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Public Information Arrival and the Fisher Effect in Emerging Markets: Evidence from Stock and Bond Markets in Turkey

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Abstract

We examine the impact of inflation on nominal stock returns and interest rates in Turkey's emerging economy, which has a moderately high, persistent, and volatile inflation rate. Empirical evidence indicates that Turkey's inflation increased more than nominal stock returns and interest rates, implying that real returns to investors declined during our sample period. Among the different sector indexes we study, the financials sector serves as the best hedge against expected inflation, and the Fisher effect appears to hold only for this sector. We also find that public information arrival plays an important role, especially in the stock market.

Key words: Fisher effect, public information arrival, emerging markets, financial services, volatility, asymmetric GARCH models.

1. Introduction

A surprising finding in the financial economics literature is that nominal stock returns and inflation are negatively related. This finding implies that the stock market is not a true hedge against inflation and that the Fisher effect, which implies a one-to-one direct relation between asset returns and expected inflation, does not hold. The results of earlier studies suggest that the debate on the issue of whether nominal asset returns move proportionally with the expected inflation rate—(so that expected real returns are independent of the expected inflation rate) is not yet settled. The empirical evidence, at least for the United States, appears inconclusive.

Our study provides evidence from a high-inflation economy, Turkey. In the last decade, the inflation rate in Turkey has ranged between 50% and has been persistent and volatile. To the best of our knowledge, ours is only the first study to examine the validity of the

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Fisher effect for the Turkish stock market.[1] In an earlier study, Berument and Malatyali (2001) test the Fisher effect for the Turkish bond market. We examine both stock and bond markets to obtain extensive evidence on the effect of anticipated inflation on financial market returns. Given the conflicting evidence on the link between bond returns and inflation in industrial countries such as the United States, we believe it is useful to provide fresh evidence from an emerging market like Turkey. Studying the bond market also allows us to compare our evidence, based on an updated sample period and a different empirical specification, with the earlier study by Berument and Malatyali (2001). In addition to using a composite stock index, we examine the impact of inflation on two key individual sector indexes in Turkey: financials and industrials. By doing so, we can provide more complete evidence on which sector (if any) serves better as a hedge against inflation.

In addition to the consumer price index (CPI), we also investigate the impact of two other key macroeconomic announcements on asset returns: exports and international reserves. The export sector is critically important to the Turkish economy. Movements in international reserves could capture monetary policy actions. They might also reflect changes in capital inflows, including changes in workers' remittances, a significant source of external revenue for the Turkish economy. Due to a lack of long-term data for tourism, construction, and other key sectors of the economy, we are not able to include such macroeconomic data in our estimations. Although data on industrial production is available, we do not include it here, because it is highly correlated (more than 50%) with exports. This choice ensures that we do not commit any bias in estimations by including highly collinear data.

In emerging markets such as Turkey's, the periodic arrival of public information may affect market volatility by bringing about frequent revisions in investors' expectations for future economic trends. Such information may cause investors to adjust their risk and return expectations, thus inducing portfolio revisions and price volatility. Furthermore, a better understanding of the impact of such public information on emerging market activity permits us to make further inferences about the usefulness of public information on asset price determination in such markets.

Evidence on the impact of public information arrival on emerging asset returns, typically measured by the publicly released economic and financial data, is mixed. The recent debate has concentrated more on the usefulness of public compared to private information, but the available empirical evidence has been based on only a few industrial countries. Thus, our results providing evidence on the role of public information in asset price movements from an emerging market like Turkey contribute to this line of debate as well.

1 Empirical studies on Turkish stock market are scant. Başci et al. (1996) examine the link between stock prices and trading volume dynamics, while Muradoğlu and Metin (1996) investigate efficiency issues. Gürsoy and Aydoğan (2002) study the ownership structure of firms listed in the Istanbul Stock Exchange. Güner and Önder (2002) examine the volatility of daily stock returns. Demirer and Karan (2002) investigate the day-of-the-week effects, while Yüksel (2002) studies the performance of the Istanbul Stock Exchange during the Russian crisis.

In section 2, we discuss the motivation of our empirical analysis by providing background information on previous empirical studies and the macroeconomic environment in Turkey. In section 3, we develop our empirical model. Section 4 describes data used and provides descriptive statistics. We report our empirical findings in section 5. In section 6, we report some robustness tests results. Section 7 concludes with a discussion of the policy implications of our findings.

