional mean and variance equations were included in this study to investigate month and Holy Days' effects on the volatility of trade deficit. For this purpose, three separate models are estimated. The first one is the standard OLS equation that includes month and Holy Days (the feast of Ramadan and the feast of Sacrifice) dummies. The second represents the conditional mean equation that includes these dummy variables, and the last model represents both the conditional mean and variance equations including these dummies. ARCH (1) model using 12 dummy variables considered in this paper can be shown as equations (7) and (8) given below:

$$\begin{aligned} TDV_t &= \phi_0 + \phi_1 D_1 + \phi_2 D_2 + \phi_3 D_3 + \phi_4 D_4 + \phi_5 D_5 + \phi_6 D_6 + \phi_7 D_7 \\ &+ \phi_8 D_8 + \phi_9 D_9 + \phi_{10} D_{10} + \phi_{11} D_{11} + \phi_{12} D_H + \sum_{i=1}^{n} \alpha_i TDV_{t-i} + e_t \end{aligned} \qquad (7)$$

$$\begin{aligned} h_t &= \alpha_0 + \alpha_1 e_{t-1}^2 + \delta_1 D_1 + \delta_2 D_2 + \delta_3 D_3 + \delta_4 D_4 + \delta_5 D_5 + \delta_6 D_6 + \delta_7 D_7 \\ &+ \delta_8 D_8 + \delta_9 D_9 + \delta_{10} D_{10} + \delta_{11} D_{11} + \delta_{12} D_H \end{aligned} \qquad (8)$$

where TDV_t is the trade deficit value in period t, the variables from D_1 to D_{11} are the 11 monthly dummies, D_H is the Holy Days dummy (the feast of Ramadan and the feast of Sacrifice)[3], e_t is the white-noise error term, and h_t

[3] The main festivities occur every year at a different time of the calendar. The feast of Ramadan is celebrated over a holiday period of three-and-a-half days and the feast of Sacrifice is celebrated over a holiday period of four-and-a-half days. The

is the conditional variance. The D_H dummy variable is constructed to analyze the impacts of the feast of Ramadan and the feast of Sacrifice on trade deficit. The variable takes on a value of 1 for the months of Holy Days and 0 otherwise. The constant term of the conditional variance equation can vary from month to month by using such model specifications. Dummy variables in this estimation process not only model conditional dependence between mean and variance but also deal with monthly effects in the conditional forecasts. To deal with potential model misspecification, Bollerslev and Wooldridge (1992) propose using the Quasi-Maximum likelihood method that produces robust t-ratios.

4. Empirical Results

The analysis uses monthly seasonally unadjusted data obtained from the International Monetary Fund's (IMF) International Financial Statistics (IFS) database covering the 1984:01–2006:12 periods. For this model specification, the trade deficit values (TDV) are calculated as a difference between total exports and imports. Export data are valued as "free on board" or "FOB"; import data including Cost, Insurance, and Freight or "CIF" (millions of dollars) are also valued. The trade deficit data for US inflation is adjusted by dividing by the consumer price index (2000=100) thus generating real data. The descriptive statistics for the trade deficit can be observed in Table 1 given below.

Table 1: Descriptive Statistics of Monthly Trade Deficits

	Mean	**Median**	**Maximum**	**Minimum**	**Std.Dev.**	**Skewness**	**Kurtosis**
TDV	-13.3743	-10.5423	-0.0285	-45.5901	9.7384	-1.0782	3.6284
JANUARY	-9.3787	-7.7707	-2.0811	-26.1230	6.4975	-0.9585	3.0927
FEBRUARY	-10.2843	-6.7709	-1.4033	-32.4171	8.0227	-1.2099	3.8019
MARCH	-12.4650	-9.2992	-1.7600	-36.0447	9.7407	-1.0948	3.1696
APRIL	-13.4051	-12.6968	-3.5822	-43.9131	10.2504	-1.3065	4.4948

Turkish Government decrees the remaining day of the week as a holiday. These holidays clearly affect retail trade, production, and financial markets.