2. Previous studies and motivation of the study

2.1. Previous studies

Our study relates to two distinct, albeit related, literatures on the Fisher effect and the role played by macroeconomic announcements in financial markets. Earlier studies find that the Fisher effect does not well. Many studies offer reasons for this unexpected finding.[2] Choudhry (2001) argues that the Fisher effect should hold better for high inflation countries because in these economies, the response of asset prices to inflation news would be relatively larger. He provides evidence from four high inflation Latin and Central American economies and reports a positive relation between stock market returns and inflation, suggesting that the Fisher effect does indeed hold better for these economies.

While earlier studies have focused on hyperinflation countries (e.g., Choudhry, 2001), this paper focuses on an economy with moderate but persistent inflation. It is possible that the evidence for the Fisher effect is sensitive to include other key macroeconomic data in estimations. Earlier studies that examine the validity of the Fisher effect do not use such data (see, for example, Gültekin, 1983; Choudhry, 2001). We view the inclusion of public announcements in the empirical models of asset returns as control variables.

There have also been several studies that examine the link between bond returns and inflation, but their focus has been on the United States and other developed economies. For example, using Treasury bill and government bond data for the 1951–1971 period, Fama and Schwert (1977, 1979) report evidence that both bonds and bills are complete hedges against expected inflation. However, using one-month and three-month Treasury bill data for the period of 1964–1987 and 1954–1987, respectively, Evans and Watchel (1992) find no evidence for the Fisher effect. These authors argue that the lack of empirical evidence for the Fisher effect can be at least partially explained by the volatile real interest rates, which were higher during early 1970s but declined in 1980s. Evans and Lewis (1995) reach an opposite conclusion, arguing that the Fisher hypothesis holds well, once research incorporates structural shifts in the inflation process in the estimation. Berument and Malatyali (2001) is the only study that examines the Fisher effect for Turkey, using bond data. They find that the nominal interest rate increases less than inflation, suggesting that the real interest rate decreases with higher inflation.

Our study is also related to the literature on the effect of macroeconomics announcements. Although many studies show that macroeconomic announcements are

2 Sellin (2001) provides a review of the literature.

one of the major determinants of short-term volatility of asset returns in developed financial markets (Jones et al., 1994; Fleming and Remolona, 1999), there is little evidence on this issue from emerging markets. Earlier studies find that publicly released variables are related to changes in current and future asset returns. For example, Hardouvelis (1988) and Karfakis and Kim (1995) provide evidence drawn from exchange rate returns. Berry and Howe (1994) focus on stock returns, while Jones et al. (1994) study bond market effects. However, more recent studies have yielded more mixed results. Ito et al. (1998) examine the importance of private information in the Tokyo foreign exchange market and find significant effects whereas Melvin and Yin (2000) and Edmonds and Kutan (2002) report significant public announcement effects in the yen/dollar market and Nikkei stock market, respectively. Reacting to these conflicting findings, researchers have turned to new asset pricing models that emphasize the role of private information. For example, microstructure theories stress the importance of how private information, revealed through order flows, affects asset prices (O'Hara, 1995, provides a review of microstructure theories).

2.2. Motivation

The Turkish economy has been subject to persistent high inflation rates since the 1970s. Following the sharp increase in world oil prices in 1973–1974, the inflation rate started rising, reaching triple-digit levels by 1979. In response to a weakening economy and a growing level of political instability, the Turkish government introduced major reforms in January 1980. In response to the reforms, the inflation rate declined somewhat during 1981–1987, but then rose again to higher levels in 1990. The rise was due to frequent nominal exchange rate depreciations, which had been designed to promote stagnating exports.

During the 1990s, due to lack of fiscal discipline and an overvalued domestic currency, the Turkish economy continued experiencing high inflation. These factors finally brought about a major economic crisis in 1994, causing the inflation rate to reach a peak level of 106.3% in the same year (table 1). Following the crisis, to stabilize exports, monetary policy authorities attempted to align the rate of nominal exchange-rate depreciation with inflation to maintain a more stable real exchange rate. Although the inflation rate declined following the crisis, it has persisted at moderately high levels, ranging between 55% and 65% (Table 1). It appears that promoting exports has been a major economic goal of policymakers since the 1970s, superseding the goal of having a stable, low inflation-rate environment.

Besides frequent nominal exchange-rate depreciations, a growing level of budget deficits has been another key factor responsible for the high-inflation rate in Turkey. The deficit in domestic currency increased from 11.8 trillion in 1990 to 14,262.8 trillion in 2000. Such a high-inflation environment has caused higher nominal interest rates, resulting in the dollarization of the economy. The dollarization ratio, which was only 5.4% in 1990, reached approximately 20% in 2000 (Table 1).