MAY	-14.3978	-12.2571	-2.3833	-45.5902	10.9145	-1.2026	4.0785
JUNE	-14.3451	-11.1613	-3.1527	-37.1280	9.9036	-0.8818	2.8192
JULY	-14.7283	-12.3194	-0.32491	-38.0258	9.7769	-0.7838	2.9167
AUGUST	-15.0893	-13.3743	-2.0026	-43.8989	11.7006	-1.0752	3.4244
SEPTEMBER	-13.4054	-12.2888	-1.6669	-38.8790	9.6368	-0.9602	3.4062
OCTOBER	-12.6098	-10.6369	-1.5300	-37.8790	9.3249	-1.1213	3.6244
NOVEMBER	-14.0916	-12.9999	-2.0583	-35.7476	9.1117	-0.9590	3.0717
DECEMBER	-16.5184	-17.0192	-0.0285	-38.8177	11.0736	-0.3929	2.1980

As seen from Table 1, the mean of the trade deficit is at maximum in December, whereas it is minimum in January. Besides, the maximum and minimum standard deviations representing the degree of volatility occur in August and January, respectively.

In analyzing the time series properties of the variable TDV, the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests are applied. According to the findings in Table 2, the unit root null hypothesis is rejected in this study.

Table 2: ADF and PP Unit Root Tests

	ADF TEST	PP TEST
TDV	-3.1542 [c]	-6.9743 [a]

[a, c] indicate the significance in %1 and %10 probability levels.

Additionally, the Beaulieu and Miron (1993) version of the Hylleberg, Engle, Granger, and Yoo (HEGY, 1990) seasonal unit root test is applied to make an accurate inference about the seasonal movements in the monthly trade deficit data. HEGY test is the general procedure that can test for unit roots at some seasonal frequencies without maintaining that unit roots are present at all seasonal frequencies. In this way, it is possible to avoid the a-priori application of the ($1-L^{12}$) filter, which supposes unit roots at all seasonal frequencies (L is the lag operator). Beaulieu and Miron (1993) derive the asymptotics of HEGY's procedure for monthly data and use

Monte Carlo Methods to compute the finite sample critical values[4]. To test hypothesis about various unit roots, the following equation by OLS is estimated.

$$(1-L^{12})TDV_t = \alpha + \sum_{k=1}^{11} m_k M_{kt} + \delta t + \sum_{i=1}^{p} \phi_i \Delta_{12} TDV_{t-i} + \sum_{k=1}^{12} \pi_k TDV_{k,t-1} + \varepsilon_t \quad (9)$$

The estimation equation includes a constant (α), 11 seasonal dummies (M), a trend (t), and 12 lags of the dependent variable. The Akaike (1974) criteria are also used to determine the number of lags. The estimation period yields a sample size of 276; therefore, critical values for 12T=240 are used. The seasonal unit root test results for monthly data are reported in Table 3[5].

Table 3: HEGY Seasonal Unit Root Test Results

Frequency	Null Hypothesis	Test Statistic
0	$\pi_1 = 0$	-4.22[a]
π	$\pi_2 = 0$	0.67
$\pi/2$	$\pi_3 = 0$	3.63
	$\pi_4 = 0$	-1.54[c]
$2\pi/3$	$\pi_5 = 0$	-4.88[a]
	$\pi_6 = 0$	1.03

[4] Beaulieu and Miron provide the asymptotic distribution of the statistics necessary to perform the test: t_1, t_2, t_k, and t_{k+1}, where k∈ (3,5,7,9,11). They write that k is "odd" if k ≠ 1 and that k is "even" if k ≠ 2 and k∈ (4,6,8,10,12). They prove that the asymptotic distribution of five t_k is the same as those of the five t_{k+1}.

[5] For frequencies 0 and π, the corresponding t-statistic for $\pi_k = 0$ (k=1,2) against the alternative that $\pi_k < 0$ (stationarity) has to be tested. For the other frequencies ($\pm\pi/6, \pm\pi/3, \pm\pi/2, 2\pi/3, \pm 5\pi/6$), the null hypothesis $\pi_k = 0$ (i=3,4,..12) implies the presence of unit roots. The critical values of the t statistic are based on the paper by Beaulieu & Miron (1993). They also obtain an F statistic to test the joint hypothesis $H_0 = \pi_k = \pi_{k+1} = 0$ (k=3,5,7,9,11), for the presence of unit roots in all the seasonal frequencies.