Turkey's inflation experience is distinctive because, although many countries, particularly those in South America, have experienced hyperinflation, in the last decade

Table 1. **Key macroeconomic indicators of the Turkish economy**
The inflation rate is computed using the consumer price index (CPI) while the nominal interest rate refers to the interbank money market rate. The budget deficit is in national currency (trillions). The dollarization ratio is given by the ratio of foreign currency deposits to the gross domestic product (GDP). All data, which are obtained from IMF's International Financial Statistics, CD ROM tape, March 2002, are expressed in percentage terms per annum.

Year	Inflation	Budget Deficit	Interest Rate	Dollarization Ratio
1990	60.3	– 11.8	51.9	5.4
1991	66.0	– 33.3	72.7	7.8
1992	70.1	– 47.3	65.4	9.2
1993	66.1	– 133.1	62.8	9.1
1994	106.3	– 150.8	136.5	14.5
1995	88.1	– 316.6	72.3	14.8
1996	80.3	– 1238.1	76.2	16.3
1997	85.7	– 2444.1	70.3	17.0
1998	84.6	– 4387.0	74.6	16.2
1999	64.9	– 10075.8	73.5	21.9
2000	54.9	– 14262.8	56.7	19.5

some of these countries have been able to lower inflation to moderate levels and even single-digits. Although Turkey never experienced hyperinflation, it has not been able to bring its moderate level of inflation rate down. Instead, inflation has been persistent at moderately high and volatile levels for about two decades.[3] At the same time, Turkey has faced an unstable political and financial environment.

Table 2 provides evidence for inflation rates in Turkey and selected Latin American countries. Argentina, which experienced hyperinflation rate in 1990 with an inflation rate of 2314%, was able to bring down inflation to single digits in the mid-1990s. In 2000, the

Table 2. **Inflation rates in Turkey and selected Latin American countries**
Inflation rates are based on the CPI, which are obtained from the IMF's International Financial Statistics, CD ROM tape, March 2002. They are expressed in percentages per annum.

Year	Turkey	Argentina	Chile	Mexico	Venezuela
1990	60.3	2314.0	26.0	26.7	40.7
1991	66.0	171.7	21.8	22.7	34.2
1992	70.1	25.0	15.4	15.6	31.4
1993	66.1	10.7	12.7	9.8	38.1
1994	106.3	4.2	11.4	7.0	60.8
1995	88.1	3.4	8.2	35.0	60.0
1996	80.3	0.2	7.4	34.4	99.0
1997	85.7	0.5	6.1	20.6	50.0
1998	84.6	0.9	5.1	16.0	35.8
1999	64.9	– 1.2	3.3	16.6	23.6
2000	54.9	– 0.9	3.8	9.50	– 98.8

3 We would like to thank Haluk Ünal for his suggestions to make this distinction more clear.

Argentinean economy actually experienced deflation (Table 2). We observe a similar pattern for other Latin American countries. In Venezuela, the inflation rate has continuously declined since mid-1995 and turned into deflation in 2000. On the other hand, Turkey had moderately high inflation levels. The jump in the inflation rate in 1994 was due to the financial crisis, resulting from the collapse of the exchange rate regime in the late 1994.

Some researchers argue that a fundamental reason for the failure of disinflation programs in Turkey is the lack of confidence in the willingness of Turkish policymakers to reduce inflation (Akyürek and Yenigün, 1999). Thus, examining the relation between asset returns and inflation in Turkey provides a natural laboratory in which to test the effectiveness of different financial instruments as a hedge against inflation in a politically unstable economy with a moderately high and volatile inflation rate, such as Turkey.

3. Empirical model

Berument and Malatyali (2001) use a symmetric GARCH model to examine the significance of inflation on bond returns in Turkey. To examine the effect of the CPI and other key public macroeconomic announcements on stock market returns and interest rates, we follow their approach, but with a significant difference. We use asymmetric GARCH models, which not only allows us to capture the time-varying properties of daily financial returns, but also to account for potential asymmetric shocks to the volatility of returns.[4] We propose to utilize the following EGARCH (1,1) specification for stock returns:

$$R_t^i = \alpha_0 + \alpha_1 R_t^{US} + \alpha_2 \Pi_{t-1} + \alpha_3 \Pi_{t-2} + \alpha_4 \Pi_{t-3} + e_t, \tag{1}$$

$$e_t \sim N(0, h_t), \tag{2}$$

$$\log h_t = \beta_0 + \beta_1 \log h_{t-1} + \beta_2 \left| \frac{\varepsilon_{t-1}}{h_{t-1}} \right| + \beta_3 \frac{\varepsilon_{t-1}}{h_{t-1}} + \beta_4 (\text{volume})_t, \tag{3}$$

where R_t^i represents individual (composite and sector index) stock returns and Π denotes inflation.