$\pi/3$	$\pi_7 = 0$	-2.14
	$\pi_8 = 0$	2.94
$5\pi/6$	$\pi_9 = 0$	-3.53[b]
	$\pi_{10} = 0$	-2.52
$\pi/6$	$\pi_{11} = 0$	-3.99[a]
	$\pi_{12} = 0$	2.34
$\pi/2$	$\pi_3 = \pi_4 = 0$	7.23[b]
$2\pi/3$	$\pi_5 = \pi_6 = 0$	12.31[a]
$\pi/3$	$\pi_7 = \pi_8 = 0$	6.69[c]
$5\pi/6$	$\pi_9 = \pi_{10} = 0$	11.86[a]
$\pi/6$	$\pi_{11} = \pi_{12} = 0$	14.43[a]

[a, b] and [c] indicate rejection of the null hypothesis at 1, 5, 10 percent levels.
* All critical values are from Beaulieu & Miron (1993).

The t-statistics in the Table 3 indicate that the data rejects the unit root null hypothesis at frequencies zero (long run), $\pi/2$, $2\pi/3$, $5\pi/6$, and $\pi/6$. At the same time, according to the F statistics in the table, the unit roots at frequencies $\pi/2$, $2\pi/3$, $\pi/3$, $5\pi/6$, and $\pi/6$ are rejected. However, the existence of a unit root at the π frequency cannot be rejected. As discussed in Beaulieu and Miron (1993), tests at this frequency have lower power than tests for seasonal unit roots at other frequencies. Therefore, the trade deficit data can be considered stationary around a deterministic trend and in the seasonal frequencies.

The time series properties of the variable TDV were examined and the three model specifications were estimated. These are as follows: The first one is OLS equation including monthly dummies and Holy Days dummy. The second one is an EGARCH (1,1) model including dummy variables in conditional mean equation, whereas the last one is an ARCH (1) model that includes monthly dummies and holy days dummy in both conditional mean and variance equations. The estimation results are given below in Table 4.

Table 4: Modeling the Turkish Trade Deficit

Dependent Variable:	TDV					
Mean Equation		OLS		EGARCH(1,1)		ARCH(1)
Intercept		-21.999[a]		-24.892[a]		-24.521[a]
D_1		5.830[a]		7.051[a]		5.520[a]
D_2		5.007[a]		5.422[a]		5.118[a]
D_3		2.825[a]		1.886[b]		2.855[a]
D_4		1.983[b]		2.351[a]		1.836
D_5		1.187		1.559[c]		1.461
D_6		1.335[c]		2.594[a]		1.034
D_7		1.145		1.804[b]		0.977
D_8		1.202		2.095[b]		1.063
D_9		2.951[a]		3.376[a]		2.632[b]
D_{10}		3.735[a]		4.373[a]		3.586[a]
D_{11}		2.240[a]		2.365[a]		2.255[c]
D_H		2.306[a]		2.784[a]		1.687[a]
AR (1)		0.983[a]		0.980[a]		0.981[a]
MA (1)		-0.448[c]		-0.453[a]		-0.495[a]
MA (3)		-		-		0.097[b]
Variance Equation						
α_0		-		4.195[a]		31.747[a]
α_1				-		0.217[c]
γ		-		0.158[a]		-
θ		-		0.127[a]		-
β_1		-		-0.953[a]		-
D_1		-		-		-21.522[b]
D_2		-		-		-22.254[b]
D_3		-		-		-27.737[a]
D_4		-		-		-21.377[b]
D_5		-		-		-26.027[b]
D_6		-		-		-28.556[a]
D_7		-		-		-28.592[a]
D_8		-		-		-21.846[b]
D_9		-		-		-29.216[a]
D_{10}		-		-		-26.218[a]
D_{11}		-		-		-27.167[a]
D_H		-		-		0.671
Diagnostic Test Results						
Q-Stat (6)		10.749		11.252		8.069
Q-Stat (12)		25.135		29.937		17.328
ARCH-LM (6)		12.584		-		-
ARCH-LM (12)		21.618		-		-
Jarque Bera		58.841		-		-

- The t-critical values for OLS estimation are obtained from 10.000 bootstrap replications. [a], [b], and [c] indicate the significance of the relevant variable at %1, %5, and %10 probability levels, respectively.

The first column reports the results from the OLS estimation. In this model, the Jarque-Bera test statistics easily rejects the null hypothesis of a normal distribution. The lack of normality would make standard calculation of p values and confidence intervals invalid. To investigate the results in more detail for OLS estimation, an additional simulation study is constructed and the bootstrap[6] critical values for the t-statistics using 10.000 bootstrap replications are generated. The findings show that by using the bootstrapped critical values for the t-statistics, all the parameters except for the dummies D_5, D_7, and D_8 are estimated as statistically significant. Estimation results of the OLS equation indicate that the trade deficit is at maximum in December (-21.999), which is taken over completely by the constant term, and is minimum in January (5.830). These findings are consistent with the descriptive statistics in Table 1. The sum of the autoregressive coefficients in this model is lower than unity, verifying that the model satisfies the stability (stationarity) condition. The feast of Ramadan and the feast of Sacrifice have significant effects on trade deficit. The Ljung-Box Q statistics with both 6 and 12 lags is performed and the null hypothesis of the existence of autocorrelation is rejected. The Engel's (1982) ARCH LM test for 6 and 12 lags indicates the presence of the ARCH effect; the null hypothesis of constant variance is rejected. Therefore, the conditional variance equation is modeled as an EGARCH (1, 1) process and the mean equation is reestimated with the conditional variance equation.

In the EGARCH (1,1) model that includes the monthly dummies and Holy days dummy in the conditional mean equation, all the parameter estimates are statistically significant. The maximum value of trade deficit (-24.892) occurs in December, whereas the minimum value of trade deficit (7.051) comes from the month of January. It is seen from the table that the stability condition is satisfied. All these findings support the estimation results given so far. In the conditional variance equation, the estimated coefficient value for asymmetry parameter θ is significant, which means that the volatility of trade deficit is asymmetric. By using the Ljung-Box Q statistics for 6 and 12 lags, the null hypothesis that the residuals are autocorrelated can be rejected. Thus, specification problems in this model are not encountered.

[6] Bootstrap provides an alternative approximation for the asymptotic distribution of a statistic.

Finally, the Modified ARCH model including dummy variables in both the mean and variance equations[7] are estimated. The estimated coefficients for monthly dummies except April, May, June, July, and August are statistically significant in the mean equation. Supporting the above findings, December (-24.521) is the month when trade deficit is maximum as can be seen from the intercept parameter, whereas the minimum trade deficit occurs in January (5.520). In the conditional variance equation, α_1 represents the lagged value of the squared residual term and it is statistically significant and positive to satisfy the nonnegativity of conditional variance. The volatility of trade deficit is at maximum in December (31.747), and is minimum in September (-29.216). It is clear that Holy Days dummy in variance equation is statistically significant. On the basis of the Ljung-Box Q statistics test for 6 and 12 lags, the null hypothesis of the existence of autocorrelation is rejected. It can be seen that there is no evidence of misspecification for this model.

Summarizing, three different model types for the mean and variance equations are employed. The findings show that the monthly dummies for the mean and variance equations are statistically significant. The Holy Days dummy is statistically significant for the mean equation. Thus, it is concluded that the month effect on both mean and variance are present in the trade deficit of Turkey. Trade deficit is found to be at maximum in December and minimum in January and the volatility content of trade deficit is maximum in December and minimum in September.

5. Concluding Remarks

By using three separate models, this paper investigates and estimates the month effects and Holy Days' effect (the feast of Ramadan and the feast of Sacrifice) on the volatility of trade deficit in Turkey during the period of 1984:01–2006:12. The estimation results indicate that the trade deficit seems to be maximum in December and minimum in January. Additionally, volatility of trade balance is found to be maximum in December and minimum in September. These findings indicate that the extreme values in the trade balance in the month of December should be of special interest for

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[7] By applying other types of conditional variance models it is found that the additional terms and asymmetry parameters (θ) are statistically insignificant. Hence, ARCH (1) is considered to be the appropriate representation.

the policymakers. Another finding of this paper is that Holy Days' effect is significant on the Turkish trade deficit but not on the volatility.

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