Equation (1) is the mean equation, modeling nominal stock returns as a function of movements in a world market index and anticipated inflation. The movements are captured by the U.S. Standard & Poors 500 (S&P500) index, R_t^{US}. Following earlier studies (Firth, 1979; Gültekin, 1983; Choudhry, 2001), expected inflation is represented by past rates of inflation in the months of $t-1$, $t-2$, and $t-3$.

Equation (2) indicates that financial market volatility is time varying. Equation (3) is the conditional variance equation that captures the time-varying volatility, In equation (3), the constant term, β_0, represents the unconditional variance and β_1 captures the

4 Campbell and Hentschel (1992) and Braun et al. (1995) report evidence that monthly U.S. stock returns exhibit time-varying volatility, while Beakert and Harvey (1997), Aggarwal et al. (1999) and Davis and Kutan (2002), among others, provide international evidence from emerging stock markets.

effect of past conditional variance on today's variance. In this model, good news $(\varepsilon_t < 0)$ and bad news $(\varepsilon_t > 0)$ have a differential effect on the conditional variance, which can be tested by the hypothesis that $\beta_3 < 0$. If this parameter is not equal to zero, then the news has an asymmetric impact. The log of the conditional variance implies that the asymmetric effect is exponential. Trading volume data are expressed in percentage terms and measured by the number of shares. In estimations, we also include lagged returns in the mean equation to remove potential serial correlation in the dependent variable.

We perform estimations by using Eviews4. Eviews uses OLS estimates as the initial values and uses the Marquardt optimization method. In estimations, we also use the Gauss–Newton method to check the sensitivity of the results to using different optimization techniques. However, the results do not change qualitatively. When Eviews cannot find a step size that improves the objective function, it issues an error message. Error messages occurred when we included inflation and other macroeconomic variables together in the conditional variance equation. To achieve global maxima, we omitted these variables in the variance equation.

We estimate a similar model for the bond market. Due to the lack of data, we exclude the volume data in the conditional variance equation. Preliminary investigation indicated that the threshold GARCH (TGARCH) model fits the data better than the EGARCH one. Therefore, we conducted estimations for this market using this specification with the following modified conditional variance equation as:

$$h_t = \beta_0 + \beta_1 h_{t-1} + \beta_2(\varepsilon_{t-1}^2 * d_{t-1}) + \beta_3 \varepsilon_{t-1}, \tag{4}$$

where the dummy variable d_t is equal to one when ε_t is less than zero. Good news $(\varepsilon_t < 0)$ and bad news $(\varepsilon_t > 0)$ have different effects on the conditional variance. Bad news has an impact of β_3, while the effect of good news is measured by $\beta_2 + \beta_3$. If β_2 is statistically significant, then the news impact is asymmetric.

4. Data and descriptive statistics

Data for the composite index are available from December 1986 to March 2001. For the industrial and financial sectors, the available data start in January 1991. We obtain the stock market (closing) prices from the website of the Central Bank of Republic of Turkey (CBRT), www.cbrt.gov.tr. Table 3 provides a description of each stock index. Interest rate data are based on average interest rates for the Treasury auctions and the average maturity dates for these auctions. The data are available from January 1987 to December 2000 and can be downloaded from the Treasury Department.[5] Data for the S&P500 index are downloaded from the Yahoo finance, http://finance.yahoo.com. We obtain data on the CPI, exports, and international reserves from the IMF's International Financial Statistics CD ROM data base.

5 We would like to thank Hakan Berument for generously sharing the data with us.

Table 3. **Descriptive statistics**

Composite (National-100) index is composed of national market companies except investment trusts. The financials index is composed of banks, insurance, leasing and factoring companies, and holding and investment companies. The industrials index is composed of food, beverage, textile, leather, wood, paper, printing, chemical, petroleum, plastic, non-metal mineral products, basic metal, metal products and machinery companies. Average interest rates for the Treasury auctions and the average maturity dates for these auctions are used in estimations and obtained from the Treasury Department. Stock returns are computed as the log-differenced price index, multiplied by 100. Inflation is computed using the CPI. Exports and international reserves are expressed in growth rates. All data are in monthly percentage terms, except interest rates. For consistency, a common sample period, from January 1991 to March 2001, is used for all the variables.

Variable	Mean	Standard Deviation	Skewness	Kurtosis
Composite	4.46	14.34	– 0.02	3.22
Industrials	4.24	14.18	0.01	3.67
Financials	4.77	15.84	– 0.07	3.32
S&P500	1.17	3.96	– 0.72	4.94
Interest rate	103.75	41.24	1.98	11.91
Inflation	4.63	3.06	3.05	23.10
Exports	0.64	9.77	– 0.30	4.31
Int. reserves	0.92	9.21	– 0.38	4.32

Table 3 reports the statistics for nominal stock market returns, interest rates, and the macroeconomic variables. The results for the stock market indicate that the financials-sector returns are relatively higher than those of the industrials. This evidence reflects the higher standard deviation of financial returns relative to the industrials. Although domestic stock returns are much higher than are their world counterparts, their standard errors are proportionally larger, too, indicating that the Turkish stocks are relatively riskier than the world stocks. The results for kurtosis and skewness indicate the non-normality of the stock returns and interest rates. To account for the non-normality of returns, in all the estimations that follow we use the robust standard errors developed by Bollerslev and Wooldridge (1992).

5. Empirical results

Table 4 reports the results for the nominal stock market returns. In table 4, all estimated mean equations include a sufficient number of lagged dependent variables to remove serial correlation in returns. We also include several (0, 1) dummy variables to account for the impact of recent global financial crises in Asia and Russia, the recent banking crises in Turkey, and the major Turkish earthquake. For space considerations, these results are not reported, but they are available on request from the authors.

The results for the mean equation indicate that the U.S. market movements do not significantly affect any of the domestic stock returns. In predicting inflation, we find that past inflation rates do not influence composite nominal stock returns. For the industrials and financials, inflation in month $t-2$ has a significant and positive impact, but it is marginal at the 10% level. We note that the effect of inflation on the financials' returns is

Table 4. **Impact of anticipated inflation on nominal stock returns: EGARCH(1, 1) model estimates** R_t^i refers to individual sector and composite index returns. Global market developments are captured by returns on the U.S. S&P500 index, R_t^{US}. The sample period for the composite returns runs from December 1986 to March 2001, while for the industrials and financials, data begin in January 1991 and end in March 2001. $\Sigma\Pi$ is the sum of the estimated inflation coefficients. Volume data are measured by the percentage growth rate in the number of shares. *P*-values are in parentheses. The *P*-value for $\Sigma\Pi$ refers to the null hypothesis that the sum of the anticipated inflation coefficients is zero.

$$R_t^i = \alpha_0 + \alpha_1 R_t^{US} + \alpha_2 \Pi_{t-1} + \alpha_3 \Pi_{t-2} + \alpha_4 \Pi_{t-3} + \varepsilon_t$$

$$\varepsilon_t \sim N(0, h_i),$$

$$\log h_t = \beta_0 + \beta_1 \log h_{t-1} + \beta_2 \left| \frac{\varepsilon_{t-1}}{h_{t-1}} \right| + \beta_3 \frac{\varepsilon_{t-1}}{h_{t-1}} + \beta_4 (\text{volume})_t,$$

	Composite	Industrials	Financials
α_0	−0.67 (0.82)	−0.40 (0.91)	−4.36 (0.20)
α_1	0.25 (0.14)	−0.00 (0.99)	0.18 (0.46)
α_2	−0.55 (0.23)	−0.78 (0.14)	−0.52 (0.28)
α_3	0.63 (0.13)	0.99 (0.00)***	1.68 (0.00)***
α_4	0.20 (0.61)	0.54 (0.11)	0.30 (0.43)
$\Sigma\Pi$	0.28 (0.66)	0.75 (0.31)	1.46 (0.04)**
β_0	2.28 (0.00)***	2.80 (0.02)**	2.16 (0.01)***
β_1	0.46 (0.00)***	0.38 (0.10)*	0.52 (0.00)***
β_2	0.45 (0.00)***	0.36 (0.10)*	0.32 (0.10)*
β_3	0.31 (0.00)***	0.26 (0.05)**	0.36 (0.00)***
β_4	0.02 (0.00)***	0.01 (0.35)	0.01 (0.37)
Log-likelihood	−695.7	−487.0	−495.4
Q (12)	9.66 (0.56)	12.43 (0.33)	5.48 (0.91)
Q^2(12)	6.41 (0.85)	11.30 (0.41)	5.20 (0.92)

larger than that on industrials: a 1% increase in inflation in month $t-2$ raises industrials' returns by 0.99 and financials' returns by 1.68%, respectively.

For the null hypothesis that the sum of the estimated inflation coefficients is zero, the estimated *F*-statistics reject the hypothesis only for the financial sector, which indicates that this sector serves as the best hedging tool against anticipated inflation. The sum of the inflation coefficients in the financials equation is 1.46 and significant at the 5% level, suggesting that a 1% increase in anticipated inflation is associated with a 1.46% increase in returns, *ceteris paribus*. These results suggest the Fisher effect holds mainly for the financial sector in Turkey.

In Table 4, the results for the conditional variance show that estimated EGARCH (1, 1) model performs well, indicating significant asymmetric effects. The volume variable enters significantly only in the case of composite returns and with the expected sign. The reported Q and Q^2 test statistics indicate that the estimated models perform satisfactorily, in that there is no evidence of significant serial correlation or ARCH effects, respectively.

Table 5. **Impact of anticipated inflation on nominal interest rates: TARCH(1, 1) model estimates**
The sample period runs from January 1987 to December 2000. The dependent variable is the change in the level of interest rate. $\Sigma\Pi$ is the sum of the estimated inflation coefficients. P-values are in parentheses. The P-value for $\Sigma\Pi$ is based on the null hypothesis that the sum of the coefficients is zero.

$$R_t = \alpha_0 + \alpha_1 \Pi_{t-1} + \alpha_2 \Pi_{t-2} + \alpha_3 \Pi_{t-3} + \varepsilon_t,$$
$$\varepsilon_t \sim N(0, h_t),$$
$$h_t = \beta_0 + \beta_1 h_{t-1} + \beta_2(\varepsilon_{t-1}^2 d_{t-1}) + \beta_3 \varepsilon_{t-1}^2,$$

	Estimated Coefficient	P-value
α_0	− 0.13	0.81
α_1	0.06	0.66
α_2	− 0.13	0.49
α_3	0.22	0.25
$\Sigma\Pi$	0.15	0.28
β_0	0.24	0.29
β_1	0.51	0.00***
β_2	− 2.20	0.00***
β_3	2.33	0.00***
Log-likelihood	− 593.5	
Q (12)	13.20	0.35
Q^2(12)	14.03	0.26

Table 5 reports the results for the bond market.[6] None of the estimated inflation coefficients is significant at any reasonable level of significance. Nor is the F-statistics for the sum of the estimated inflation coefficients significantly different from zero, which demonstrates that the bond market does not act well as a hedge against anticipated inflation in Turkey. The reported Q and Q^2 test statistics indicate that neither significant serial correlation nor additional ARCH effects are respectively present in the data.

6. Robustness analysis

The results in Tables 4 and 5 might be distorted because we exclude other key macroeconomic determinants of financial market returns from the estimations. Therefore, as a robustness check, we include exports and international reserves announcements in the estimations. The export sector plays a key role in economic growth in Turkey. Movements in international reserves, which include remittances from Turkish workers in Germany, capture the impact of capital flows on asset markets and might also reflect monetary policy changes, such as open market operations and sterilization activities.

Tables 6 and 7 report the results for the stock and bond markets, respectively. To account

6 To control for global influences, we have also used a U.S. Treasury index. However, this variable was insignificant. Therefore, we excluded it from final estimations.

Table 6. **Reaction of nominal stock returns to anticipated inflation, exports and international reserves** R_t^i refers to individual sector and composite index returns. Global market developments are captured by the U.S. S&P 500 index returns. The sample period for the composite returns runs from December 1986 to March 2001, while the industrials' and financials' data begin in January 1991 and end in March 2001. $\Sigma\Pi$, ΣX, and ΣR are the sum of the estimated coefficients for inflation, the growth of exports and the growth of international reserves, respectively. Volume data are measured by the growth rate of the number of shares. P-values are in parentheses. The P-value for $\Sigma\Pi$, ΣX, and ΣR refers to the null hypothesis that the corresponding sum of the coefficients is zero guarantee.

$$R_t^i = \alpha_0 + \alpha_1 R_t^{US} + \alpha_2\Pi_{t-1} + \alpha_3\Pi_{t-2} + \alpha_4\Pi_{t-3} + \alpha_5 X_{t-1} + \alpha_6 X_{t-2} + \alpha_7 X_{t-1},$$
$$+ \alpha_8 R_{t-1} + \alpha_9 R_{t-2} + \alpha_{10} R_{t-3} + \varepsilon_t,$$
$$\varepsilon_t \sim N(0, h_t),$$
$$\log h_t = \beta_0 + \beta_1 \log h_{t-1} + \beta_2\left|\frac{\varepsilon_{t-1}}{h_{t-1}}\right| + \beta_3\frac{\varepsilon_{t-1}}{h_{t-1}} + \beta_4(\text{volume})_t,$$

	Composite	Industrials	Financials
α_0	− 0.33 (0.99)	− 4.11 (0.32)	− 9.27 (0.00)***
α_1	0.35 (0.03)**	0.00 (0.99)	0.21 (0.10)*
α_2	− 0.94 (0.04)**	− 0.81 (0.10)*	0.61 (0.01)***
α_3	0.86 (0.04)**	0.96 (0.03)**	1.04 (0.01)***
α_4	0.29 (0.46)	0.95 (0.04)**	0.43 (0.21)
$\Sigma\Pi$	0.23 (0.76)	1.10 (0.18)	2.08 (0.00)***
α_5	− 0.03 (0.73)	0.02 (0.90)	0.17 (0.10)*
α_6	0.05 (0.66)	0.13 (0.43)	0.33 (0.00)***
α_7	− 0.04 (0.66)	− 0.04 (0.71)	0.01 (0.88)
ΣX	0.02 (0.94)	0.11 (0.78)	0.49 (0.05)**
α_8	0.36 (0.00)***	0.40 (0.00)***	0.44 (0.00)***
α_9	− 0.07 (0.47)	0.13 (0.28)	0.04 (0.71)
α_{10}	− 0.01 (0.91)	− 0.00 (0.98)	− 0.08 (0.41)
ΣR	0.28 (0.10)*	0.53 (0.00)***	0.40 (0.00)***
β_0	3.56 (0.00)***	2.77 (0.00)***	0.99 (0.02)**
β_1	0.21 (0.10)*	0.36 (0.01)***	0.58 (0.00)***
β_2	0.40 (0.00)***	0.36 (0.10)*	1.07 (0.00)***
β_3	0.51 (0.00)***	0.50 (0.00)***	0.42 (0.00)***
β_4	0.01 (0.00)***	0.01 (0.00)***	0.02 (0.00)***
Log-likelihood	− 689.70	− 479.9	− 487.8
Q (12)	9.38 (0.59)	7.62 (0.75)	9.93 (0.62)
Q^2(12)	14.48 (0.27)	13.54 (0.26)	8.91 (0.71)

***, **, and * denote significance level at 1%, 5%, and 10%, respectively.

for the impact of anticipated exports and international reserves, we use three lagged values of these variables. The stock market results in table 6 indicate that adding exports and international reserves in estimations shows the significant effects of the U.S. market on the composite and financials returns. A 10% increase in the U.S. stock returns indicates a 3.5% and 2.1% increase in the composite and financials returns, respectively.

When we look at the impact of inflation on the composite index returns, we see that inflation coefficient in month $t-1$ is now negative and significant. This ''puzzling sign'' is also reported in earlier findings (i.e., Gültekin, 1983). The inflation coefficient in month $t-3$ becomes positive and significant. However, the sum of the estimated coefficients is not significantly different from zero, which suggests that the composite index is not a good hedge against anticipated inflation in Turkey.

In periods $t-2$ and $t-3$, the estimated inflation coefficients for the industrials are positive and significant, but they are negative and significant in month $t-1$. The reported F-statistic indicates that the sum of estimated inflation coefficients is not significantly different from zero, indicating a lack of evidence for the Fisher effect for this sector as well.

In the results for the financials, anticipated inflation continues to have the most significant impact. All the estimated inflation coefficients are positive, and individually and jointly significant. The sum of the coefficients is 2.08: a 1% increase in the expected inflation rate raises the financial returns by 2.08%, all else constant. These results indicate that the financials sector serves as the best hedge against anticipated inflation in Turkey, and the Fisher effect largely holds for this sector. These results are also consistent with those reported in Table 4.

When we examine the significance of other macroeconomic variables, we see that exports announcements are significant only for the financial sector. The sum of the estimated exports coefficients for the financials is 0.33 and significant at the 1% level, which suggests that, as an important source of liquidity in the financial sector, exports revenues have a significant impact on the market activity.

The results in table 6 show that the effect of international reserves in month $t-1$ is significant and positive for all stock returns, and that the sum of the estimated coefficients for all equations is positive and significantly different from zero. This finding indicates that changes in international reserves have a significant impact on common stocks. Its biggest impact is on the industrial sector: a 10% increase in international reserves raises this sector's returns by 5.3%, *ceteris paribus*.

The results for the conditional variance equation in Table 6 indicate that the ARCH effects reported in Table 4 are not qualitatively affected when we include exports and international reserves in the estimations. However, the estimated log-likelihood values are higher, suggesting that including these variables improves the performance of the empirical model.

Table 7 reports the results for the bond market. Although the estimated inflation coefficients cumulatively continue to be statistically insignificant, an inflation announcement in month $t-1$ has a significant and positive impact on nominal interest rates. The estimated coefficient, 0.46, suggests that a 1% increase in the expected inflation rate raises nominal interest rates by about 0.5%. This evidence is consistent with that reported in Berument and Malatyali (2001). In their study, as a proxy for the expected inflation rate they use last month's inflation. They report a coefficient of 0.57, which is close to our estimate of 0.46. Overall, both their results and ours suggest that the nominal interest rate increases less than inflation, suggesting that the real interest rate decreases with higher inflation.

When we look at the rest of the variables, only exports appear to have a significant effect on interest rates. However, this impact is marginal. Exports in month $t-3$ have a negative impact on nominal interest rates, but the economic significance is relatively limited: a 10% increase in exports tends to lower the interest rate by 1%, holding everything else constant.

Table 7. **Impact of anticipated inflation, exports and international reserves on nominal Interest rates** The sample period runs from January 1987 to December 2000. P-values are in parentheses. $\Sigma\Pi$, ΣX, and ΣR are the sum of the estimated coefficients for inflation, the growth of exports and growth of international reserves, respectively.

$$
\begin{aligned}
R_t^i = {} & \alpha_0 + \alpha_1\Pi_{t-1} + \alpha_2\Pi_{t-2} + \alpha_3\Pi_{t-4} + \alpha_4 X_{t-1} + \alpha_5 X_{t-2} + \alpha_6 X_{t-3} + \alpha_7 R_{t-1} \\
& + \alpha_8 R_{t-2} + \alpha_9 R_{t-3} + \varepsilon_t, \\
\varepsilon_t & \sim N(0, h_t), \\
h_t & = \beta_0 + \beta_1 h_{t-1} + \beta_2(\varepsilon_{t-1}^2 d_{t-1}) + \beta_3\varepsilon_{t-1}^2,
\end{aligned}
$$

	Estimated Coefficient	P – value
α_0	0.65	0.52
α_1	0.46	0.03**
α_2	– 0.30	0.15
α_3	0.05	0.78
$\Sigma\Pi$	0.21	0.33
α_4	– 0.01	0.88
α_5	0.02	0.95
α_6	– 0.03	0.65
ΣX	– 0.02	0.34
α_7	0.01	0.82
α_8	– 0.03	0.57
α_9	– 0.08	0.10*
ΣR	– 0.10	0.46
β_0	2.66	0.13
β_1	0.59	0.004**
β_2	– 1.25	0.05**
β_3	1.40	0.03**
Log-likelihood	– 595.9	
Q (12)	12.71	0.31
$Q^2(12)$	16.76	0.16

***, **, and * denote significance level at 1%, 5%, and 10%, respectively.

The negative relation between exports and interest rates is as we expected. Higher export revenues, which raise the level of international reserves, increase the liquidity of the market, causing a drop in interest rates due to the liquidity effect.

7. Implications and conclusions

To the best of our knowledge, this paper represents the first study to investigate the validity of the Fisher effect and the significance of public information arrival for stock price determination in the Istanbul Stock Exchange. Our study complements an earlier study by Berument and Malatyali (2001), who initially provided evidence from the bond market in Turkey. The results in both studies suggest that inflation increases more than do nominal asset returns, implying that real returns to investors decline with inflation.

Except for the financials sector, we find no strong evidence in favor of the Fisher effect in Turkish financial markets. This finding should be interpreted cautiously, however, because of the joint hypotheses problem associated with the Fisher hypothesis tests. In other words, our results may be driven by a violation of the market efficiency assumption.

We show that the financials sector reacts more strongly to the release of the CPI than do any other sectors, and that this sector serves best as a hedging tool against anticipated inflation in Turkey. It appears that the relation between nominal asset returns and inflation in high-inflation countries is not as puzzling as are the findings in the industrial countries like the United States.

We also document evidence on the importance of public arrival information in Turkish asset markets. We note that there is an ongoing debate in the literature about the significance of public information relative to that of private information. Our results indicate that public information plays a large role in Turkish financial markets, especially in the stock market.

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